Great Expectations: Past Wages and Unemployment Durations *

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Abstract

We examine the influence of wage expectations on unemployment durations for workers after exogenous lay-offs. To measure wage expectations, we use a wage decomposition that distinguishes between workers' and firms' wage components. Assuming that workers can only observe their own wages but do not know the overall distribution of firm rents, high firm rents may lead to distortions in workers' wage expectations. In this case, workers mistake high rents for a compensation of high productivity. We estimate hazard rate models for unemployment durations in Austria and find that in particular younger workers are longer unemployed if they previously worked in firms which paid high firm-wage components.

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1 Introduction

Job search theory offers a framework to explain the duration of individual unemployment spells. In this framework, unemployed workers search sequentially for the best job offer. If offers arrive at random and the distribution of offers is known to the worker, it is optimal for the searcher to accept the first offer which is at or above the reservation wage. This strategy balances search costs and possible increases in income by further search.

Knowing or learning about the distribution of wage offers is a non-trivial task for job-seekers. While the job searcher is learning over time — updating the prior beliefs by using currently sampled job offers — the choice of an initial prior is important.[1] One way for a job searcher to form a prior for the wage offer distribution is to use his or her past wage. If the past wage correctly corresponds to the worker's outside opportunities, it will be a perfect starting point. If, on the other hand, the worker was overpaid in the last job, e.g., because of seniority wages, this may result in an incorrectly high reservation wage due to a distorted perception of the wage offer distribution. In consequence, the overly high reservation wage will result in the rejection of wage offers the worker would have accepted if the reservation wage had been formed based on the correct wage distribution. Empirically, this will translate into relatively longer unemployment durations, the durations of

[1] In special situations, e.g., if the searcher's prior beliefs follow a Dirichlet distribution and the searcher is updating her priors according to Bayes' rule, this does not matter: even if the wage offer distribution is unknown, the qualitative properties of optimal search strategies remain the same (Rothschild, 1974). But as Rothschild (1974) points out, "(the results) are still quite special, as the proofs depend on the process of revising beliefs to accommodate new information having a particular—and not terribly natural—local property" (p694).

which being determined by how quickly the searcher updates his or her prior of the wage offer distribution.[2]

We study workers who exogenously lost their jobs due to plant closures and analyze their unemployment durations. Because the past wage as such is not sufficiently informative about future earnings potentials, we decompose the past wage into worker-specific, human-capital specific and firm-specific components. This decomposition allows the estimation of firm-specific rents, which reflect deviations from an industry average, i.e., we obtain an indicator of whether the past wage differed from the market-wide wage distribution or not. Our data cover all Austrian workers for more than three decades, which allows us to reliably decompose the last wage before the plant closures and to study the unemployed workers' subsequent labor market spells.

These considerations are also related to recent discussions in behavioral economics about overconfidence (Della Vigna, 2007). Realistic workers will base their expectations only on their true productivity, i.e., the worker-specific component, while overconfident workers might mistake (parts of) the firm rent for their own productivity. Overconfidence on the part of job searchers might therefore prolong unemployment durations. While there is considerable field evidence on overconfidence in e.g., trading patterns of individuals (Barber and Odean, 2001) or in CEO behavior (Malmendier and Tate, 2005), there is no direct evidence on labor market or search behavior.[3]

[2]Winter-Ebmer (1998) studies the relation between the wage distribution in the last firm and unemployment duration and finds that average wages have no impact whereas some parameters of the wage distribution do.

[3]Dubra (2004) assumes that searchers are overconfident and explores search behavior and welfare effects whereas there is a larger experimental literature on bargaining behavior, e.g., Babcock and Loewenstein (1997).

Our analysis is also relevant for the discussion about employment patterns of elderly workers. It is often argued that Continental European labor markets are rigid, especially because of age- and tenure-related wage schedules, and in addition to earnings-related (Bismarckian) welfare state benefits, elderly workers might easily become too expensive, given their productivity (Saint-Paul, 2009). If elderly workers become unemployed, past wages that are in excess of their productivity due to seniority-based wages might lead to excessive reservation wages and long unemployment durations. We provide empirical evidence for this argument by examining the association of firm-specific rents with the unemployment durations of older workers.

2 Empirical Strategy

We analyze the relationship between wages and subsequent unemployment durations using proportional discrete time hazard rate models. We use the Prentice and Gloeckler (1978) model and a model augmented with a discrete mixture distribution to summarize unobserved individual heterogeneity, as proposed by Heckman and Singer (1984), using Jenkins' (2004) Stata module.

Suppose there are $i = 1, ..., N$ workers who become unemployed at time $t = 0$ and are observed for s time periods. At each point in time, the worker either remains unemployed or finds new employment. The discrete hazard rate in period t is (Prentice and Gloeckler, 1978):

$$h_t = 1 - \exp(-\exp(\beta_0 + X_{it}\beta)), \tag{1}$$

where β_0 is an intercept and the linear index function, $X_{it}\beta$, incorporates the impact of the covariates. (See also Jenkins (1995).) Workers who leave the sample for other reasons, e.g., moving abroad, are treated as censored.

Suppose that each worker belongs to a group of an unobserved type, e.g., low or high ability in obtaining a job. This can be parameterized by allowing the intercept term β_0 to differ across types (Heckman and Singer, 1984). In a model with types $z = 1, ..., Z$, the hazard function for worker belonging to type z is:

$$h_{z,t} = 1 - \exp(-\exp(m_z + \beta_0 + X_{it}\beta)), \qquad (2)$$

and the probability of belonging to type z is p_z. The m_z are the mass points of a multinomial distribution where m_1 is normalized to equal zero and $p_1 = 1 - \sum_{z=2}^{z=Z} p_z$. The z-th mass point equals $m_z + \beta_0$.

This econometric specifications allows to control for time-varying covariates and to investigate the importance of unobserved heterogeneity for leaving unemployment. The vector of characteristics, X_{it}, includes time-invariant characteristics, e.g., the firm size at the time of unemployment, and time-varying characteristics, such as e.g., the replacement rate of the unemployed. In addition to these (standard) controls, we also control for whether the worker enjoyed above-average firm rents or not, estimated from a decomposition of the wages. We expect that workers who had received above-average firm rents to remain unemployed longer, all other things equal.

3 Data

We use linked employer-employee data from the Austrian Social Security Database (ASSD) which contains detailed information on all workers covered by the Austrian social security system from 1972 to 2009.[4] We restrict our sample to workers who were laid off due to plant closure between 1990 and 1996 and are between 20 and 55 (50 for females) years of age at time of dismissal. We exclude the construction and the tourism sector because of strong seasonality in employment in these sectors. We also limit our sample to workers with a minimum tenure of six weeks in the firm. Typically, a sample of job searchers is composed of workers who were fired in their old job due to inadequate performance, workers who were fired due to labor demand volatility and workers who quit voluntarily. Both workers fired for cause and those quitting voluntarily pose a problem for an analysis of wage expectations, because the separation is an endogenous event. We therefore concentrate on workers from plant closures, where the cause of unemployment is an exogenous event.

Plant closures are not directly observed in the data set, but identified indirectly by the disappearance of the identifier from the data. To ensure that so identified plant closures are truly plant closures, we define firms as closing firms only if one of the following requirements is fulfilled. First, either the majority of workers are subsequently not employed, (2) the majority of workers are associated with a new firm identifier but account for less than 50% of the new firm's workforce or (3) the majority of workers is employed

[4]See Zweimüller et al. (2009) for a description of the data.

in different firms.

In total, we observe 37,432 male and 28,078 female workers in 31,704 firms in closing plants. From these, we exclude workers for whom we cannot decompose the wages.[5] Unemployment duration is counted as the number of days starting from the day the worker is laid off until the worker starts a new job. Unemployment spells that last longer than 1,500 days are coded as censored. Spells that end with death, self employment, maternity leave, subsidized employment or sick leave lasting for more than 6 months are also coded as censored.

Table 1 shows the composition of our sample in detail. More than 50% of all workers in the sample start a new job immediately after the plant closure. Of the remaining 16,574 male and 10,448 female workers who became unemployed for at least one day, approximately 90% (85%) find new employment within 1,500 days.

3.1 Decomposition of wages

We construct an indicator of whether the past wage differed from the market-wide wage distribution or not by decomposing the wages into worker-specific, human-capital specific and firm-specific components. For this wage decomposition, we use the universe of all blue-collar workers, 1980 to 2000, about 3.8 million workers in about 460,000 firms. Wages are decomposed following

[5] Notice that firm fixed components in the wage decompositions are only identified if we observe at least one worker moving in or out of a firm within a pre-defined time interval. Accordingly, worker fixed components are only observed for those workers who are observed in at least two different firms.

Abowd et al. (1999):

$$\underbrace{y_{it}}_{\log(\text{wage})} = \underbrace{\phi_{J(it)}}_{\text{firm-fixed component}} + \underbrace{\theta_i}_{\text{person-fixed component}} \tag{3}$$
$$+ \underbrace{X'_{it}\beta}_{\text{returns to productivity}} + \epsilon_{it},$$

where

$$E\left[\epsilon_{it}|i, t, J(i,t), x_{it}\right] = 0. \tag{4}$$

The parameter ϕ_j in equation (3) gives the difference in earnings in firm $j = 1, ..., J$, relative to the average firm. This is our indicator of the distortion of wage expectations as it indicates a relatively low or high wage in the past job. The parameter θ_i captures all (unobserved) constant differences between workers and may be seen as a proxy for ability. The parameter β captures economy wide returns to productivity and experience for the time-varying characteristics of worker i, x_{it}. Using this interpretation of the parameters, a highly productive worker is one with θ_i greater than the average and a firm that pays high wages is one with ϕ_j greater than the average (Abowd et al., 2004).[6]

We are able to identify worker and firm fixed components for 3,818,508 workers and 459,144 firms. Firm and worker fixed components show weak negative correlation for male (-0.01) and female workers (-0.006).[7]

[6] We use Ouazad's (2008) Stata module. Standard errors are obtained via bootstrapping.

[7] Detailed estimation results are shown in the appendix.

It is important to stress that the assumption behind equation (4) requires the error term to be independent of any observable effects in x_{it}, the person-fixed component θ_i or the firm-fixed component ϕ_j. In other words, it assumes exogenous mobility. If there is positive assortative matching, i.e., good firms employ good workers, then the correlation between θ_i and ϕ_j should be positive (and large).[8] Following De Melo (2008), we use the correlation between a worker's fixed component, θ_i and the mean of the co-workers' fixed components, $\bar{\theta}_{-i}$, to detect sorting. This correlation is small, $corr(\theta_i, \bar{\theta}_{-i}) = 0.095$, and indicates that there is little sorting in our data.[9]

A comparison of the average wages before and after plant closure shows that workers with high firm wage components (fwc) experience a higher wage loss than those with low fwc. (See Figure 2.) Figure 3 indicates that this loss in wages is accompanied by a loss in fwc for those coming from high fwc firms. Workers who had low fwc experienced on average a gain in their fwc. The relative wage loss for previously high fwc earners is at least partly due to significant changes in fwc. Workers who were previously low fwc earners experience little change in their fwc.

Figure 1 in the Appendix shows the resulting distribution of firm rents in our sample.

[8] Indeed, similar to Abowd, Kramarz, Lengermann and Pérez-Duarte (2004) we find that the components are (weakly) negatively correlated. Shimer (2005) shows that a model with coordination frictions may lead to positive but imperfect correlation between workers' productivity and firms' types. Abowd et al. (2004), in a simulation of Shimer's (2005) results, obtain a negative correlation between person and firm components. While Abowd et al. (2004) caution that the mere examination of the correlation between the person and the firm components is not sufficient to provide evidence for or against sorting in the labor market, these parameters are first-order approximations of the true model.

[9] We bootstrap the correlation using 50 replications.

Table 2 provides average unemployment durations by gender and by wage components. The average unemployment duration was 121 days for men and 127 days for women. We see that on average men who had a low firm component remain unemployed for fewer days, 115 days, than men who had a high wage component, who remain unemployed for 127 days. For women, the association is reversed, women who had a low firm component remain unemployed for 128 days and those who had a high firm component remain unemployed for 120 days on average.

Workers who have a low person-fixed component remain on average unemployed for much longer than those who have a high person-fixed component, 135 vs. 107 days for men and 135 vs. 112 days for women. We also see that the average unemployment duration increases with age, young male workers remain unemployed for some 113 days on average and older male workers about 149 days. Female workers show a similar pattern.

4 Results

In Table 4 we present results from a non-parametric discrete-time hazard rate models, estimated separately for men and women. Our preferred specifications are in columns (2) and (5), which include our parameters from the wage decomposition and including mass points for unobserved heterogeneity. Columns (1) and (4) exclude the wage components, whereas columns (3) and (6) disregard heterogeneity. Person-specific and firm-specific components are introduced as dummy variables which indicate below or above average values.

We find that individuals who have a high person-specific component leave unemployment earlier, in particular men. This variable is a proxy for fixed personal traits, such as ability, productivity or work effort, and the positive effect on unemployment durations is therefore expected. This variable captures different aspects of worker heterogeneity to the mass points, because the inclusion of mass points in the econometric specification does not change these coefficients significantly.

Workers who enjoyed above average firm-specific components in their past wages, have, in fact, lower hazard rates. This is estimated for both men and women. It seems that these workers do not fully distinguish between their productivity and the firm-specific component of their previous wages. The resulting longer unemployment durations are compatible with the hypothesis that these workers base their expectations about wage offers not only on their person-specific component, but also on the rent they enjoyed in the past firm. These workers could be characterized as being overconfident of their own abilities and productivity. In other words, they attribute the wage they earned in the past firm fully towards their own capabilities and disregard the randomness which might have played a role in the rent they enjoyed in the last firm. Interestingly, we do not find differences in overconfidence for the genders, which contrasts results from investing behavior in finance (Barber and Odean, 2001).

The additional variables, such as the replacement rate or age, which are included as controls, yield stable estimates across specifications. A higher replacement rate reduces the hazard rate and older workers search somewhat longer than younger workers. Workers who had worked in larger firms have a higher hazard rate than those who had worked in smaller firms, which could be explained by the larger network of contacts in larger firms.

In Table 5, we report results for separate age groups, young workers who are between 20 and 30 years of age, prime-age workers who are between 30 and 45 years of age and older workers who are older than 45 years of age. Interestingly, while there are little differences in the estimated coefficients to the results above, the only distortions in search behavior we find are for young workers. For both men and women, young workers who had a high firm component have significantly longer unemployment duration than those who had a low firm component. This result might be explained by younger workers having less experience in assessing their productivity in the labor market. It might also be explained by a higher prevalence of overconfidence among the younger than among the older workers. We do see somewhat higher overconfidence for men than for women, but the difference is not statistically significant.

More importantly, we do not see any evidence for distorted wage expectations for prime-age workers and, in particular, for older workers. It seems that these workers do not have excessive wage expectations stemming from rents in previous firms; Saint-Paul's (2009) argument for the unemployability of elder workers due to immoderate reservation wages is not supported by our evidence.

Distorted expectations about productivity and overconfidence might be related to job tenure. Workers who are new in a firm might have a better understanding of the outside opportunities, which they faced when they searched for the current job, and might therefore have less distorted views about the wage offer distribution and their own productivity. On the other hand, workers who are new in a firm may need time to understand and assess their own contribution to the firm and to distinguish the returns to their own productivity from firm rents. To investigate this issue, we separate our sample into workers with short and long tenure in the previous firm.[10] A tenure is short if it lasted up to 500 days. The estimation results are tabulated in Tables 6 and 7.

We see that young men are always overconfident, i.e., we estimate a negative coefficient for the high wage component, regardless if they were displaced from a short- or long-tenured job. Women, in contrast, only exhibit overconfidence if their previous job duration was short. The coefficient for the firm-specific component is zero for young women who came from long-tenured jobs. As was to be expected from Table 5, there is no effect for prime-age and older workers.

5 Conclusion

According to psychological research workers might attribute (excessively) high wages to their own abilities rather than to pure luck in obtaining employment with a firm that pays high rents. Such a distorted assessment could result in a systematic misjudgement of the wage offer distribution a job searcher faces. We study job search behavior of workers who were made

[10]Due to small sample sizes, we pool prime age and older workers.

redundant due to plant closures in Austria and find that, in particular, young workers are overconfident: high firm rents in the past wages lead to significantly longer unemployment durations.

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A Tables and Graphs

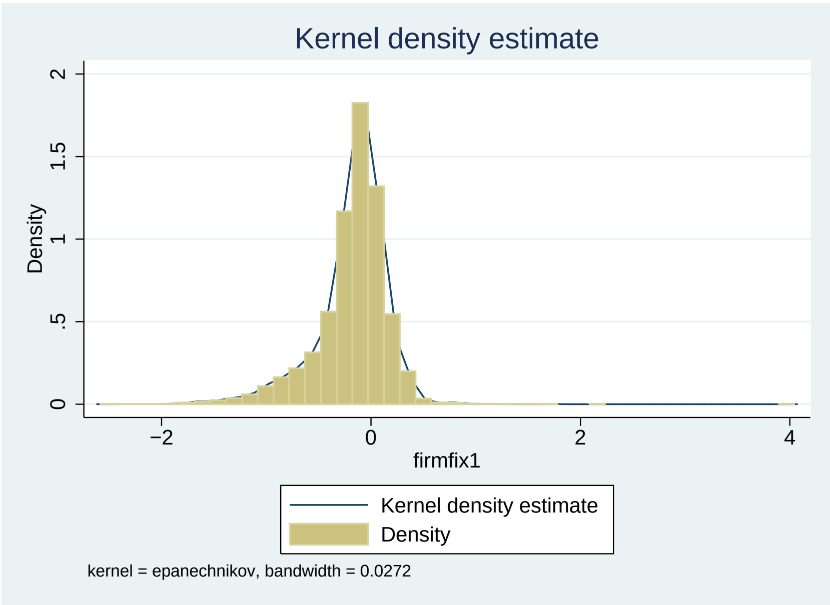

Figure 1: Distribution of Firm Wage Effects in ASSD

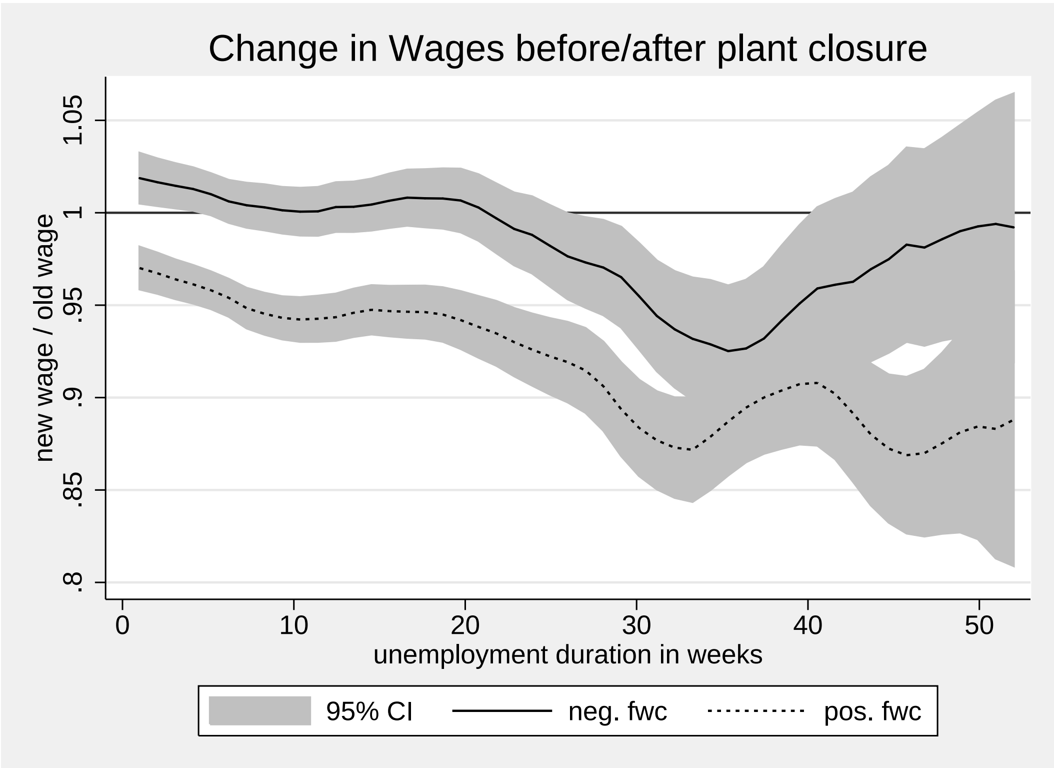

Figure 2: Change in wages before and after plant closure by level of firm wage components (fwc)

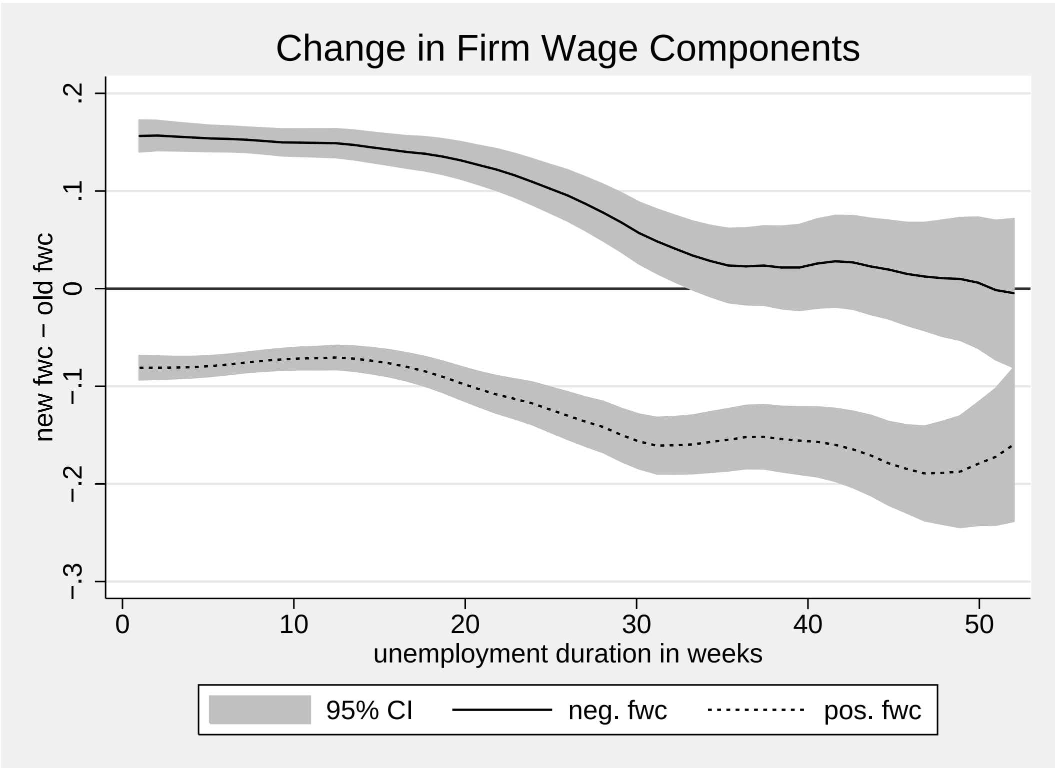

Figure 3: Change in firm wage components before and after plant closure by level pre-displacement fwc

Table 1: Transitions after Plant Closure.

	male	female
status after plant closure:		
job to job transition	18,079 (52.2%)	13,899 (57.1%)
unemployed after plant closure	16,574 (47.8%)	10,448 (42.9%)
transition after unemployment:		
reemployed	14,998 (90%)	8,862 (85%)
retired after unemployment	185 (1%)	69 (1%)
censored	1,401(9%)	1,517 (14%)

Note: 34,653 male and 24,347 female blue-collar workers.

Table 2: Average unemployment durations (days), by gender and wage components.

	male		female	
	mean	N	mean	N
all	121	16574	127	10448
low firm component	115	8913	128	9147
high firm component	127	7661	120	1301
low person component	135	8156	135	6914
high person component	107	8418	112	3534
young (20-30)	113	6785	118	4460
prime age (30-45)	117	6994	131	4609
old (45+)	149	2795	143	1379

Table 3: Estimation results from wage decomposition.

	Coef.	Std. Err.[1]	z	P> $\|z\|$
ltenure	.0145164	.0005595	25.94	0.000
experience 1-3[2]	.0194157	.0014297	13.58	0.000
experience 4-5	.0318344	.0023718	13.42	0.000
experience 6-8	.1133133	.0023854	47.50	0.000
experience 9-12	.1318127	.0023171	56.89	0.000
experience 13-17	.1474905	.0029602	49.82	0.000
experience 17+	.1661588	.0042473	39.12	0.000
age	.0230653	.000056	411.60	0.000
age2	-.0219547	.0000756	-290.25	0.000
firmsize	-1.54e-06	1.05e-07	-14.72	0.000

*Note:*Additional explanatory variables: year, region, industry dummies
[1]: Standard errors obtained via bootstrapping (50 repetitions)
[2]: baseline: 0-1 years of experience

Table 4: Estimated hazard rates from unemployment to employment, by gender.

	Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)
high firm component (0/1)	-	-0.054**	-0.057***	-	-0.055*	-0.024
	-	(0.027)	(0.022)	-	(0.030)	(0.025)
high person component (0/1)	-	0.217***	0.184***	-	0.098***	0.099***
	-	(0.022)	(0.017)	-	(0.029)	(0.024)
replacement rate	-0.032***	-0.032***	-0.024***	-0.031***	-0.031***	-0.025***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
age	-0.015***	-0.013***	-0.013***	-0.008***	-0.007***	-0.005***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)
log(firmsize)	0.135***	0.128***	0.115***	0.190***	0.193***	0.148***
	(0.008)	(0.008)	(0.006)	(0.012)	(0.012)	(0.009)
masspoint	1.334***	1.329***	-	1.244***	1.247***	-
	(0.028)	(0.028)	-	(0.049)	(0.049)	-
P(masspoint)	0.647	0.659	-	0.724	0.729	-
	0.0181	0.0182	-	0.0487	0.0494	-
Obs.	16574	16574	16574	10448	10448	10448
logl	-34591	-34539	-35040	-22011	-22003	-22207

Note: Discrete-time proportional hazard rate models. Additional variables are 5 year, 8 region (*Bundesland*) and 15 industry dummy variables (*onace_17*). See Zweimüller et al. 2009).

Table 5: Estimated hazard rates from unemployment to employment, by gender and age group.

	male workers			female workers		
	20 – 30	30 – 45	45+	20 – 30	30 – 45	45+
	(1)	(2)	(3)	(4)	(5)	(6)
high firm component (0/1)	-0.145***	0.016	0.007	-0.103**	-0.028	0.063
	(0.042)	(0.041)	(0.074)	(0.048)	(0.045)	(0.083)
high person component (0/1)	0.202***	0.257***	0.263***	0.152***	0.028	0.202**
	(0.033)	(0.033)	(0.059)	(0.042)	(0.045)	(0.087)
replacement rate	-0.032***	-0.032***	-0.033***	-0.033***	-0.029***	-0.030***
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
age	0.004	-0.001	-0.107***	-0.034***	0.005	-0.080***
	(0.006)	(0.004)	(0.007)	(0.007)	(0.004)	(0.020)
log(firmsize)	0.105***	0.112***	0.183***	0.142***	0.214***	0.253***
	(0.012)	(0.013)	(0.022)	(0.018)	(0.018)	(0.034)
masspoint	1.233***	1.287***	1.471***	1.375***	1.077***	1.010***
	(0.044)	(0.045)	(0.078)	(0.066)	(0.084)	(0.145)
P(masspoint)	0.669	0.697	0.538	0.714	0.738	0.609
	0.0301	0.0294	0.0541	0.0347	0.0576	0.189
Obs.	6785	6994	2795	4460	4609	1379
logl	-14196	-14601	-5546	-9088	-9957	-2883

Note: Discrete-time proportional hazard rate models corresponding to columns (2) and (5) in Table 4. Additional variables as in Table 4.

Table 6: Estimated hazard rates from unemployment to employment for male workers, by pre-displacement tenure.

	(1)	(2)	(3)	(4)	(5)	(6)
	all		young		prime age / old	
	short tenure	long tenure	short tenure	long tenure	short tenure	long tenure
high firm component (0/1)	-0.076**	-0.025	-0.146***	-0.167**	-0.018	0.038
	(0.035)	(0.043)	(0.051)	(0.075)	(0.049)	(0.053)
high person component(0/1)	0.163***	0.266***	0.166***	0.235***	0.178***	0.292***
	(0.028)	(0.034)	(0.040)	(0.058)	(0.040)	(0.042)
replacement rate	-0.030***	-0.035***	-0.032***	-0.033***	-0.028***	-0.036***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)
age	-0.009***	-0.017***	-0.003	0.016	-0.017***	-0.035***
	(0.002)	(0.002)	(0.007)	(0.010)	(0.003)	(0.003)
log(firmsize)	0.113***	0.152***	0.085***	0.157***	0.135***	0.150***
	(0.010)	(0.013)	(0.015)	(0.023)	(0.015)	(0.017)
masspoint	1.127***	1.575***	1.100***	1.435***	1.120***	1.604***
	(0.039)	(0.043)	(0.055)	(0.080)	(0.059)	(0.052)
P(masspoint)	0.652	0.664	0.640	0.724	0.680	0.631
	0.0321	0.0211	0.0472	0.0388	0.0493	0.0268
Obs.	8621	7953	4134	2651	4487	5302
logl	-18682	-15787	-8910	-5260	-9750	-10462

Note: A tenure is short if it was shorter or equal to 500 days. Discrete-time proportional hazard rate models corresponding to column (2) in Table 4.
Additional variables as in Table 4.

Table 7: Estimated hazard rates from unemployment to employment for female workers, by pre-displacement tenure.

	(1)	(2)	(3)	(4)	(5)	(6)
	all		young		prime age / old	
	short tenure	long tenure	short tenure	long tenure	short tenure	long tenure
high firm component(0/1)	-0.050	-0.051	-0.128**	-0.051	0.054	-0.050
	(0.043)	(0.043)	(0.062)	(0.075)	(0.059)	(0.052)
high person component(0/1)	0.075*	0.138***	0.089*	0.271***	0.050	0.051
	(0.040)	(0.041)	(0.054)	(0.067)	(0.060)	(0.053)
replacement rate	-0.027***	-0.034***	-0.028***	-0.040***	-0.026***	-0.031***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)	(0.001)
age	-0.005**	-0.011***	-0.024**	-0.058***	-0.006	-0.010***
	(0.002)	(0.002)	(0.009)	(0.011)	(0.004)	(0.004)
log(firmsize)	0.146***	0.243***	0.095***	0.218***	0.208***	0.242***
	(0.016)	(0.017)	(0.023)	(0.030)	(0.023)	(0.021)
masspoint	1.045***	1.499***	1.167***	1.798***	0.842***	1.283***
	(0.060)	(0.094)	(0.081)	(0.158)	(0.099)	(0.113)
P(masspoint)	0.628	0.815	0.645	0.830	0.611	0.791
	0.0554	0.0297	0.0589	0.0354	0.137	0.0489
Obs.	4941	5507	2588	1872	2353	3635
logl	-10709	-11250	-5477	-3569	-5206	-7642

Note: A tenure is short if it was shorter or equal to 500 days. Discrete-time proportional hazard rate models corresponding to column (5) in Table 4.
Additional variables as in Table 4.