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Relative Wage Positions and Quit Behavior:

New Evidence from Linked Employer-Employee Data *

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Abstract

We use a large linked employer-employee data set to analyze the importance of relative wage positions in the context of individual quit decisions as an inverse measure of job satisfaction. Our main findings are: (1) Workers with higher relative wage positions within their firms are on average more likely to quit their jobs than workers with lower relative wage positions; and (2) workers, who experience a loss in their relative wage positions, are also more likely to have a wage cut associated with their job-to-job transition. The overall results therefore suggest that the status effect is dominated by an opposing signal effect.

JEL-Classification: D03, J31, J62, J63, M52

Keywords: comparison income, mobility, signaling, status, wages

1. Introduction

The empirical analysis of the impact of relative wage positions on workers' decisions to voluntary quit their job in the current firm is important in the context of two streams of the economic literature. On the one hand, the recent labor turnover literature points to the importance of fair wages and status concerns of workers as well as to the paradox that many workers experience a wage cut after a job-to-job transition (e.g., Galizzi and Lang, 1998; Postel-Vinay and Robin, 2006; Jolivet et al., 2006). On the other hand, our results can be incorporated into the broader literature about interdependent preferences and the determinants of subjective well-being (e.g., Hamermesh, 1975; Frank, 1985; Easterlin, 1995; Solnick and Hemenway, 1998; Clark et al., 2008), because quits are driven to some extent by utility maximizing behavior. With respect to both streams of the literature, we can contribute new empirical findings from linked employer-employee data, which allows us to compute measures for workers' relative wage positions within their firms and to assess their impact on important decisions in real world data, namely decisions to quit full-time employment. Our sample contains almost four million yearly observations of more than one million full-time employed male prime-age workers in nearly seven thousand West German firms for the period from 1996 to 2005.

Our main results are that relative wage positions have a significant impact on the probability to voluntary quit a job and on the probability to accept a wage cut when changing firms. We find that a potential status effect is dominated by a potential signal effect, because workers with higher relative wage positions within their firms are on average more likely to quit a job than workers with lower relative wage positions. The former might expect fewer opportunities for further career advancement in their current firm so that they switch to a different firm. This seems to be the case even if they have to accept a short-term wage cut in exchange for long-term career opportunities. Overall, we find evidence that workers' quit probabilities as inverse measure for utility (or subjective well-being) are not strongly influenced by status concerns. The opposing signal effect seems to be of much larger importance than the status effect, which has also been discussed in a recent study by Clark et al. (2009), who find in Danish linked employer-employee data that a worker's satisfaction is on average higher if co-workers earn higher mean wages holding the worker's own wage constant.

The paper is structured as follows. The next section illustrates briefly the basic theoretical framework and our research hypotheses. Section 3 describes our data set, main variables, and econometric models. In Section 4 we present our econometric results for the impact of relative wage positions on the individual quit probability and for the consequences of quits on absolute wages and relative wage positions. We conclude with a short summary in Section 5.

2. Basic theoretical framework and hypotheses

The relationship between wages and individual quit behavior can be modeled in the framework of utility maximizing worker behavior. Utility *U* of individual *i* who works in firm *j* in period *t* in equation (1) is a simplified function of the individual absolute wage (w_{ijt}^{abs}), the individual relative wage position within the firm (w_{ijt}^{rel}), the relative wage position across firms (w_{it}^{rel}), and other individual and job characteristics (X_{ijt}). Moreover, we assume that the individual probability to voluntary quit a job in firm *j* in period t is negatively correlated with utility as described in equation (2) (Freeman, 1978; Akerlof et al., 1988; Clark et al., 1998; Clark, 2001; Clark and Georgellis, 2006; Lévy-Garboua et al., 2007).

$$(1) \quad U_{ijt} = U_{ijt}\left(w_{ijt}^{abs}, \underbrace{w_{ijt}^{rel}}_{+/-}, \underbrace{w_{it}^{rel}}_{+}, X_{ijt} \right)$$

$$(2) \quad \Pr\left(Quit_{ijt} = 1 \mid \underbrace{U_{ijt}}_{-} \right) = \Pr\left(Quit_{ijt} = 1 \mid w_{ijt}^{abs}, \underbrace{w_{ijt}^{rel}}_{-/+}, \underbrace{w_{it}^{rel}}_{-}, X_{ijt} \right)$$

Standard economic theory (e.g., search models, efficiency wage models) accounts usually only for absolute wages, which should positively affect a worker's utility (Salop and Salop, 1976; Salop, 1979; Akerlof, 1982). Our main focus is however on workers' relative wage positions, which have received increasing attention in happiness research in the last two decades (Clark et al., 2008). The impact of the relative wage position within a firm (w_{ijt}^{rel}), which includes wages of co-workers as comparison income, is however ambiguous (Clark et al., 2008; Clark et al., 2009). If the individual wage is held constant, higher wages of co-workers are associated with a lower relative wage position of an individual worker within his firm. On the one hand, a lower relative wage might be perceived as unfair and of low social status (Adams, 1965; Frank, 1984a, 1984b; Garner, 1986; Akerlof and Yellen, 1990; Clark et al., 2009), which consequently decreases utility and increases the quit probability, ceteris paribus. This is called the 'status effect'. On the other hand, the relative wage position within a firm can also cover a 'signal effect' as it provides workers with information about their own future income and career prospects (Hirschman and Rothschild, 1973; Senik, 2008; Clark et al., 2009). Higher wages of co-workers might signal better career prospects in the firm, which increases utility and decreases the quit probability. If a worker is however already high up in the pay scale, he cannot expect to have further career advancements in the current firm and consequently he might decide to quit his job and to join another firm.[1]

Hypothesis 1a: Workers are less likely to quit their job if they have a higher relative wage position within their firm ('status effect'>'signal effect').

Hypothesis 1b: Workers are more likely to quit their job if they have a higher relative wage position within their firm ('signal effect'>'status effect').

Job utility and quit probability are certainly not only affected by the present individual absolute wage and the relative wage position within the own firm but also by outside wages (Stiglitz, 1974; Clark and Oswald, 1996; Galizzi and Lang, 1998; Kim, 1999; Fairris, 2004; Bingley and Westergard-Nielsen, 2006). A low wage in comparison to workers with similar characteristics in different firms (low relative wage position across firms (w_{it}^{rel})) implies that a worker can gain from quitting his current job and moving to another firm for three reasons. First, workers should be, ceteris paribus, more likely to change firms if they can earn higher absolute wages in other firms. Second, workers might perceive their wages unfair and of lower status if comparable workers in other firms earn higher relative wages. Third, a firm, in which workers earn higher relative wages than in other firms, might signal better career prospects so that workers of other firms might be convinced to join the 'high wage' firm. Overall, workers of 'high wage' firms, who have already a higher relative wage position than workers in other firms, have lower incentives to quit as they cannot gain much.

[1] See Clark et al. (2009) for an extensive discussion of status and signal effects.

Hypothesis 2: Workers are less likely to quit their job if they have a higher relative wage position across firms.

In a further step, our paper aims to shed some more light on the empirical observation that many mobile workers experience wage cuts. Table 1 outlines some results of recent studies on wage cuts induced by mobility. Mobility to a lower wage is common in the U.S. and Germany.[2] Fitzenberger and Garloff (2007) report for Germany that about one quarter of mobility events is associated with a wage cut. Jolivet et al. (2006) find that about 36 percent of job-to-job transitions in Germany and 23 percent of transitions in the U.S. are to lower wages. Nosal and Rupert (2007) provide evidence for the U.S. that about two out of five (voluntarily) mobile individuals change to lower wages.

[Insert Table 1 about here]

Theoretical approaches explain voluntary wage cuts mostly as investments in future wage growth (Galizzi and Lang, 1998; Postel-Vinay and Robin, 2002; Connolly and Gottschalk, 2009). These approaches have in common that individual decisions are reduced to a monetary maximization problem, in which workers maximize the long-run value of job opportunities. Nosal and Rupert (2007) suggest furthermore that job-specific (non-wage) amenities affect the job choice and consequently individual mobility decisions. We expect relative wage positions to be one such amenity because of its fairness and status aspects. Thus, relative wage positions within a firm should affect the voluntary acceptance of wage cuts. The total effect is again ambiguous due to counter acting signal and status effects. On the one hand, workers might accept wage

[2] See Jolivet et al. (2006) for a larger set of countries.

cuts if they can improve their status in the new firm, which is measured as a higher relative position in the new firm. On the other hand, workers might be more likely to accept a wage cut if they have better career prospects in the new firm, i.e., higher future wages, which is signaled by a lower current relative wage position.

Hypothesis 3a: Workers trade off absolute wages and relative wage positions when changing firms ('status effect'>'signal effect').

Hypothesis 3b: Lower absolute wages and lower relative wage positions go hand in hand when workers change firms ('signal effect'>'status effect').

3. Data and methodological remarks

3.1 Data set

As we are interested in relative wage positions within firms, our estimation framework requires information about workers, co-workers, and their firms. It is furthermore desirable that the data set contains not only a small subsample of workers in each firm but as many workers as possible so that relative wage positions in each firm can be computed accurately. The German linked employer-employee data set of the Institute for Employment Research (‘Institut für Arbeitsmarkt und Berufsforschung (IAB)’, LIAB in the following) fulfills these prerequisites (Alda et al., 2005). The LIAB links employer side information from the IAB Establishment Panel with employee information from process-produced person specific data. The IAB Establishment Panel is a yearly survey that includes a random sample of firms with at least one employee covered by social security. The sample is drawn from stratification cells of establishment size classes and industries. The firms are asked about their employment structure, personnel policy, industrial relations etc. The process-produced person specific data stem basically from the notification procedure for unemployment, pension, and health insurances. Employers have to notify the social security agencies about all employees that are covered by social security at the start and at the end of an employment relationship as well as on the last day of each year.

The underlying data set is set up as a panel of cross-sections from 1993 to 2006 at the corresponding record date of June 30. In the last period of the sample, the individuals cannot exactly be assigned to be movers or non-movers because we do not observe their subsequent employment status. The year 2006 is hence not subject to the analysis. Moreover, our analysis focuses on the years from 1996 onwards because sample size was considerably enlarged and information about collective contracts are only available for this time horizon. In sum 10 periods are available for our study.

We restrict the sample to male full-time workers in the main job aged between 30 and 55 years. Winkelmann (1994) shows that German workers hold on average four lifetime jobs and half of all lifetime job transitions are executed in the first ten years of their careers. Job shopping provides one interpretation for this pattern (Topel and Ward, 1992). During the beginning of the career, young individuals learn about their abilities and productivity, the employer-employee match quality, and their own fair market wage. As we are interested in the effect of relative wage positions rather than learning related job shopping, we focus on workers exceeding 30 years of age. Analogously to Galizzi and Lang (1998), we limit the sample to men who are less than 55 years old as those workers might be more concerned with retirement decisions. Furthermore, only full-time employed German citizens are included because no information about working hours is available in the data. Our analysis concentrates on establishments located in West Germany because of different labor market conditions in East and West Germany (e.g., unemployment, wages) and the fact that our data contain mostly West German firms. A methodological reason for the restriction is that some of the control variables are left-censored before unification in 1990. The data consequently report only a lower boundary for tenure and experience in East Germany.

The data reveal some implausible low daily wages (Jacobebbinghaus, 2008). The analysis is responsive to these values and excludes wages below the marginal employment ceiling of 400 Euros per month. Hence, employees who earn less than 13.33 Euros per day are dropped from the analysis. Another problem of the data is that wages are censored at the upper earnings limit for social security contributions. This implies that all wages above this ceiling are set to the corresponding value of the ceiling. To reduce the impact of the censoring, the sample is restricted to workers who do not have more than a high-school degree with an apprenticeship degree.[3]

Establishments with less than ten workers under the above restrictions are not subject to the analysis because we need to estimate earnings functions for single firms (degrees of freedom) and need sufficient wage variance within firms for our analysis of relative wage positions. Consideration of more than ten observations decreases the number

[3] Note that imputation methods are available but imputation procedures increase the uncertainty about the relative wage positions of workers within an establishment. Moreover, it seems questionable if regular workers compare themselves with high wage employees in upper management positions.

mobility events rapidly.[4] The final sample for our analysis contains 3,867,569 yearly observations from 1,115,437 workers in 6,791 different firms in an unbalanced panel design for the period from 1996 to 2005. Only few observations (0.15 percent) comply with the upper earnings limit for social security contributions in our sample so that wage censoring is not of much concern.

Our analysis of quit behavior relies on the assumption that individuals leave their employer voluntarily. Voluntary transitions are defined as an unconstrained choice of the worker. This criterion is hardly to implement using this data set, but the identification of voluntary quits is crucial for our analysis. The following conditions need to be met for the identification of quits. The worker is full-time employed in two successive periods in two different establishments. To assure that mobility is likely to be induced by the worker and not by the firm, information on the individual's employment relationship eight days before the new job started is consulted. If the worker was full-time employed at another establishment eight days before entering the new establishment, he is assumed to have voluntarily quit the job at his past employer because he switched establishments with virtually no unemployment spell.[5] Other types

[4] Table A.1 in the appendix presents number of mobility events and number of observations for different samples with respect to annual observations per establishment.

[5] Following our definition of quits we cannot assure that all but that most transitions are voluntary. Jolivet et al. (2006, p. 882) note: "Surely, many of the quick job re-accessions at very short durations correspond to voluntary job changes [...]. Yet some of them are likely to reflect involuntary reallocation - essentially job losses followed by the immediate finding of a replacement job." Note that only 0.2 percent of the observations in our sample experience quits. This very low share is reasoned by the nature of our data set and voluntary quit variable, which is defined as changes from one firm to another firm in our data set. As

of mobility than voluntary quits (e.g., layoffs) are included for the computation of comparison wages, but they are excluded from the regression analysis of the determinants of quits. In sum, 7,785 mobility events are observed and 7,516 individuals are mobile up to four times.

3.2 Wage measures

We have introduced three general wage variables in the theory section, which need to be generated from the data. At first, the individual absolute wage (w_{ijt}^{abs}) is measured straightforward and is the log mean daily wage in Euros of individual *i* in firm *j* in year *t*. Moreover, we construct five different measures to analyze the relative wage position within a firm. Following the literature (Freeman, 1978; Akerlof et al., 1988; Topel and Ward, 1992; Clark and Oswald, 1996; Galizzi and Lang, 1998; Clark et al., 2009), the average wage of workers in a firm ($\bar{w}_{jt}$) is used as comparison income.[6] Holding the individual wage constant, an increase of the average wage is associated with a lower individual relative wage position. As a second measure for comparison income, we use predicted inside wages ($\hat{w}_{ijt}^{inside}$) obtained from separately estimated earnings functions

our sample contains a little less than one percent of the entire relevant population of firms in West Germany, we can only observe a low share of quits. Our randomized sample of firms should however mitigate this possible problem, because workers with observed quits should not be different from workers who voluntary change to firms not included in our sample.

[6] The data reports a lower boundary of the average wage within the establishment because of the censoring. As only 0.15 percent of our sample is censored, this problem can be neglected in our analysis.

for every firm in every year.[7] Included worker characteristics are schooling, potential experience, squared experience, and occupation.[8] The earnings function looks as in equation (3), in which α denotes the firm-specific constant, γ the coefficients for worker characteristics X, and ε the residual term.

$$(3) \qquad w_{ijt}^{inside} = \alpha_j + \gamma' X_{it} + \varepsilon_{it}$$

We also construct measures which might be intuitively more appealing in the context of relative wage positions as they actually measure the individual wage position. Following Brown et al. (2008)[9], we construct the wage rank as well as the wage range of a worker within his firm so that both variables lie in the unit interval (0, 1). Values of one indicate that the individual is at the top of the pay scale.[10] The wage rank measures

[7] This approach closely follows Clark and Oswald (1996) and Senik (2008), who include predicted wages conditional on schooling, occupation, sector, region, and other variables in satisfaction equations. The authors interpret the predicted wages as comparison income of individuals.

[8] We do not include tenure in these estimates because the comparison group should also include comparable workers at later career stages (career prospects). Descriptive statistics are presented in Table A.2 in the Appendix.

[9] Brown et al. (2008) draw on insights from research in psychology and the range frequency theory (Parducci, 1965) to analyze the impact of wage positions within a firm on workers' satisfaction with different job related items. They find that workers with higher relative wage positions are more satisfied with their pay, influence, achievement, and respect.

[10] The exact values cannot be computed due to the upper censoring of wages, which might lead to a compression of our rank and range measures. Since only 0.15 percent of our sample is censored, this issue is not of large concern in our analysis.

the normalized rank of individual i in firm j in period t as proportion of the number of workers in firm j in period t ($w_{ijt}^{rank} = (\text{'rank of worker } i \text{ in } j \text{ in } t\text{'} - 1) / (\text{'number of workers in } j \text{ in } t\text{'} - 1)$). A higher rank indicates that the worker is higher up the pay scale in his firm. The wage range measures the normalized distance of individual i's wage in firm j in period t to the maximum wage in his firm as proportion of the wage spread between the highest and the lowest wage in the firm ($w_{ijt}^{range} = (w_{ijt} - w_{jt}^{\min}) / (w_{jt}^{\max} - w_{jt}^{\min})$). The individual wage rank indicates in an ordinal sense and the individual wage range in a cardinal sense a worker's position in his firm's wage hierarchy. The impact of both variables can be compared to assess if the ordinal rank (wage rank) or the cardinal rank (wage range) is more important to workers (Fields and Fei, 1978; Brown et al., 2008).

A further measure of the relative wage position within a firm, which is very closely related with the previous two measures and especially with wage rank, is calculated on the basis of the empirical cumulative distribution function (CDF) for each establishment in each period (w_{ijt}^{CDF}). Equally paid workers get the same cumulative value. Analogously to wage rank and wage range, a larger value implies a higher relative wage position within the firm and the variable is restricted to the unit interval (0, 1).

At last, the relative wage position across firms is measured as the predicted comparison wage. In an ideal setting we would be able to observe the full distribution of individual outside wage offers a worker can choose from. As this is not the case in reality, we solve this problem by estimating an earnings function across all individuals in all firms for every year and then predict the outside wage ($\hat{w}_{it}^{outside}$) for every individual by the obtained results. The predicted wage can also be interpreted as an expected outside wage offer, i.e., the average wage a worker with the same characteristics in the same sector and in the same geographical area earns. The earnings function looks as in equation (4), in which α denotes the annual constant, γ the coefficients for worker characteristics X (schooling, potential experience, squared experience, and occupation)[11], δ the coefficients for sector S and geographical area A, and ε the residual term.

$$(4) \qquad w_{it}^{outside} = \alpha + \gamma' X_{it} + \delta_1 S_{it} + \delta_2 A_{it} + \varepsilon_{it}$$

Table 2 summarizes the definitions of our wage measures. Table 3 presents means, standard deviations, and the correlations between the constructed wage measure. As already noted by Brown et al. (2008, p. 372), the wage measures are of course somewhat correlated and contain quite similar information. Therefore, we only account for one of the relative measures in a single specification when estimating the determinants of quits and compare their effects. In case of w_{ijt}^{rank} and w_{ijt}^{range}, we compare the impact of ordinal and cardinal ranks. Furthermore, w_{ijt}^{CDF} provides a valuable robustness check on the effect of the ordinal wage rank.

[Insert Table 2 about here]

[Insert Table 3 about here]

[11] We did not incorporate tenure or workplace characteristics in the earnings function because our aim is to predict individual wages across firms and to use the predictions as comparison income. Descriptive statistics are presented in Table A.2 in the Appendix.

3.3 Econometric models

Our basic estimation framework looks as in equation (5), in which *Quit** denotes the latent individual quit probability, α the constant, β the coefficients of our wage variables w_{ijt}^{abs} and w_{ijt}^{r} for which we incorporate the different relative wage measures discussed in the previous section ($\overline{w}_{jt}$, $\hat{w}_{ijt}^{inside}$, w_{ijt}^{rank}, w_{ijt}^{range}, w_{ijt}^{CDF}, $\hat{w}_{it}^{outside}$), γ the coefficients of worker characteristics X (schooling degree, tenure, squared tenure, potential experience[12], squared experience, occupation), δ the coefficients of firm characteristics Y (share of unskilled workers in establishment, number of employees, works council, collective contract, sector, federal state), λ time fixed effects, and ε the remaining residual term. For descriptive statistics of the variables see Table A.2 in the Appendix.

$$(5) \quad Quit_{ijt}^{*} = \alpha + \beta_1 w_{ijt}^{abs} + \beta_2 w_{ijt}^{r} + \gamma' X_{it} + \delta' Y_{jt} + \lambda_t + \varepsilon_{ijt}$$

As the quit probability (*Quit**) cannot be observed but only the actual quit behavior (see equation (6)), our dependent variable is binary and we apply the individual random effects Probit model in equation (7), in which Φ is the cumulative density function of the standard normal distribution and ν_i is the individual random effect.[13]

[12] Potential experience is calculated with respect to individual labor market entry. Hence, possible spells in unemployment or apprenticeships directly following after school are accounted for.

[13] Since quits are a rare event, linear models cannot be applied. Moreover, the incidental parameter problem arises for individual fixed effects Probit models because our panel is too short with an average panel length of 3.5 years (Heckman, 1981). We thus apply an individual random effects Probit model to

$$\text{(6)} \quad Quit_{ijt} = \begin{cases} 1 & \text{if worker } i \text{ quits his job in firm } j \text{ in period } t \\ 0 & \text{if worker } i \text{ stays in firm } j \text{ in period } t \end{cases}$$

$$\text{(7)} \quad \Pr\left(Quit_{ijt} = 1\right) = \Phi\left(\alpha + \beta_1 w_{ijt}^{abs} + \beta_2 w_{ijt}^{r} + \gamma^{'} X_{it} + \delta^{'} Y_{jt} + \lambda_t + \nu_i\right)$$

In addition to the determinants of individual quit behavior, we also analyze the consequences of quits. Many empirical studies report a large share of workers who experience an individual wage loss when changing firms, which might be explained by factors like future wage growth, non-pecuniary rewards, and other job characteristics (Bartel and Borjas, 1981; Bartel, 1982; Ruhm, 1987; Akerlof et al., 1988; Polsky, 1999; Yankow, 2003; Nosal and Rupert, 2007; Connolly and Gottschalk, 2008; Schneck, 2009a). We extend this perspective by our measures for the relative wage position within firms introduced in the previous section (w_{ijt}^{rank}, w_{ijt}^{range}, w_{ijt}^{CDF}). As discussed in Section 2, utility and quit probabilities also depend on status from relative wage positions as well as on signals for career advancement opportunities. For example, a quitting worker might experience a loss in absolute wages but is compensated by a gain in status. It is, therefore, straightforward to compare not only the differences between individual wages in the old and the new firm but also the differences between relative wage measures in the new and the old firm. For this purpose, non-parametric methods like kernel density estimates of the differences can give first insights.

Moreover, it is possible to regress the absolute wage difference on the difference in relative wage positions to assess possible tradeoffs in the utility function. Equation (8)

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exploit the panel nature of the data set. Because of our interest in average comparison wages within a firm, which do not vary across workers in one firm, we do not control for firm fixed effects.

presents the estimation framework, in which $\left(w^{new}-w^{old}\right)_{it}$ is the difference of individual absolute wages between the new and the old firm, α the constant, η the coefficients of the differences in relative wage measures (w_{ijt}^{rank} , w_{ijt}^{range} , w_{ijt}^{CDF}) between the new and the old firm, γ the coefficients of worker characteristics X (schooling degree, change in establishment size class, potential experience, squared experience,), λ time fixed effects, and ε the usual remaining residual term. This estimation framework further allows us to investigate which socio-demographic groups (e.g., low skilled) are more affected by wage cuts.

$$(8)\qquad \left(w^{new}-w^{old}\right)_{it}=\alpha+\eta^{'}\left(w^{r,new}-w^{r,old}\right)_{it}+\gamma^{'}X_{it}+\lambda_{t}+\varepsilon_{it}$$

If workers accept lower absolute wages in the new firm because they are compensated with higher status, i.e., higher relative wage positions, we would expect the coefficients η‘s to be negative. However, if the signal effect of better career opportunities dominates, we would expect the coefficients η‘s to be positive. As robustness check an additional Probit regression for accepting a wage cut can be estimated that is presented in equation (9). The dependent variable takes the value one in case of a wage cut and zero otherwise (see equation (10)) so that the expected signs of the coefficients η‘s reverse compared to the regression in equation (8).

$$(9)\qquad \Pr\left(WageCut_{it}=1\right)=\Phi\left(\alpha+\eta^{'}\left(w^{r,new}-w^{r,old}\right)_{it}+\gamma^{'}X_{it}+\lambda_{t}\right)$$

$$(10)\qquad WageCut_{it}=\begin{cases}1 & \text{if } \left(w^{new}-w^{old}\right)_{it}<0\\ 0 & \text{if } \left(w^{new}-w^{old}\right)_{it}\geq 0\end{cases}$$

4. Econometric results

4.1 Determinants of quits

This section presents the results of the individual random effects Probit model as discussed in equation (7) in Section 3.3. Likelihood ratio tests reject the null hypothesis of no individual unobserved heterogeneity in all specifications, which indicates that the random effects Probit model is more appropriate than the simple cross section Probit model. In the following, we discuss only marginal effects of our wage variables at the means of all covariates and under the assumption that the individual error term is zero. Note that the estimated absolute marginal effects might seem very small and not of economic significance at first glance. As the mean probability is however also very small, the relative marginal effects are in fact quite sizeable.[14] The complete estimation output and the corresponding coefficients are presented in Table A.3 in the Appendix.[15]

The first specification in Table 4 contains the results for the quit probability without relative wage measures. The absolute wage (w_{ijt}^{abs}) has a significant positive effect on the quit probability, which holds also in the next specifications. An increase of the absolute wage by one log point increases the quit probability by about 0.02 percentage points or by about 60 percent, respectively. This result is counter-intuitive as we would

[14] For example, an absolute marginal effect of 0.0001 is a relative marginal effect of 33.3 percent if the mean predicted probability is only 0.0003.

[15] We also performed all subsequent estimates with a subsample of individuals who work in firms with at least one quit. As our main results are robust, the results are only displayed in Table A.4 in the Appendix.

expect that a worker's utility depends positively on his wage. Galizzi and Lang (1998) report also that workers with a higher absolute wage have on average a higher quit probability. One reason might be better outside job opportunities for better paid workers because differences in wages might reflect to some degree unobserved productivity differences in the estimates.

[Insert Table 4 about here]

The next specifications include our variables of main interest, i.e., the relative wage measures discussed in Section 3.2. In these estimates the absolute wage is only a control variable to discuss the results for our relative wage measures from a ceteris paribus perspective, i.e., we interpret the effects when holding the individual absolute wage constant. Specification two includes the average wage in a worker's firm ($\overline{w}_{jt}$). Holding the own wage constant a higher average wage is a higher comparison income, meaning that the own relative wage position is lower. Because the average wage has a negative effect on the individual quit probability, workers with a lower relative wage position have on average a lower quit probability. If the mean log wage in a firm increases by one point, workers' quit probabilities decrease on average by about 0.025 percentage points or about 70 percent, respectively. In the third specification, we include the predicted inside wage ($\hat{w}_{ijt}^{inside}$) as comparison income. The effect is again negative and can be interpreted in the same way as before. These findings correspond with previous findings about quits (Galizzi and Lang, 1998; Bingley and Westergard-Nielsen, 2006).

Specification four includes the wage rank within the firm (w_{ijt}^{rank}), which has a significant positive effect on the quit probability. This result is consistent with our previous estimates for comparison incomes because workers with a higher wage rank have on average a higher quit probability and vice versa. The effect of wage range (w_{ijt}^{range}) in specification five has a negative but not significant effect, which might indicate that the cardinal rank (wage range) is less important than the ordinal rank (wage rank). That the ordinal rank increases the quit probability and is of high significance, is also found in specification six for the position in the wage CDF (w_{ijt}^{CDF}).

Overall, the results show that workers with higher relative wage positions within their firm have on average significant higher quit probabilities.[16] Consequently, we find more support for our Hypothesis 1b than 1a. As we cannot distinguish between the status and the signal effect of relative wage positions, we can only conclude that at least in our sample the signal effect dominates the status effect. Therefore, it seems as workers react more strongly to opportunities for career advancement than to fairness and status concerns. Our findings somehow contradict the results of Brown et al. (2008), who find that workers with higher wage ranks are more satisfied, which would lead to a lower quit probability in our context. Other empirical results for satisfaction are however mixed. Clark et al. (2009) report that individual satisfaction is higher if co-workers earn higher wages. Clark and Oswald (1996) find that a higher comparison income decreases job satisfaction and satisfaction with pay. McBride (2001) and Stutzer (2004) report lower subjective well-being if comparison and aspiration income is higher. Senik (2008) finds mixed evidence for different countries.

[16] Note that differences in relative wage positions are unlikely to reflect unobserved productivity differences as these should be covered by the individual absolute wage.

Specification seven includes the predicted outside wage ($\hat{w}_{it}^{outside}$) as a proxy for the relative wage position across firms and possible outside wage offers. The effect is positive, which can be interpreted in the way that holding the own wage constant, a higher comparison wage across all firms, which is associated with a lower relative wage position in the current firm and the chance of higher earnings in other firms, increases the quit probability. This is in line with our Hypothesis 2 that workers are less likely to quit their job if they have already a higher relative wage position across firms because they cannot gain much from changing firms.[17] Bingley and Westergard-Nielsen (2006) find analogously that predicted alternative wages positively affect the individual quit probability. Fairris (2004) reports evidence from establishment data that firms, which pay on average wages below the industry and geographical area means, have higher quit rates.

4.2 Consequences of quits

In this section, we analyze the consequences of quits with respect to absolute wages and relative wage positions within a firm. More precisely, we are interested in the tradeoff between absolute wages and relative wage positions after a job change and the question whether mobile workers are compensated for wage cuts by increasing their relative wage positions, i.e., by a gain in status. As discussed in Section 2, previous studies have found most voluntary job mobility to be associated with wage gains but also that a

[17] As the standard error is quite large, the effect of the predicted outside wage is however statistically not significant.

significant share of quits is accompanied with wage cuts. In our sample, which includes now only observations with a job-to-job transition between firms, 28.5 percent of workers experience a wage cut when changing the firm. This number has about the same size as in previous studies for Germany (Jolivet et al., 2006; Fitzenberger and Garloff, 2007).

Table 5 presents descriptive statistics about changes in absolute wages $\left(w^{new} - w^{old}\right)_{it}$ and relative wage positions within the new and the old firm ($\left(rank^{new} - rank^{old}\right)_{it}$, $\left(range^{new} - range^{old}\right)_{it}$, $\left(CDF^{new} - CDF^{old}\right)_{it}$). On average workers gain 0.03 log points in wages when changing the firm. The consequences are, however, quite heterogeneous as can be seen from the separated analysis for workers with wage cuts and wage markups. Workers with a wage cut receive on average 0.12 log points lower wages, while workers with a wage markup receive on average 0.09 log points higher wages. For our ordinal wage rank measures w_{ijt}^{rank} and w_{ijt}^{CDF}, we find that the average mobile worker has a lower relative wage position in the new firm. The cardinal wage range measure (w_{ijt}^{range}) is on the other hand slightly positive. One might be tended to misleadingly conclude that the average gain in absolute wages and loss in ordinal relative wage positions is in support of our Hypotheses 3a that workers tradeoff absolute wages and relative wage positions when changing firms ('status effect'). If we look at workers with wage cuts, we see however that those workers also suffer from lower relative wage positions, whereas workers with wage markups also gain in their relative wage positions. Thus, the first descriptive findings are more in line with our Hypothesis 3b that lower (higher) absolute wages and lower (higher) relative wage positions go hand in hand when workers change firms. The rationale behind this finding is that workers are more likely to accept wage cuts if they start at a lower relative wage position in the new firm as they have more space for career advancements ('signal effect').

[Insert Table 5 about here]

In the following, we present Epanechnikov kernel density estimators for changes in relative wage positions, which distinguish between mobile workers with wage cuts and wage markups, to shed some more light into the heterogeneous consequences of quits. Figure 1 displays exemplary the distributions of changes in the wage positions in the CDF, which look very similar for wage rank and wage range. The aforementioned results for the separate samples of movers with wage cuts and wage markups hold in general also here. Most workers who suffer from wage cuts in the period of mobility also lose in relative wage position. In contradiction, most mobile workers with wage markups do not experience much change in their relative wage positions. It can nevertheless be seen that more mobile workers with wage markups gain than lose with respect to the relative wage position. Overall only few workers seem to accept wage cuts in order to improve their relative wage positions and to gain additional status.

[Insert Figure 1 about here]

The kernel density estimators do not account for further determinants of consequences of quits. Thus, we use linear regressions to regress changes in absolute wages on changes in relative wage positions and a set of control variables (see equation (8) in Section 3.3 for the econometric model and Table A.5 in the Appendix for descriptive statistics). The results in Table 6 support our previous findings that changes in absolute wages and changes in relative wage positions are positively correlated. We further estimate Probit models for the determinants of accepting wage cuts (see equation (9) in Section 3.3). Table 7 presents the estimated marginal effects. The results show that workers who improve their relative wage positions are less likely to experience a wage cut.[18]

[Insert Table 6 about here]

[Insert Table 7 about here]

The results in Table 6 and Table 7 also give insights into the question which groups are more likely to suffer from wage cuts. Workers with lower educational levels are more likely to experience a wage cut when changing firms. This finding might be reasoned by bad job opportunities for unskilled workers. Specifications one and three suggest that workers who are mobile to larger establishments[19] are significantly less likely to suffer from wage cuts. This result corresponds with findings about positive wage premiums in larger firms, which are reasoned for example by efficiency wages (e.g., Brown and Medoff, 1989; Idson and Oi, 1999). Potential experience has a negative impact on wage

[18] As a robustness check, we repeated the Probit estimates for a subsample of quitting workers who switch between large firms with at least 1,000 employees. The results show that the relative wage effects are even larger than in the complete sample (see Table A.6 in the Appendix).

[19] Mobility to larger establishments is defined as a binary variable, which takes the value one if the new firm is in a larger establishment size class as defined in Table A.2 in the Appendix. Mobility to larger establishments is quite common as 56 percent of transitions are executed to establishments in a larger establishment size class.

changes associated with quits but an insignificant effect on the probability to accept a wage cut.

In sum, our analysis does not provide evidence that wage cuts are accepted in exchange for an increase in status associated with higher relative wage positions. Workers who suffer from decreasing relative wage positions are in fact also more likely to suffer from wage cuts induced by mobility and vice versa. On the one hand, this finding can be interpreted from the point of view that mobile workers with wage cuts are 'double losers' because of their additional loss in status. If these workers, on the other hand, do not care much about status but about their chance for career advancement, a lower relative wage position might signal such better future career opportunities and consequently the quit decision would be rational. Consistent with this argument, Fairris (2004) finds evidence that firms have on average lower quit rates if internal promotions and seniority are important and job ladders are long. Overall, we find more support for our Hypothesis 3b than 3a because the signal effect seems again to dominate the status effect.

5. Conclusion

Our main results are that relative wage positions have a significant impact on the probability to voluntary quit a job and on the probability to accept a wage cut when changing firms. We find that a possible status effect is dominated by a possible signal effect, because workers with higher relative wage positions within their firms are more likely to quit a job than workers with lower relative wage positions. The former might expect fewer opportunities for further career advancement in their current firm so that they switch to a different firm. This might be even the case if they have to accept a short-term wage cut in exchange for new career opportunities. Workers with lower relative wage positions within their firm have on the other hand still much space for career advancement in their current firm, which would make quits unnecessary in this context. Our results imply that better relative wage positions are not the often cited factor to reduce quits, because they have the counter acting effect of signaling workers few further career advancement opportunities.

As quits are driven to some extent by utility maximizing behavior, our results can also be incorporated into the broader literature about the determinants of subjective well-being. That the status effect is dominated by a counter acting signal effect has also been found in other recent studies about comparison income and satisfaction (Senik, 2008; Clark et al., 2009). One limitation of our study, which we have in common with previous studies, is that we cannot separately identify status and signal effects of relative wage positions. Future research should therefore emphasize the distinction between status and signal and try to separate their effects. Our paper is nevertheless important because it shows that some previous results on comparison income, which are mostly based on survey data and laboratory experiments, are also found in real world data about important decisions in peoples' life, which do not suffer from a subjectivity bias and from the critique of unrealistic laboratory environments (Falk and Heckman, 2009).

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Appendix

Table A.1: Mobility events and annual observations

	Sample Means			
Annual observations per establishment	N ≥ 10	N ≥ 15	N ≥ 50	N ≥ 100
Number of mobility events	7,785	7,672	7,037	6,323
Number of observations	3,867,569	3,843,919	3,639,961	3,382,528

Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table A.2: Descriptive statistics for the complete sample

Variable	Mean	Std. Deviation
Quit = 1	0.0020	0.0448
w_{ijt}^{abs}	4.6437	0.2414
$\overline{w}_{jt}$	4.6415	0.1585
$\hat{w}_{ijt}^{inside}$	4.6414	0.2048
w_{ijt}^{rank}	0.5002	0.2837
w_{ijt}^{range}	0.6309	0.2371
w_{ijt}^{CDF}	0.5127	0.2954
$\hat{w}_{it}^{outside}$	4.6378	0.1701
Tenure (years)	12.9425	7.7708
Tenure squared	227.8940	220.7135
Potential experience (years)	19.2799	6.1179
Potential experience squared	409.1431	225.5760
Professional status (dummies)		
Unskilled worker[i]	0.2788	0.4484
Skilled worker/ Craftsman[ii]	0.3423	0.4745
Technician/ Foreman[iii]	0.0333	0.1793
Clerk[iv]	0.3456	0.4756
Highest schooling degree (dummies)		
Secondary school leaving certificate[v]	0.1197	0.3246
Secondary school leaving certificate and apprenticeship[vi]	0.8183	0.3856
(Technical) college entrance qualification[vii]	0.0090	0.0944
(Technical) college entrance qualification and apprenticeship[viii]	0.0529	0.2239
Share of unqualified workers within the establishment	0.2700	0.2475
Establishment size class (dummies)		
Workforce of establishment in [10;49]	0.0122	0.1096
Workforce of establishment in [50;199]	0.0670	0.2500
Workforce of establishment in [200;999]	0.2634	0.4405
Workforce of establishment in [> 1000]	0.6575	0.4746
Works council	0.9667	0.1795
Collective bargaining	0.9528	0.2120
Sector (dummies)		
Agriculture	0.0009	0.0297
Mining	0.0474	0.2126
Building	0.0144	0.1192
Credit	0.0559	0.2297
Traffic	0.0498	0.2176
Retail	0.0297	0.1698
Hotel	0.0055	0.0742
Education	0.0118	0.1079
Service	0.0308	0.1728
Welfare	0.0285	0.1665

Public utility	0.0598	0.2371
Production	0.6654	0.4719
Federal region (dummies)		
Schleswig Holstein	0.0277	0.1641
Hamburg	0.0623	0.2418
Lower Saxony[ix]	0.1416	0.3486
Bremen	0.0221	0.1470
North Rhine-Westphalia[x]	0.2672	0.4425
Hesse (Hessen)	0.0870	0.2819
Baden-Wuerttemberg	0.1339	0.3405
Bavaria[xi]	0.1579	0.3647
Rhineland-Palatinate and Saarland[xii]	0.1003	0.3004
Time-fixed effects (dummies)		
Year: 1996	0.1102	0.3131
Year: 1997	0.0926	0.2899
Year: 1998	0.0883	0.2838
Year: 1999	0.0849	0.2787
Year: 2000	0.0987	0.2982
Year: 2001	0.1069	0.3090
Year: 2002	0.1124	0.3158
Year: 2003	0.0956	0.2940
Year: 2004	0.1091	0.3118
Year: 2005	0.1013	0.3018
Number of observations	3,867,569	
Number of individuals	1,115,437	
Number of establishments	6,791	

Note: German terms:
i) nicht formal qualifiziert
ii) Facharbeiter
iii) Meister, Poliere
iv) Angestellter
v) bis mittlere Reife ohne Berufsausbildung
vi) bis mittlere Reife mit Berufsausbildung
vii) bis (Fach-)Hochschulreife ohne Berufsausbildung
viii) bis (Fach-)Hochschulreife mit Berufsausbildung
ix) Niedersachsen
x) Nordrhein Westfalen
xi) Bayern
xii) Rheinland Pfalz, Saarland
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table A.3: Random-effects Probit results for quit probability

Quit=1	(1)	(2)	(3)	(4)	(5)	(6)	(7)
w_{ijt}^{abs}	0.1623	0.2280	0.2874	0.0645	0.1914	0.0129	0.1612
	(0.0241)	(0.0279)	(0.0318)	(0.0313)	(0.0369)	(0.0322)	(0.0242)
$\bar{w}_{jt}$		-0.1921					
		(0.0405)					
$\hat{w}_{ijt}^{inside}$			-0.2879				
			(0.0472)				
w_{ijt}^{rank}				0.1107			
				(0.0228)			
w_{ijt}^{range}					-0.0349		
					(0.0334)		
w_{ijt}^{CDF}						0.1571	
						(0.0227)	
$\hat{w}_{it}^{outside}$							0.0782
							(0.1804)
Tenure	-0.0417	-0.0419	-0.0421	-0.0421	-0.0417	-0.0421	-0.0417
	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
Tenure squared	0.0009	0.0009	0.0009	0.0009	0.0009	0.0009	0.0009
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Potential experience	0.0076	0.0074	0.0103	0.0073	0.0077	0.0071	0.0062
	(0.0037)	(0.0037)	(0.0037)	(0.0037)	(0.0037)	(0.0037)	(0.0050)
Potential experience squared	-0.0004	-0.0004	-0.0004	-0.0004	-0.0004	-0.0004	-0.0004
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Secondary school leaving	0.1482	0.1507	0.1617	0.1507	0.1482	0.1527	0.1418
certificate & apprenticeship	(0.0177)	(0.0177)	(0.0178)	(0.0177)	(0.0177)	(0.0177)	(0.0229)
(Technical) college entrance	0.1147	0.1215	0.1362	0.1215	0.1149	0.1190	0.1035
qualification	(0.0445)	(0.0445)	(0.0446)	(0.0445)	(0.0445)	(0.0445)	(0.0514)
(Technical) college entrance	0.2929	0.2970	0.3175	0.2973	0.2929	0.2976	0.2807
qualification & apprenticeship	(0.0237)	(0.0237)	(0.0241)	(0.0237)	(0.0237)	(0.0237)	(0.0367)
Skilled worker/ Craftsman	0.0105	0.0084	0.0217	0.0064	0.0104	0.0057	0.0049
	(0.0127)	(0.0128)	(0.0129)	(0.0128)	(0.0127)	(0.0128)	(0.0182)
Technician/ Foreman	-0.1036	-0.1166	-0.0516	-0.1191	-0.1027	-0.1237	-0.1283
	(0.0305)	(0.0306)	(0.0317)	(0.0307)	(0.0305)	(0.0306)	(0.0647)
Clerk	0.0536	0.0467	0.0999	0.0459	0.0541	0.0419	0.0311
	(0.0147)	(0.0148)	(0.0166)	(0.0147)	(0.0147)	(0.0148)	(0.0540)
Year: 1997	0.0591	0.0598	0.0610	0.0595	0.0588	0.0605	0.0583
	(0.0220)	(0.0220)	(0.0220)	(0.0220)	(0.0220)	(0.0220)	(0.0221)
Year: 1998	0.0656	0.0689	0.0707	0.0682	0.0650	0.0702	0.0637
	(0.0222)	(0.0222)	(0.0222)	(0.0222)	(0.0222)	(0.0222)	(0.0226)
Year: 1999	0.2463	0.2532	0.2544	0.2518	0.2447	0.2552	0.2427
	(0.0208)	(0.0208)	(0.0208)	(0.0208)	(0.0208)	(0.0208)	(0.0224)
Year: 2000	0.3131	0.3211	0.3236	0.3195	0.3117	0.3230	0.3093
	(0.0193)	(0.0194)	(0.0194)	(0.0194)	(0.0193)	(0.0194)	(0.0211)

Year: 2001	0.2396	0.2503	0.2527	0.2482	0.2384	0.2527	0.2345
	(0.0197)	(0.0199)	(0.0199)	(0.0198)	(0.0198)	(0.0198)	(0.0229)
Year: 2002	0.0339	0.0471	0.0494	0.0447	0.0318	0.0501	0.0274
	(0.0215)	(0.0217)	(0.0216)	(0.0216)	(0.0216)	(0.0216)	(0.0262)
Year: 2003	0.0673	0.0851	0.0888	0.0814	0.0624	0.0902	0.0582
	(0.0222)	(0.0225)	(0.0225)	(0.0224)	(0.0227)	(0.0224)	(0.0305)
Year: 2004	0.2348	0.2544	0.2578	0.2505	0.2304	0.2599	0.2241
	(0.0201)	(0.0205)	(0.0205)	(0.0204)	(0.0205)	(0.0204)	(0.0317)
Year: 2005	0.2140	0.2355	0.2390	0.2310	0.2090	0.2417	0.2029
	(0.0206)	(0.0212)	(0.0211)	(0.0210)	(0.0212)	(0.0210)	(0.0330)
Workforce [50;199]	0.0470	0.0490	0.0480	0.0485	0.0477	0.0528	0.0469
	(0.0348)	(0.0348)	(0.0348)	(0.0348)	(0.0348)	(0.0349)	(0.0348)
Workforce [200;999]	-0.0045	0.0013	0.0017	0.0011	-0.0028	0.0071	-0.0048
	(0.0341)	(0.0342)	(0.0342)	(0.0342)	(0.0341)	(0.0342)	(0.0341)
Workforce [>1000]	-0.0491	-0.0354	-0.0331	-0.0360	-0.0458	-0.0287	-0.0495
	(0.0344)	(0.0346)	(0.0346)	(0.0346)	(0.0346)	(0.0346)	(0.0344)
Schleswig Holstein	-0.1483	-0.1513	-0.1513	-0.1507	-0.1483	-0.1523	-0.1465
	(0.0322)	(0.0322)	(0.0322)	(0.0322)	(0.0322)	(0.0322)	(0.0325)
Hamburg	-0.0611	-0.0538	-0.0553	-0.0549	-0.0609	-0.0541	-0.0666
	(0.0223)	(0.0223)	(0.0223)	(0.0223)	(0.0223)	(0.0223)	(0.0256)
Lower Saxony	-0.0349	-0.0341	-0.0325	-0.0339	-0.0351	-0.0349	-0.0363
	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0186)
Bremen	0.4534	0.4551	0.4565	0.4545	0.4530	0.4547	0.4520
	(0.0228)	(0.0228)	(0.0228)	(0.0228)	(0.0228)	(0.0228)	(0.0230)
North Rhine-Westphalia	-0.1270	-0.1254	-0.1264	-0.1258	-0.1276	-0.1256	-0.1287
	(0.0169)	(0.0169)	(0.0169)	(0.0169)	(0.0169)	(0.0169)	(0.0173)
Hesse	-0.2241	-0.2197	-0.2221	-0.2203	-0.2238	-0.2211	-0.2254
	(0.0224)	(0.0224)	(0.0224)	(0.0224)	(0.0224)	(0.0224)	(0.0226)
Baden Wuerttemberg	0.2126	0.2243	0.2246	0.2222	0.2114	0.2249	0.2071
	(0.0168)	(0.0170)	(0.0169)	(0.0169)	(0.0168)	(0.0169)	(0.0210)
Bavaria	-0.1001	-0.0999	-0.1013	-0.0991	-0.0999	-0.1009	-0.0998
	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0183)	(0.0183)
Works council	-0.1685	-0.1531	-0.1527	-0.1566	-0.1705	-0.1510	-0.1685
	(0.0221)	(0.0224)	(0.0223)	(0.0222)	(0.0222)	(0.0223)	(0.0221)
Structure	-0.1504	-0.1713	-0.1615	-0.1680	-0.1491	-0.1732	-0.1505
	(0.0170)	(0.0171)	(0.0169)	(0.0170)	(0.0171)	(0.0168)	(0.0170)
Collective bargaining	0.1609	0.1600	0.1628	0.1599	0.1605	0.1610	0.1608
	(0.0214)	(0.0215)	(0.0215)	(0.0215)	(0.0214)	(0.0215)	(0.0214)
Agriculture	-0.3532	-0.3789	-0.3839	-0.3722	-0.3484	-0.3837	-0.3340
	(0.2018)	(0.2022)	(0.2022)	(0.2022)	(0.2017)	(0.2023)	(0.2066)
Mining	0.1313	0.1263	0.1203	0.1277	0.1311	0.1265	0.1361
	(0.0204)	(0.0205)	(0.0205)	(0.0205)	(0.0204)	(0.0205)	(0.0232)
Building	-0.0488	-0.0541	-0.0584	-0.0524	-0.0488	-0.0557	-0.0395
	(0.0333)	(0.0333)	(0.0333)	(0.0333)	(0.0333)	(0.0333)	(0.0395)
Credit	0.1313	0.1513	0.1264	0.1482	0.1305	0.1524	0.1319
	(0.0174)	(0.0180)	(0.0175)	(0.0178)	(0.0175)	(0.0177)	(0.0175)
Traffic	0.0735	0.0673	0.0625	0.0692	0.0745	0.0661	0.0786
	(0.0199)	(0.0200)	(0.0200)	(0.0200)	(0.0200)	(0.0200)	(0.0232)

Retail	0.0220	0.0176	0.0020	0.0203	0.0227	0.0165	0.0338
	(0.0236)	(0.0236)	(0.0239)	(0.0236)	(0.0236)	(0.0236)	(0.0360)
Hotel	0.0280	0.0160	0.0072	0.0197	0.0297	0.0146	0.0432
	(0.0543)	(0.0543)	(0.0544)	(0.0543)	(0.0543)	(0.0543)	(0.0645)
Education sector	-0.3200	-0.3299	-0.3478	-0.3266	-0.3185	-0.3307	-0.3072
	(0.0570)	(0.0572)	(0.0573)	(0.0572)	(0.0570)	(0.0572)	(0.0642)
Service	0.1278	0.1178	0.1070	0.1209	0.1301	0.1150	0.1399
	(0.0215)	(0.0217)	(0.0219)	(0.0216)	(0.0216)	(0.0217)	(0.0353)
Welfare	0.0222	0.0083	-0.0170	0.0110	0.0229	0.0072	0.0406
	(0.0257)	(0.0259)	(0.0265)	(0.0258)	(0.0258)	(0.0258)	(0.0498)
Public utility	0.1604	0.1449	0.1283	0.1493	0.1618	0.1434	0.1755
	(0.0180)	(0.0183)	(0.0188)	(0.0181)	(0.0180)	(0.0181)	(0.0391)
Constant	-3.9959	-3.4355	-3.3426	-3.6146	-4.1082	-3.4132	-4.3153
	(0.1134)	(0.1631)	(0.1549)	(0.1370)	(0.1566)	(0.1398)	(0.7455)
Log Likelihood	-53034.6	-53023.3	-53016.0	-53022.7	-53034.0	-53010.5	-53034.5
LR test of rho=0 (P-Value)	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001
Number of observations				3,867,569			
Number of individuals				1,115,437			

Note: Random-effects Probit (coefficients). Robust standard errors in parentheses. 'LR' denotes Likelihood-Ratio.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table A.4: Random-effects Probit results for quit probability for observations in firms with quits

Quit = 1	$\bar{x}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
w_{ijt}^{abs}	4.6877	0.0002155* (0.00011)	0.0006863*** (0.00014)	0.00119*** (0.00016)	-0.0006136*** (0.00015)	0.001154*** (0.00018)	-0.0006818*** (0.000903)	0.00024** (0.00011)
$\bar{w}_{jt}$	4.6868		-0.001458*** (0.00021)					
$\hat{w}_{ijt}^{inside}$	4.6858			-0.0024*** (0.00027)				
w_{ijt}^{rank}	0.4929				0.0009153*** (0.00012)			
w_{ijt}^{range}	0.6819					-0.001201*** (0.00017)		
w_{ijt}^{CDF}	0.5100						0.000903*** (0.00012)	
$\hat{w}_{it}^{outside}$	4.6500							-0.0019816** (0.00092)
Control variables		yes	yes	yes	yes	yes	yes	yes
$\Pr(y=1\mid\bar{x})$		0.00121433	0.00120318	0.00120544	0.00120382	0.00120752	0.00120174	0.00121523
Number of obs.					1,339,892			
Number of indiv.					569,639			

Note: Sample includes only observations of firms with at least one quitting worker in a year. Random-effects Probit (marginal effects at $\bar{x}$).
Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.10.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table A.5: Descriptive statistics for mobile individuals

Variable	Mean	Std. Deviation
$\left(w^{new} - w^{old}\right)_{it}$	0.0336	0.1552
Wage Cut = 1	0.2847	0.4513
$\left(rank^{new} - rank^{old}\right)_{it}$	-0.0312	0.2573
$\left(range^{new} - range^{old}\right)_{it}$	0.0034	0.2045
$\left(CDF^{new} - CDF^{old}\right)_{it}$	-0.0325	0.2536
Secondary school leaving certificate	0.0641	0.2449
Secondary school leaving certificate and apprenticeship	0.8077	0.3941
(Technical) college entrance qualification	0.0121	0.1092
(Technical) college entrance qualification and apprenticeship	0.1161	0.3204
Mobility to larger establishment	0.5603	0.4964
Blue-collar to white-collar transition	0.0434	0.2038
Potential experience	17.9665	6.3326
Potential experience squared	362.8906	227.8344
Year: 1996	0.0674	0.2508
Year: 1997	0.0673	0.2506
Year: 1998	0.0645	0.2456
Year: 1999	0.0975	0.2966
Year: 2000	0.1563	0.3632
Year: 2001	0.1344	0.3411
Year: 2002	0.0796	0.2708
Year: 2003	0.0722	0.2588
Year: 2004	0.1382	0.3451
Year: 2005	0.1225	0.3279
Number of observations	7,785	
Number of individuals	7,516	

Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table A.6: Probit results for acceptance of wage cuts for job switches between large establishments

Wage Cut = 1	$\bar{x}$	(1)	(2)	(3)
$\left(rank^{new} - rank^{old}\right)_{it}$	-0.0193	-0.8861***		
		(0.0556)		
$\left(range^{new} - range^{old}\right)_{it}$	0.0062		-1.0497***	
			(0.0681)	
$\left(CDF^{new} - CDF^{old}\right)_{it}$	-0.0201			-0.9389***
				(0.0570)
Secondary school leaving certificate	0.0894	reference	reference	reference
Secondary school leaving certificate and apprenticeship	0.7973	-0.2094***	-0.1538***	-0.2126***
		(0.0389)	(0.0399)	(0.0391)
(Technical) college entrance qualification	0.0110	-0.1301*	-0.0887	-0.1127*
		(0.04818)	(0.0700)	0.0492
(Technical) college entrance qualification and apprenticeship	0.1023	-0.1953***	-0.1730***	-0.1882***
		(0.0193)	(0.0255)	(0.0196)
Potential experience	17.8107	0.0009	-0.0038	0.0063
		(0.0096)	(0.0100)	(0.0099)
Potential experience squared	349.2447	-0.0002	-0.0001	-0.0003
		(0.0003)	(0.0003)	(0.0003)
Annual dummies		yes	yes	yes
Pseudo R^2		0.2053	0.1463	0.2252
$\Pr(y=1 \mid \bar{x})$		0.2082	0.2240	0.2029
Number of Observations			2,092	

Note: Sample includes only observations who change firms within the highest establishment size class (more than 1000 employees). Probit (marginal effects at $\bar{x}$). Robust standard errors clustered for 2,022 individuals in parentheses. *** p<0.01, ** p<0.05, * p<0.10.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Tables and figures included in text

Table 1: Recent studies about job mobility and wages

Authors	Country (data set)	Mobility (in percent) ... to lower wage	to same wage	to higher wage
Fitzenberger and Garloff (2007)	Germany (IABS)	22.2 - 24.5	3.8 - 7.1	70.7 - 72.7
Jolivet et al. (2006)	Germany (ECHP)	36.3	3.3	60.4
	U.S. (PSID)	23.3	21.1	55.6
Nosal and Rupert (2007)	U.S. (PSID)	42.1 - 42.4	8.4 - 4.8	49.5- 52.8

Note: Fitzenberger and Garloff (2007) refer to establishment-to-establishment transitions between two successive years. The authors use different subsamples for their analysis on wage cuts which do not differ much. Nosal and Rupert (2007) consider individuals who report an employer change. Jolivet et al. (2006) define mobility as job-to-job transition if the interval between jobs was one month or less (Germany) or less than three weeks (U.S.).
'IABS': IAB employment subsample 1975-2001.
'ECHP': European Community Household Panel Survey.
'PSID': Panel Study of Income Dynamics.

Table 2: Definition of wage measures

w_{ijt}^{abs}	Log wage of individual *i* in period *t* in establishment *j*	$w_{ijt}^{abs} = ln(Wage_{ijt})$
$\overline{w}_{jt}$	Average log wage paid in establishment *j* in period *t*	$\frac{1}{N_{jt}} \sum_{i_{jt}=1}^{N_{jt}} w_{ijt}$ *N*: number of employees in our sample
$\hat{w}_{ijt}^{inside}$	Predicted comparison wage in own establishment (given individual characteristics) in period *t*	Annual regression for establishment *j*: $\hat{w}_{ijt}^{inside} = \hat{\alpha}_j + \hat{\gamma}' X_{it}$ *X*: potential experience (squared), dummies for occupation, and schooling
w_{ijt}^{rank}	Ordinal relative wage position of individual *i* in establishment *j* in period *t*	$w_{ijt}^{rank} = \frac{\text{wage rank}_{\text{ijt}} - 1}{\text{wage rank}_{\text{jt}}^{\max} - 1}$ Workers with equal wages within establishment j have the same rank. In such cases we calculate the average rank of workers with same wages. For example, if the two lowest paid workers are paid the same, both employees exhibit a non-normalized wage rank of 1.5.
w_{ijt}^{range}	Cardinal relative wage position of individual *i* in establishment *j* in period *t*	$w_{ijt}^{range} = \frac{(w_{ijt} - w_{jt}^{\min})}{(w_{jt}^{\max} - w_{jt}^{\min})}$
w_{ijt}^{CDF}	Empirical cumulative distribution function (CDF) of w_{ijt} in establishment *j* in period *t*	$w_{ijt}^{CDF} = \Pr(W_{jt} \le w_{ijt})$ W_{jt} is the set of wages within establishment *j* in period *t*. w_{ijt} denotes the individual wage of individual *i* working in establishment *j* in period *t*.
$\hat{w}_{it}^{outside}$	Predicted comparison wage across all individuals in all firms (given individual characteristics) in period *t*	Annual regression: $\hat{w}_{it}^{outside} = \hat{\alpha} + \hat{\gamma}' X_{it} + \hat{\delta}_1 S_{it} + \hat{\gamma}_2 A_{it}$ *X*: potential experience (squared), dummies for occupation and schooling; *A*: regional dummies; *S*: sector.

Table 3: Descriptive statistics and correlations of wage measures

Variable	Mean ($\bar{x}$) (Std.dev.)	Correlation coefficients w_{ijt}^{abs}	$\overline{w}_{jt}$	$\hat{w}_{ijt}^{inside}$	w_{ijt}^{rank}	w_{ijt}^{range}	w_{ijt}^{CDF}	$\hat{w}_{it}^{outside}$
w_{ijt}^{abs}	4.6437 (0.2414)	1						
$\overline{w}_{jt}$	4.6415 (0.1585)	0.6476***	1					
$\hat{w}_{ijt}^{inside}$	4.6414 (0.2048)	0.8403***	0.7687***	1				
w_{ijt}^{rank}	0.5002 (0.2837)	0.6750***	-0.0407***	0.3689***	1			
w_{ijt}^{range}	0.6309 (0.2371)	0.8116***	0.4033***	0.6353***	0.6521***	1		
w_{ijt}^{CDF}	0.5127 (0.2954)	0.7064***	0.0151***	0.4099***	0.9721***	0.6775***	1	
$\hat{w}_{it}^{outside}$	4.6378 (0.1701)	0.6838***	0.4971***	0.8133***	0.4277***	0.5178***	0.4552***	1
Number of observations				3,867,569				

Note: *** p<0.01, ** p<0.05, * p<0.10.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table 4: Random-effects Probit results for quit probability

Quit = 1	$\bar{x}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
w_{ijt}^{abs}	4.6437	0.0002066*** (0.00003)	0.0002881*** (0.00004)	0.0003624*** (0.00005)	0.0000814** (0.00004)	0.0002437*** (0.00005)	0.0000163 (0.00004)	0.0002052*** (0.00003)
$\bar{w}_{jt}$	4.6415		-0.0002428*** (0.00005)					
$\hat{w}_{ijt}^{inside}$	4.6414			-0.000363*** (0.00006)				
w_{ijt}^{rank}	0.5002				0.0001398*** (0.00003)			
w_{ijt}^{range}	0.6309					-0.0000444 (0.00004)		
w_{ijt}^{CDF}	0.5127						0.0001979*** (0.00003)	
$\hat{w}_{it}^{outside}$	4.6378							0.0000995 (0.00023)
Control variables (see Table A.2)		yes	yes	yes	yes	yes	yes	yes
$\Pr(y=1 \mid \bar{x})$		0.00034897	0.00034621	0.00034541	0.00034595	0.00034904	0.00034493	0.00034886
Number of obs.					3,867,569			
Number of individuals					1,115,437			

Note: Random-effects Probit (marginal effects at $\bar{x}$). Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.10. The corresponding coefficients of the Probit estimates and the complete results are presented in Table A.3. Table A.2 contains descriptive statistics.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table 5: Descriptive statistics for consequences of quits

		Mean	Std. Deviation
$\left(w^{new} - w^{old}\right)_{it}$	all	0.0336	0.1552
	wage cut	-0.1171	0.1419
	wage markup	0.0936	0.1140
$\left(rank^{new} - rank^{old}\right)_{it}$	all	-0.0312	0.2573
	wage cut	-0.1747	0.2682
	wage markup	0.0260	0.2291
$\left(range^{new} - range^{old}\right)_{it}$	all	0.0034	0.2045
	wage cut	-0.1110	0.2083
	wage markup	0.0489	0.1841
$\left(CDF^{new} - CDF^{old}\right)_{it}$	all	-0.0325	0.2536
	wage cut	-0.1861	0.2538
	wage markup	0.0286	0.2263

Note: 7,785 transitions of 7,516 individuals are observed. 5,569 moves are executed to higher wages and 2,216 moves to lower wages (wage cut).
Source: Cross-sectional model of the LIAB (Years 1996-2005).

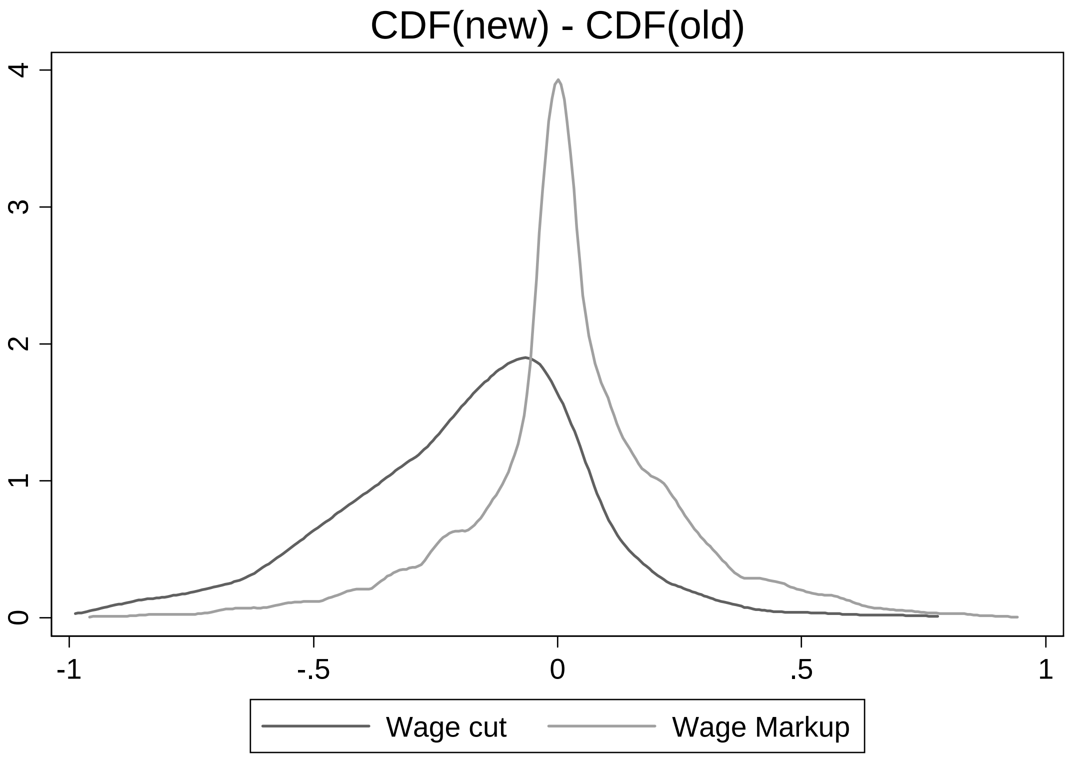

Figure 1: Kernel density estimator for changes in relative wage positions

Table 6: Linear regression results for wage change

$\left(w^{new} - w^{old}\right)_{it}$	(1)	(2)	(3)
$\left(rank^{new} - rank^{old}\right)_{it}$	0.2377***		
	(0.0108)		
$\left(range^{new} - range^{old}\right)_{it}$		0.3443***	
		(0.0136)	
$\left(CDF^{new} - CDF^{old}\right)_{it}$			0.2586***
			(0.0105)
Secondary school leaving certificate	reference	reference	reference
Secondary school leaving certificate and apprenticeship	0.0122*	0.0122*	0.0121*
	(0.0066)	(0.0068)	(0.0066)
(Technical) college entrance qualification	0.0255*	0.0181	0.0234*
	(0.0139)	(0.0122)	(0.0139)
(Technical) college entrance qualification and apprenticeship	0.0123	0.0105	0.0092
	(0.0077)	(0.0078)	(0.0077)
Mobility to larger establishment	0.0465***	0.0165***	0.0457***
	(0.0043)	(0.0041)	(0.0043)
Potential experience	-0.0044***	-0.0032**	-0.0045***
	(0.0016)	(0.0016)	(0.0016)
Potential experience squared	0.0001	0.0000	0.0001
	(0.0000)	(0.0000)	(0.0000)
Constant	0.0570***	0.0408***	0.0605***
	(0.0161)	(0.0161)	(0.0160)
Annual dummies	yes	yes	yes
R^2	0.1896	0.2335	0.2125
Number of observations		7,785	

Note: OLS (coefficients). Robust standard errors clustered for 7,516 individuals in parentheses. *** p<0.01, ** p<0.05, * p<0.10. Table A.5 contains descriptive statistics.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

Table 7: Probit results for acceptance of wage cuts

Wage Cut = 1	(1)	(2)	(3)
$\left(rank^{new} - rank^{old}\right)_{it}$	-0.6534***		
	(0.0252)		
$\left(range^{new} - range^{old}\right)_{it}$		-0.8496***	
		(0.0304)	
$\left(CDF^{new} - CDF^{old}\right)_{it}$			-0.7315***
			(0.0255)
Secondary school leaving certificate	reference	reference	reference
Secondary school leaving certificate and apprenticeship	-0.0803***	-0.0883***	-0.0824***
	(0.0216)	(0.0229)	(0.0217)
(Technical) college entrance qualification	-0.1459***	-0.1368***	-0.1387***
	(0.0337)	(0.0343)	(0.0343)
(Technical) college entrance qualification and apprenticeship	-0.1659***	-0.1681***	-0.1606***
	(0.0170)	(0.0172)	(0.0171)
Mobility to larger establishment	-0.0746***	-0.0022	-0.0737***
	(0.0115)	(0.0126)	(0.0116)
Potential experience	0.0022	0.0021	0.0037
	(0.0047)	(0.0046)	(0.0047)
Potential experience squared	-0.0001	-0.0002	-0.0002
	(0.0001)	(0.0001)	(0.0001)
Annual dummies	yes	yes	yes
Pseudo R^2	0.1283	0.1297	0.1480
$\Pr(y = 1 \mid \bar{x})$	0.2591	0.2584	0.2543
Number of observations		7,785	

Note: Probit (marginal effects at $\bar{x}$). Robust standard errors clustered for 7,516 individuals in parentheses. *** p<0.01, ** p<0.05, * p<0.10. Table A.5 contains descriptive statistics.
Source: Cross-sectional model of the LIAB (Years 1996-2005).

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