



# Great expectations: Past wages and unemployment durations<sup>☆</sup>

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## ABSTRACT

Decomposing wages into worker and firm wage components, we find that firm-fixed components are sizeable parts of workers' wages. If workers can only imperfectly observe the extent of firm-fixed components in their wages, they might be misled about the overall wage distribution. Such misperceptions may lead to unjustified high reservation wages, resulting in overly long unemployment durations. We examine the influence of previous wages on unemployment durations for workers after exogenous lay-offs and, using Austrian administrative data, we find that younger workers are, in fact, unemployed longer if they profited from high firm-fixed components in the past. We interpret our findings as evidence for overconfidence generated by imperfectly observed productivity.

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## 1. Introduction

Job search theory offers a framework to explain the duration of unemployment spells. In this framework, unemployed workers search sequentially for a job. If job offers arrive at random and the distribution of offers is known, 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 expected gains from 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 prior beliefs with recently sampled job offers – the choice of an initial prior is important.<sup>1</sup> A job searcher may use his or

her past wage as a prior for the wage offer distribution. If the past wage equals the worker's productivity, it will be a perfect starting point. If, however, the wage was greater than the worker's productivity, e.g., because of seniority wages, this may result in an overly high reservation wage due to a distorted perception of the worker's productivity. In consequence, the overly high reservation wage will result in the rejection of wage offers the worker would have accepted had the reservation wage been based on the correct wage distribution. Empirically, this will translate into relatively longer unemployment durations, which are being determined by how quickly the searcher updates his or her prior of the wage offer distribution.<sup>2</sup>

We study workers who exogenously lost their jobs due to plant closures and analyze their unemployment durations. A random sample of unemployed workers would be problematic for two reasons: workers dismissed for a cause might be negatively selected and, more importantly, workers who quit their job voluntarily typically do so because they are looking for a better(-paid) job. These workers will bias the analysis. Using unemployed workers from plant closures solves this problem, because plant closures hit all workers alike.

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<sup>1</sup> 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" (p. 694).

<sup>2</sup> Winter-Ebmer (1998) studies the relation between the wage distribution in the last firm and unemployment durations and finds that average wages have no association with unemployment durations, while other parameters of the wage distribution, e.g., inequality in the old firm, do.

Because wages may contain components which are not related to a worker's productivity, such as rents, seniority pay or efficiency wage components, we decompose past wages into worker-specific, human-capital specific and firm-specific components. The decomposition separates wage components which reflect a worker's productivity, including unobservable productivity components (fixed effects), from a firm-specific part of the wage. The firm-specific component is normalized to reflect deviations from the industry average. 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. Using the same decomposition procedure, Gruetter and Lalive (2009) show that firm fixed effects indeed play an important role in the wage determination in Austria. Their estimates suggest that around 27% of total variation in wages can be explained by unobserved firm effects.

These considerations are related to recent discussions in behavioral economics about overconfidence (Della Vigna, 2007). Workers who judge their productivity correctly will base their expectations of the wage offer distribution only on those parts of the wage which reflect their productivity. Overconfident workers might mistake (parts of) the firm component for their own productivity and attribute the firm component to their own effort and ability. While there is 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 little direct evidence on labor market or search behavior. Hoch (1985) found that MBA students overestimate the number of job offers they will receive and the magnitude of their salary.<sup>3</sup> Explicitly considering the role of behavioral aspects in search processes, Falk et al. (2006) show in a laboratory experiment that people are substantially uncertain about their own abilities and that this misperception of relative abilities affects search outcomes. While some high ability people do not search enough, some low ability people search too much. This idea is related to the notion of loss aversion in finance, where people are reluctant to sell a stock once its price goes below the price the investor has paid before (Kahneman et al., 1991): here workers might be reluctant to sell their human capital for a price below the previously acknowledged value.

Our analysis is also relevant for the discussion of the employment patterns of older workers. For example, Saint-Paul (2009) argues 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, older workers might easily become too expensive, given their productivity. If older workers receive wages that are in excess of their productivity due to seniority-based wages and they mistakenly assume that such wages reflect their true productivity, they will have reservation wages that are too high and end up with long unemployment durations. Our analysis can shed some light on this discussion.

We find that only young workers can be described as overconfident. Workers who previously had a high firm-fixed component tend to search longer for a new job than those who had a low firm component. They presumably expect to find high-paying jobs and turn down more realistic job offers at the start of their unemployment spells. We explore that the pattern might be caused by misconceptions of the true wage distribution, e.g., because workers have not been actively searching for new employment in the past. Our analyses along such lines suggest that overconfidence is the more probable explanation for the relatively longer unemployment durations of workers who had high firm components. In addition, we do not find evidence that older workers remain unemployed because they systematically misjudge potential wages, given their productivity.

<sup>3</sup> Dubra (2004) assumes in a theoretical model that searchers are overconfident and explores search behavior and corresponding welfare effects. There is also a larger experimental literature on bargaining behavior, e.g., Babcock and Loewenstein (1997).

## 2. Empirical strategy

We model unemployment durations with proportional discrete time hazard rate models. The reservation wage,  $\xi$ , is the minimum acceptable wage offer which results from equating expected gains from continued search with the expected value from accepting a wage offer. In a simple search model (Lancaster and Chesher, 1983), it is given by:

$$\xi = b + \frac{\lambda}{\rho} \int_{\xi}^{\infty} (1 - F(\xi)) dw,$$

where  $1 - F(\xi)$  is the probability that the wage offer  $w$  is greater than the reservation wage.  $b$  denotes the value of the outside option and  $\lambda$  is the probability of a wage offer. Future incomes are discounted with a factor  $\rho$ . The hazard rate from unemployment,  $h$ , is given by the product of two probabilities, the probability of receiving an offer,  $\lambda$ , and the probability that the offer is acceptable,  $1 - F(\xi)$ :  $h = \lambda[1 - F(\xi)]$ . Thus the workers' belief of the true wage offer distribution is directly linked to the hazard of leaving unemployment: if workers believe that the wage distribution is to the right of the actual wage distribution, they will reject wage offers as unacceptable and remain unemployed longer. The reverse is true for workers who think that the wage offer distribution is to the left of the actual wage offer distribution, they will leave unemployment earlier. We argue that such distortions may arise from wage components that do not reflect a worker's productivity, but may be caused by firm fixed wage components.

We use the Prentice and Gloeckler (1978) model, augmented with a discrete mixture distribution to account for unobserved individual heterogeneity, as proposed by Heckman and Singer (1984).<sup>4</sup>

Suppose there are  $i = 1, \dots, N$  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):

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

where  $\beta_0$  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., retiring, 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  $\beta_0$  to differ across types (Heckman and Singer, 1984). In a model with types  $z = 1, \dots, Z$ , the hazard function for worker belonging to type  $z$  is:

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

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  $p_1 = 1 - \sum_{z=2}^Z p_z$ . The  $z$ -th mass point equals  $m_z + \beta_0$ .

This econometric specification allows 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 start of the unemployment spell, and time-varying characteristics, such as e.g., the benefit replacement rate of the unemployed. In addition to these (standard) controls, we also control for whether the worker enjoyed above-average firm components or not, estimated from a decomposition of the wages. We expect that workers who had received above-average firm-fixed components will remain unemployed longer, all other things equal. The indicator of the firm-fixed component is by construction orthogonal to the person-specific wage component and

<sup>4</sup> We use Jenkins' (2004) Stata module to estimate the hazard models.

should therefore capture such an effect of a distorted perception of the wage offer distribution on the unemployment duration.<sup>5</sup>

### 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.<sup>6</sup> Because of strong seasonality in employment (Del Bono and Weber, 2008), we exclude construction and tourism workers. We also limit our sample to workers with a minimum tenure of six weeks in the last 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 their separation from the firm is an endogenous event. We therefore concentrate on workers from plant closures where the cause of unemployment is an exogenous event. Our sample consists of workers who were laid off due to plant closures between 1990 and 1996 and who were between 20 and 55 (50 for females) years of age at that time.

Plant closures are not directly observed in the data, but identified indirectly by the disappearance of a firm's identifier. To ensure these disappearances are true plant closures, and not merely caused by e.g., administrative recoding, we define firms as closing firms only if one of the following requirements is fulfilled. First, the majority of workers is not immediately employed after the disappearance of the identifier, (2) the majority of workers is employed in a single firm with a different identifier, but the workers account for less than 50% of the new firm's workforce, or (3) the majority of workers is spread out over different firms.

In total, we observe 28,078 female and 37,432 male workers being laid off from 31,704 closing firms within 60 days before plant closure.<sup>7</sup> From these, we exclude workers for whom we cannot decompose the wages, which reduces our estimating sample to 24,424 female and 34,746 male workers in 30,192 closing firms.<sup>8</sup> The unemployment duration is 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 1500 days are censored. Spells that end with death, self employment, maternity leave, or subsidized employment and sick leave lasting for more than 6 months are also censored.

#### 3.1. Decomposition of wages

Following Gruetter and Lalive (2009) we derive our proxy for the distortion of the reservation wage by decomposing wages into worker-specific, human-capital specific and firm-specific components. For this we use the universe of all blue-collar workers for the years 1980 to 2000 and not only our sample of workers who worked in firms that closed down. These are 3,818,508 workers in 459,144 firms after deleting observations where we cannot identify the wage components.

<sup>5</sup> Hogan (2004) uses an estimate of the wage offer distribution in a duration model to investigate the relationship between the reservation wage and unemployment durations. However, this strategy requires information on expected wages which are not available in our data.

<sup>6</sup> See Zweimüller et al. (2009) for a description of the data.

<sup>7</sup> To check for robustness, we follow Schwerdt (2011) and also sample early leavers: workers who were laid off up to 180 days before plant closure. See below.

<sup>8</sup> Notice that firm-fixed components in the wages are only identified if we observe at least one worker moving in or out of a firm. Similarly, worker-fixed components are only identified for workers who are observed in at least two different firms.

The wages are decomposed following Abowd et al. (1999):

$$\underbrace{y_{ijt}}_{\log(\text{wage})} = \underbrace{\phi_j}_{\text{firm-fixed component}} + \underbrace{\theta_i}_{\text{person-fixed component}} + \underbrace{X'_{ijt}\beta}_{\text{returns to productivity}} + \varepsilon_{ijt}, \quad (3)$$

where

$$E[\varepsilon_{it} | \theta_i, \phi_j, t, X_{ijt}] = 0. \quad (4)$$

The parameter  $\phi_j$  in Eq. (3) gives the difference in earnings in firm  $j = 1, \dots, J$ , relative to the average firm. This is our indicator of firm-fixed components as it indicates a relatively low or high wage in the past job, controlling for observed and unobserved worker heterogeneity. The parameter  $\theta_i$  captures all (unobserved) time-invariant differences between workers and may be seen as a proxy for ability. The parameter vector  $\beta$  captures economy wide returns to productivity and experience for the time-varying characteristics of worker  $i$  in firm  $j$  at time  $t$ ,  $X_{ijt}$ .<sup>9</sup>

It is important to stress that the identifying assumption behind Eq. (4) requires the error term to be independent of any observable effects in  $X_{ijt}$ , the person-fixed component  $\theta_i$  or the firm-fixed component  $\phi_j$ . In other words, it assumes exogenous mobility, which might be a strong assumption. If there is positive assortative matching, i.e., good firms employ good workers, then the correlation between  $\theta_i$  and  $\phi_j$  should be positive (and large).<sup>10</sup> Here, in contrast, we find that firm and worker fixed components are weakly negatively correlated, the correlations are  $-0.01$  for male and  $-0.006$  for female workers. However, Abowd et al. (2004) caution that the mere examination of the correlation between person and firm components is not sufficient to provide evidence for or against sorting in the labor market. We therefore follow De Melo (2008) and additionally calculate the correlation between a worker's fixed component,  $\theta_i$ , and the mean of the co-workers' fixed components,  $\bar{\theta}_{-i}$ . This correlation is small,  $\text{corr}(\theta_i, \bar{\theta}_{-i}) = 0.095$ , and indicates that there is little sorting in our data.<sup>11</sup> Unobserved heterogeneity of the workers may not be considered a problem—it is captured by the worker fixed effect—as far as we can safely assume that this fixed effect is time-invariant.

We normalize the firm-fixed component to reflect deviations from the industry average and we characterize workers by whether they worked in firms that paid above the average (“high-wage firms”) and those that paid below the average (“low-wage firms”):

$$I(\phi_j) = \begin{cases} 1 & \text{if } \phi_j > 0 \\ 0 & \text{if } \phi_j \leq 0 \end{cases}, \quad (5)$$

and we re-write Eq. (2) as

$$h_{z,t} = 1 - \exp\left(-\exp\left(m_z + \beta_0 + X_{it}\beta + I(\phi_j)\right)\right). \quad (6)$$

Although post-unemployed wages were on average higher for workers with previously high firm-fixed components than for those with low firm components, we see that workers from “high wage firms” experienced a relative wage loss and those from “low wage firms” a relative wage gain. Workers from “high wage firms” were slightly older than workers coming from “low wage firms”, and they had on average shorter tenures.

<sup>9</sup> We use Ouazad's (2008) Stata module. Standard errors are obtained via bootstrapping. Detailed estimation results are shown in the appendix.

<sup>10</sup> 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.

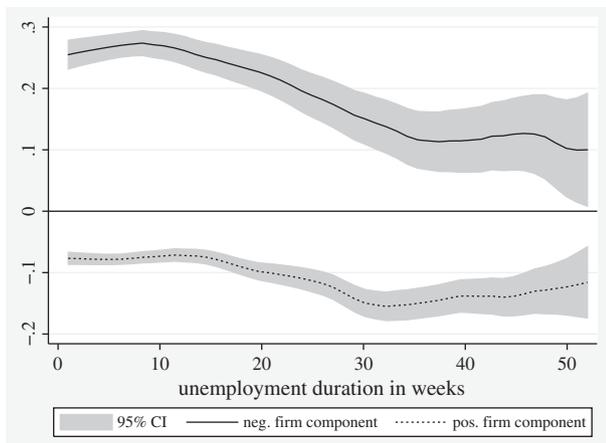
<sup>11</sup> We bootstrap the correlation using 50 replications.

If our decomposition of wages is valid, firm-fixed wage components are random and workers who worked in firms with high firm wage components should lose the firm component, and vice versa. This is, in fact, what we find. For both males and females, workers who worked in low-wage firms have on average higher wages in their new jobs, and workers from high-wage firms have lower wages in their new jobs. Fig. 1 looks at this pattern in more detail. For each elapsed unemployment duration, we plot the average wage change between the old and new job, distinguishing between workers who earlier had worked in low-wage or high-wage firms. Workers who had enjoyed positive firm-fixed components in the past clearly have lower wages in their new jobs, independent of elapsed unemployment duration. In contrast, workers who had worked in low-wage firms experience wage increases at shorter unemployment durations. For longer unemployment durations, they also experience a wage loss, however, this is less pronounced than for workers who had worked in high-wage firms.<sup>12</sup> Overall, workers who remain unemployed longer than about 20 weeks face a negative trend in wages, which is probably a combination of selection and stigma effects.

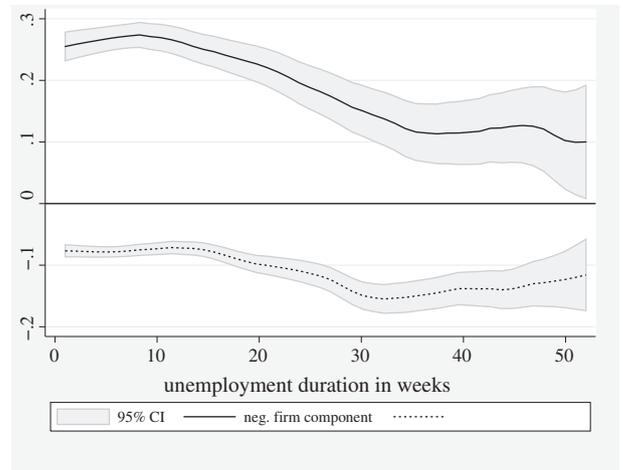
Fig. 2 indicates that the convergence in wages is, in fact, driven by a convergence in firm-fixed components. We see large increases in firm components – up to 25% – for workers who had worked in low-wage firms and losses of about 10% for those from high-wage firms. These patterns are confirming once more the validity of the wage decomposition procedure. See Table 1 for descriptive statistics.

**4. Results**

We present results from non-parametric discrete-time hazard rate models, estimated separately for men and women, in Table 2. The explanatory variables in all specifications include the replacement rate, the worker's age at the time of plant closure, the (old) firm's size at the time of plant closure, and indicator variables for year, industry and region. The specification presented in columns (1) and (4) does not contain the wage components and serves as a benchmark for our specifications below. The estimates indicate that the higher the replacement rate, the lower the hazard of finding employment, a finding that is consistent with previous research on unemployment durations (e.g., Meyer, 1990).<sup>13</sup> Older workers search somewhat longer than younger workers.



**Fig. 1.** Relative change in wages before and after plant closure by level of pre-displacement firm wage component.



**Fig. 2.** Relative change in firm-fixed components before and after plant closure by level of pre-displacement firm component.

The specification in columns (2) and (5) augments the benchmark specification with the estimated firm- and person-fixed wage components which are introduced as dummy variables indicating below or above average values.<sup>14</sup> Workers who had high firm components in their previous employment have lower hazard rates, indicating that they search longer than comparable workers who had low firm components. The result holds both for men and women; however, for women, the coefficient is not statistically significant at conventional levels. The resulting longer unemployment durations for workers from high-wage firms are compatible with our hypothesis that these workers base their wage expectations not only on their person-specific component, but also on the firm component.

These workers could be characterized as being overconfident of their own abilities and productivity. In other words, they appear to attribute the wage they earned in the past firm largely towards their own capabilities and disregard the randomness which might have played a role in the firm-fixed wage component they enjoyed in the

**Table 1**  
Descriptive statistics by gender and firm component category.

	Male		Female	
	Low <sup>[a]</sup>	High <sup>[a]</sup>	Low <sup>[a]</sup>	High <sup>[a]</sup>
Daily wage old job	40.8 (12.8)	55.0 <sup>[b]</sup> (14.4)	25.9 (8.9)	35.4 <sup>[b]</sup> (11.4)
Unemployment duration	117 (128)	128 <sup>[b]</sup> (130)	126 (134)	123 <sup>[b]</sup> (115)
Daily wage new job	41.1 (14.5)	49.1 <sup>[b]</sup> (15.0)	28.8 (10.8)	32.8 (12.1)
Age	33.6 (9.7)	34.7 <sup>[b]</sup> (10.1)	32.5 (8.8)	32.7 (9.1)
Tenure (days)	1308 (1820)	1267 (1822)	1226 (1563)	1219 (1610)
Workers with high person effect (%)	52.5	55.4	38.2	30.2
% of workers	54.6	45.4	70.9	29.1 <sup>[b]</sup>

Notes: Means (standard deviations in parentheses). Wages are deflated to prices of 1990. [a] “Low” and “high” indicate below and above average firm-fixed components respectively. [b] Difference between “low” and “high” statistically significant at the 5% level.

<sup>12</sup> Due to smaller sample sizes at longer unemployment durations, the confidence intervals are large at durations longer than 30 weeks.

<sup>13</sup> See Lalive et al. (2006) for an analysis concerning different financial incentives in the Austrian context.

<sup>14</sup> Robustness tests when we change the cutoff to median values or to the top 25% or introduce firm rents as a continuous variable are available on request: they are qualitatively and quantitatively very similar.

**Table 2**  
Estimated hazard rates from unemployment to employment, by gender.

	Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)
High firm component (0/1)	−0.057***	−0.054**	(0.027)	−	−0.024	−0.055*
	−	(0.022)		−	(0.025)	(0.030)
High person component (0/1)	−	0.184***	0.217***	−	0.099***	0.098***
	−	(0.017)	(0.022)	−	(0.024)	(0.029)
Replacement rate	−0.032***	−0.024***	−0.032***	−0.031***	−0.025***	−0.031***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age	−0.015***	−0.013***	−0.013***	−0.008***	−0.005***	−0.007***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Mass point	1.334***	−	1.329***	1.244***	−	1.247***
	(0.028)	−	(0.028)	(0.049)	−	(0.049)
P(mass point)	0.647	−	0.659	0.724	−	0.729
	0.0181	−	0.0182	0.0487	−	0.0494
Obs.	16,574	16,574	16,574	10,448	10,448	10,448

Notes: Discrete-time proportional hazard rate models. Additional variables are log(firm size), 5 year, 9 region and 15 industry dummy variables. \*\*\*, \*\* and \* indicate significance at the 1, 5 and 10% level.

last firm. In addition, we find that individuals with high person-specific components 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 association with the hazard rate is therefore to be expected.

The specification in columns (3) and (6) additionally controls for unobserved heterogeneity in the search process by estimating two mass points for the distribution of abilities.<sup>15</sup> In addition to heterogeneity captured by the fixed person effects, which indicate differences in unobserved productivity, the mass points control for unspecific differences in job finding (and accepting) probabilities. It turns out that the inclusion of mass points does not change our estimated coefficients to a large extent. The only exception is the coefficient for the high firm component for females, which is now more precisely estimated. Interestingly, the estimates imply only minor differences for male and female workers, and especially the associations between high firm components and the hazard rate are virtually identical.<sup>16</sup> The estimated coefficient of  $-0.054$  for male workers translates into a 1 percentage point lower hazard rate. This implies that after six months the probability of being unemployed is almost 3pp higher for those workers who enjoyed above average firm-fixed components prior to unemployment.<sup>17</sup>

#### 4.1. Age differences

Results from the psychological literature suggest that older adults have greater insight into the limitations of their knowledge than younger ones (Pliske and Mutter, 1996; Musiélak et al., 2006); in other words, overconfidence might be related to age and the associated differences in postformal cognitive development (Pliske and Mutter, 1996).

We split our population into young, prime-age and old workers. The results are tabulated in Table 3 using the comprehensive

<sup>15</sup> When we include more than two mass points, all additional mass points are not statistically significant at conventional levels.

<sup>16</sup> To assess the consequences of missing information on hours worked in the female estimations we ran the following check: We replaced the firm-fixed wage component in the hazard rate estimations for females with firm-fixed components that were calculated from a male worker sample. Since part time work is negligible for male workers the calculation of firm components in a male-only sample is not affected by missing information of hours worked. Our results are robust to this procedure.

<sup>17</sup> We calculate the survival rate for a representative person with mean characteristics (mean age, tenure, replacement rate) for a person with and one without high firm-fixed components. This results in a hazard rate of 0.20 and 0.21 for workers with low and high firm-fixed components.

specification with mass points. While all other coefficients—in particular the person fixed effect and the benefit replacement rate—have almost exactly the same influence across age groups, the effect of the firm-fixed component differs across age groups. Only for workers below the age of 30 we find a negative and significant effect of high firm component on the hazard of leaving unemployment. This is consistent with other patterns of overconfidence by young people as know e.g., from traffic accidents.<sup>18</sup> Repeating the exercise from above we find that young male workers with high firm-fixed components experience an almost 3pp lower hazard rate translating into a 5pp higher probability of being unemployed after 6 months.

More importantly perhaps, 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 which were caused by firm components; Saint-Paul's Saint-Paul (2009) argument for the unemployability of older workers due to misguided reservation wages is not supported by our evidence.

#### 4.2. Overconfident or inexperienced?

In our interpretations above we stressed overconfidence, i.e., attributing spurious wage components to one's own ability or effort, as the cause for the differences in unemployment durations. However, our empirical results could also be driven by systematic errors in deriving the reservation wages, for example, because workers with long tenures are less informed about the state of the labor market than workers who have short tenures. Workers with long tenures may have inferior knowledge of the relevant wage offer distribution and might therefore put too much emphasis on past wages. Alternatively, 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. To investigate this issue, we separate our sample into workers with short and long tenures in the previous firm.<sup>19</sup> A tenure is short if it lasted up to 500 days.

The results are tabulated in Table 5. For males, we do not see any difference between short- and long-tenured workers. Young workers do search longer if they came from a high-wage firm, regardless of

<sup>18</sup> Young drivers' higher probability of being involved in car accidents is found to be linked to young drivers being overconfident in their own driving abilities. See e.g., Brown (1982) and Rumar (1985).

<sup>19</sup> Due to small sample sizes, we pool prime age and older workers.

**Table 3**  
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*** (0.042)	0.016 (0.041)	0.007 (0.074)	−0.103** (0.048)	−0.028 (0.045)	0.063 (0.083)
High person component (0/1)	0.202*** (0.033)	0.257*** (0.033)	0.263*** (0.059)	0.152*** (0.042)	0.028 (0.045)	0.202** (0.087)
Replacement rate	−0.032*** (0.001)	−0.032*** (0.001)	−0.033*** (0.002)	−0.033*** (0.001)	−0.029*** (0.001)	−0.030*** (0.002)
Age	0.004 (0.006)	−0.001 (0.004)	−0.107*** (0.007)	−0.034*** (0.007)	0.005 (0.004)	−0.080*** (0.020)
Mass point	1.233*** (0.044)	1.287*** (0.045)	1.471*** (0.078)	1.375*** (0.066)	1.077*** (0.084)	1.010*** (0.145)
P(mass point)	0.669 0.0301	0.697 0.0294	0.538 0.0541	0.714 0.0347	0.738 0.0576	0.609 0.189
Obs.	6785	6994	2795	4460	4609	1379

Notes: Discrete-time proportional hazard rate models corresponding to columns (2) and (5) in Table 2. Additional variables as in Table 2. \*\*\*, \*\* and \* indicate significance at the 1, 5 and 10% level.

their tenure; prime-age and old workers do not search differently with respect to low-wage or high-wage firms. This is strong evidence against the misconception argument. For females, the results are similar: while we do find firm-fixed component effects for young short-tenured workers, the effect is smaller and insignificant for females with a longer job tenure in the past. If long tenures had caused the misconception, we would have expected the opposite result. Again, there is no effect for prime-age or old women.

A similar test concerns the number of previous jobs, where workers who were more exposed to *realized* rather than *offered* wages might have a more realistic perception of the wage offer distribution. We proxy this exposure by the number of previous jobs and define a worker who had 5 or more previous jobs as a “job hopper” and one with fewer than 5 previous jobs as a “stayer”. The lower panel in Table 5 reports the estimated coefficients for these groups of workers, interacted with age. Again, we obtain large negative coefficients for young male workers, regardless whether they were stayers or hoppers, indicating that high firm components have a distortive effect on their wage expectations. While the coefficient for stayers is somewhat higher as compared to hoppers, the difference is not statistically significant. For females, the results are similar; both young job hoppers as well as stayers face longer unemployment durations in case of higher previous firm fixed

components—but the effect for stayers is statistically insignificant. Overall these tests imply that our results are not driven by young workers' lack of job search experience.

It might be argued that Bayesian updating might explain our results as well: Workers who become unemployed need to learn the distribution of wage offers. Workers who previously earned high wages may form a different prior relative to workers who earned low wages and will have to update their beliefs more often before they achieve a realistic belief of the wage offer distribution. Our decomposition of wages into a firm-specific and individual-specific part speaks against this hypothesis: A high firm wage component prolongs unemployment durations, but a high individual-specific part reduces unemployment durations. Workers do not appear to form a simple prior of the wage distribution where they relate previous wages to potential offers. (However, it might be possible that workers are using a Bayesian updating strategy, if only for the firm-specific wage component.)

A further alternative explanation is liquidity constraints. Card et al. (2007) show that workers search longer if they receive cash-on-hand on becoming unemployed, e.g., from severance pay. Severance pay in Austria is awarded according to tenure with the firm. In Table 5 we see that the effect of high firm-specific components is fairly similar for workers who had short or long tenure. We therefore believe that

**Table 4**  
Estimated hazard rates from unemployment to employment, including early leavers.

	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)	−0.109*** (0.032)	0.002 (0.030)	−0.040 (0.050)	−0.112* (0.066)	0.012 (0.039)	−0.059 (0.046)
High person component (0/1)	0.226*** (0.029)	0.222*** (0.026)	0.213*** (0.044)	0.130*** (0.040)	−0.016 (0.043)	0.065 (0.079)
Replacement rate	−0.033*** (0.001)	−0.032*** (0.001)	−0.033*** (0.001)	−0.035*** (0.001)	−0.030*** (0.001)	−0.028*** (0.001)
Age	0.020*** (0.005)	0.000 (0.003)	−0.109*** (0.006)	−0.025*** (0.007)	−0.001 (0.004)	−0.077*** (0.017)
Mass point	1.321*** (0.034)	1.403*** (0.036)	1.538*** (0.056)	1.460*** (0.053)	1.244*** (0.055)	1.054*** (0.128)
P(mass point)	0.633 (0.0216)	0.733 (0.0189)	0.597 (0.0364)	0.666 (0.0312)	0.712 (0.0343)	0.686 (0.109)
Obs.	10,125	11,152	5038	5208	6192	1963

Notes: Discrete-time proportional hazard rate models corresponding to columns (2) and (5) in Table 2. Additional variables as in Table 2. \*\*\*, \*\* and \* indicate significance at the 1, 5 and 10% level.

**Table 5**

Estimated effect of high firm-fixed component (0/1) on the hazard rates from unemployment to employment for male and female workers, by pre-displacement tenure and number of pre-displacement jobs.

	(1)		(2)		(3)		(4)	
	Male		Female		Male		Female	
	Young	Prime age/old						
Short tenure <sup>a</sup>	-0.146*** (0.051)	-0.018 (0.049)	-0.128*** (0.062)	0.054 (0.059)	-0.128*** (0.062)	0.054 (0.059)	-0.128*** (0.062)	0.054 (0.059)
Long tenure <sup>b</sup>	-0.167*** (0.075)	0.038 (0.053)	-0.051 (0.075)	-0.050 (0.052)	-0.051 (0.075)	-0.050 (0.052)	-0.051 (0.075)	-0.050 (0.052)
Hopper <sup>2</sup>	-0.136*** (0.053)	0.029 (0.042)	-0.093* (0.053)	-0.005 (0.044)	-0.093* (0.053)	-0.005 (0.044)	-0.093* (0.053)	-0.005 (0.044)
Stayer <sup>2</sup>	-0.173*** (0.072)	-0.106 (0.069)	-0.123 (0.104)	-0.066 (0.088)	-0.123 (0.104)	-0.066 (0.088)	-0.123 (0.104)	-0.066 (0.088)

Notes: Discrete-time proportional hazard rate models corresponding to column (3) in Table 2. Only the coefficients for high firm-fixed components are reported, additional variables as in Table 2 (Abowd et al., 1999). A tenure is short if it was shorter or equal to 500 days (Abowd et al., 2004). Workers with less/more than 4 different previous jobs are defined as stayers/hoppers. \*\*\*, \*\* and \* indicate significance at the 1, 5 and 10% level. a and b indicate the copy editor included the first 2 references (Abowd et al.) instead.

liquidity constraints do not explain the longer unemployment durations of workers who earned high firm specific wage components.

#### 4.3. Are plant closures random (enough)?

Because we use workers who are remaining with the firm until the very end of the firm, we may have a selected sample of workers who have below average characteristics of the unemployed. ("Good" workers may have had other opportunities and left prior to the firm's closing-down.) This aspect has been studied with similar Austrian data by Schwerdt (2011), who suggests that an analysis that uses plant closures to obtain an unbiased sample of the unemployed should experiment with different sampling periods prior to plant closure.

We therefore repeat the analysis including all employees who have been laid off from the firm within six months before plant closure. Extending the sampling period from two to six months increases out sample size to 26,315 male and 13,363 female workers experiencing at least one day of unemployment after lay-off.

Table 4 shows that our results are remarkably stable and, again, young males and females have significant longer unemployment durations if they had high firm-fixed components. The coefficients are similar in size to the ones in Table 3. It is therefore unlikely that our result of a negative impact of high firm components on unemployment durations for younger workers is caused by workers who left the plants early.

## 5. Conclusion

Assessing one's own productivity is important for job search and matching in the labor market. A realistic perception of one's productivity will enable the job searcher to match efficiently with an employer. According to psychological research, workers often attribute (excessively) high wages to their own abilities rather than to pure luck in obtaining employment with a firm that pays high firm-specific wage components. Such unrealistic assessment could result in a systematic misjudgment of the wage offer distribution that a job searcher faces and might lead to a distorted job search process.

We study job search behavior of workers who were made redundant due to plant closures in Austria and find that only young workers can be characterized as being overconfident: high firm-fixed components in the last job lead to significantly longer unemployment durations. We do not find such a pattern for prime-age or older workers. These results challenge the view that

the high unemployment rates of older workers in Europe are due to excessive wage claims, which in turn are caused by a slow adaptation to changing circumstances.

Several areas of scientific research have pointed to overconfidence as a source of costs: Finance (Malmendier and Tate, 2005), Medicine (Berner and Graber, 2008), and even sports (McGraw et al., 2004). McGraw et al. (2004) show that people can increase their utility by reducing their overconfidence. In particular, they demonstrate that simply providing additional information increases utility by lowering confidence without negative effects on performance. Such additional information could easily be provided by job centers, for example, starting wage distributions by elapsed unemployment durations or by previous wages. However, accurate information might not be the single solution to overconfidence. Another, more costly approach might be to provide feedback on wage expectations, e.g., a web based service provided by job centers. However, Eberlein et al. (2005) caution that some people ignore feedback more or less completely, while others become confused. In any case, a large scale implementation of such services requires careful testing and additional research.

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