Labor Market Mobility in Germany

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Abstract

This thesis contributes to the recent discussion about the flexibility of the German labor market. The empirical studies analyze individual mobility between jobs using German data. Specifically, chapters 2 and 3 rely on integrated employer-employee data, while in chapter 4 household data is applied. After a brief introduction about the relevance of labor market mobility in Germany (chapter 1), chapter 2 focuses on monetary consequences of individual between-establishment transitions. Counterfactual wage trajectories are estimated in order to compare the wage trajectories at different employers simultaneously. The main finding is that only few immediate wage cuts pay off because of steeper wage growth in the new job. Chapter 3 enhances the literature by an examination of the relationship between quit decisions and the relative wage position within an establishment. The main assumptions are that individuals compare themselves to colleagues within the same establishment and that workers form rational expectations about where they lie in the pay ordering. Voluntary mobility with wage cuts is analyzed in chapter 4, in which the effects of different subjective comparisons between the previous and the current job on the decision to accept earnings losses are investigated.

Keywords: Mobility, wage cut, relative wage position.

Kurzzusammenfassung

Diese Dissertation beschäftigt sich mit der Flexibilität des deutschen Arbeitsmarktes, wobei die empirischen Studien individuelle Arbeitsplatzwechsel anhand deutscher Daten untersuchen. Dabei beruhen Kapitel 2 und 3 auf integrierten Betriebs- und Personendaten, während in Kapitel 4 Haushaltspaneldaten herangezogen werden. Nach einer kurzen Übersicht zur Bedeutung der individuellen Mobilität (Kapitel 1) folgen die empirischen Analysen. Kapitel 2 beschäftigt sich mit dem Vergleich von kontrafaktischen Lohnkurven, um den Lohn eines mobilen Arbeitnehmers im Ausgangs- und Zielbetrieb vergleichen zu können. Der Einfluss der relativen Lohnposition auf die Entscheidung den Betrieb zu wechseln ist Gegenstand von Kapitel 3. Dabei wird zugrunde gelegt, dass Individuen ihre relative Lohnposition anhand von Lohnvergleichen mit den Kollegen innerhalb ihres Betriebes abschätzen. Schließlich behandelt Kapitel 4 die Frage warum Personen bei einem Wechsel des Arbeitgebers Lohnabschläge akzeptieren. Die Studie basiert dabei vor allem auf subjektiven Vergleichen zwischen dem aktuellen und dem vorherigen Arbeitsplatz.

Schlagwörter: Mobilität, Arbeitsplatzwechsel, Lohnabschlag, relative Lohnposition.

Contents

1 Introduction

The German labor market is characterized by a significant degree of dynamism, leading to a continuous creation and destruction of jobs. As a result, employees change their labor market status, moving from one job to another, from employment to non-employment, from employment to unemployment, from unemployment to non-employment, and vice versa. Labor market mobility, on the one hand, might increase efficiency and productivity because of reallocation of resources where they are most productive. On the other hand, this progress pushes responsibility for careers as well as uncertainty about income security onto workers. Peter Capelli (1999, p. 17) describes this with the following words: *"THE OLD employment system of secure, lifetime jobs with predictable advancement and stable pay is dead."* This thesis analyzes job-to-job mobility in Germany and, thus, contributes to the aspect of the worker's responsibility for the own career. In fact, todays employees are characterized by a large degree of self-determination and flexibility. This also includes the improvement of the own career by finding a new job and quitting the previous one. In one of his speeches, Earl Nightingale (1921–1989, American author) implicitly motivates workers to quit their jobs to a new one for climbing up the career ladder because *"Jobs are owned by the company, you own your career"*.

The seminal literature on the on-the-job search introduced the possibility that workers search for new jobs while employed. These studies, then, intended to explain quit rates and individual quit behavior. This early literature, however, mainly focused on wage maximization problems in voluntary mobility. This implies that workers form their decisions to change jobs only by comparison of wages which can be obtained in different firms. The next chapter of this thesis empirically contributes to this literature and examines whether mobility to a new job pays off in the long run. This work differs from other studies because it applies an innovative methodology based on firm-specific estimation of counterfactual wage trajectories. More recent on-the-job search literature change focus from wage-maximization to utility-maximization where bundles of various job characteristics (including wages) affect the decision to quit a job voluntarily. Chapter 3 enhances this literature via an empirical examination of the relationship between quit decisions and the relative wage position within a firm. This is an important determinant of the own career because it might signal future career prospects and could be interpreted as reputation or status within a firm. The analysis explicitly controls for the monetary component, in order to account for wage effects in labor market quit decisions. Another aspect of this chapter is whether workers are mobile to lower wages if they can improve their relative wage position in the new firm. Individual quits with earnings losses are further analyzed in chapter 4, in which the effects of different job-specific (non-wage) characteristics on the decision to accept wage cuts are examined. Therefore, this study contributes to the so far sparse literature about the reasons for this behavior. It is important to note that empirical analyzes always have some drawbacks. The most serious problem in this thesis concerns the definition of voluntary job-to-job mobility. The data, in fact, do not allow for distinct definitions regarding voluntary quits in chapters 2 and 3. The studies, however, refer to measures which are arbitrary to those used by other researchers and provide robustness checks with respect to different definitions of voluntary quits to a new job. In the following part of this introduction, the main findings and conclusions are summarized.

Chapter 2 is to conclude on whether staying at the same employer or moving to a different employer pays off in the medium term. This study is the first which estimates individual counterfactual wage profiles in order to compare wages at different employers simultaneously. This allows conclusions about the monetary consequences of individual mobility. The results show that most mobile workers achieve permanently higher wages when being mobile once. A substantial share of mobile workers, in turn, suffers wage cuts when changing the establishment. A major finding of this study is that less than one in three immediate wage cuts is beneficial in the future. One explanation is that job-specific amenities play a role for transitions to permanently lower wages. An important topic for further research is comparison of the basic findings achieved with German linked employer-employee data with those obtained in other countries. A further promising field for future research is the inclusion of job-related amenities to conclude about the individual trade-off reasoning between job-specific amenities and wages.

Chapter 3 deals with the effect of individual relative wage positions which can be interpreted as a worker-specific amenity. The main point of interest is the effect of the relative wage position on the decision to voluntarily quit a job. The results show that relative wage positions have a significant impact on the probability to voluntarily quit a job. In addition, the study concentrates on both, linear as well as nonlinear effects. When considering linear effects, the analysis suggests that workers with high relative wage positions within their firms are more likely to quit a job in comparison to workers with low relative wage positions. This finding is consistent with the results of a study conducted with Italian data. The results can also be incorporated into the literature about interdependent preferences and the determinants of subjective well-being because the position in the pay ordering within a firm also has significant impact on job satisfaction, as found in a recent study conducted in Denmark. This thesis, however, goes beyond the scope of existing studies and introduces considerations about a nonlinear relationship as well as individual trade-off reasoning when changing jobs. The analysis of nonlinear effects of relative wage positions in quit decisions reveals a U-shape between the wage rank and the decision to quit. An explanation is that workers at the bottom of the within-establishment pay scale are more sensitive to status considerations and those at the top to signal considerations. In other words, workers in low relative wage positions quit their jobs because of their low status while workers in high relative wage positions quit because of low future career prospects. Finally, this chapter reveals that relative wage positions are significantly correlated with the probability to be mobile to lower wages. Workers who improve their relative wage position compared to the previous establishment are, on average, less likely to pay for mobility by lower wages. Trade-off reasoning between wages and relative wage positions, thus, is not evident in this study.

The last part of this thesis concentrates on mobility to lower wages. Studies show that more than one in three individuals pay for a new job by lower wages in the U.S., Germany, France, and Denmark. Literature, however, lacks detailed information on the reasons for mobility to lower wages. Chapter 4 sheds light on this behavioral pattern and examines the relationship between a variety of different subjective improvements in a diversity of job-specific amenities between two jobs and the willingness to pay for them. The results suggest that individual trade-off reasoning is evident. Specifically, individuals are found to pay for improvements in workload by lower wages. As chapter 2 shows that transitions to permanently lower wages are common, it might be hypothesized that workers trade off permanently lower wages with subjective improvements in job-specific characteristics (e.g., improvements in workload). The results also indicate compensating wage differentials for job-specific disamenities which are, however, to the largest extent, statistically insignificant.

In sum, this thesis shows that labor market mobility is very complex in nature. Combination of all the different aspects discussed in this thesis is a promising field of research. Nevertheless, it is problematic to find linked employer-employee data which combine long time horizons, job-specific amenities, wage information, and detailed information on the reason of mobility that allow for a joint investigation of all of the points addressed in this thesis.

2 Inter-Firm Labor Mobility and Wages

Individuals who want to leave their employer usually raise the question whether mobility to a new employer pays off in the future. This paper contributes to this question by examining the consequences of labor market mobility in the medium-term. Conclusions regarding whether an individual's wage trajectory at the new employer exceeds the one of the previous employer are drawn by application of an innovative procedure which involves the simultaneous investigation of wage trajectories at different employers. The main finding is that a considerable number of workers experience wage cuts which are of permanent nature. Only few of the transitions to lower wages pay off because of steeper wage growth in the new job.[1]

[1]This chapter was originally published as "Inter-Firm Labor Mobility and Wages", *Jahrbuch für Wirtschaftswissenschaften (Review of Economics)*, Vol. 61, 196–211. Publication within this thesis with kind permission of Vandenhoeck & Ruprecht GmbH & Co. KG. This study uses the Cross-sectional model of the Linked-Employer-Employee Data (LIAB; years 1993-2006) from the Institute for Employment Research (IAB). Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the IAB and remote data access.

2.1 Introduction

Labor market mobility is an outstanding characteristic of labor markets (Burgess et al. 2000, OECD 1997). This is also true for the German private sector, where individuals are unlikely to stay in one job over their entire working life. In fact, Winkelmann (1994) shows that male German workers hold an average of four lifetime jobs. Jolivet et al. (2006), Nosal and Rupert (2007), and Fitzenberger and Garloff (2007) show that wage markups and wage cuts coexist with mobility. It is further illustrated that a small proportion of workers is mobile and experiences no wage changes[2]. This paper is motivated by these results and contributes to the investigation of medium-term consequences of immediate wage cuts and wage markups.

Literature on the immediate consequences of job-to-job mobility is wide spread and states that wage markups as well as wage cuts coexist. Borjas (1981) emphasizes that an individual's earnings profile is discontinuous across jobs because job mobility on average results in a higher wage. Upward mobility is empirically confirmed by other studies (e.g., Topel and Ward 1992). Jolivet et al. (2006) show that 60.4% of German job-to-job transitions are to higher wages. Downward mobility is also shown to be frequent. In Denmark, France, and Germany, more than one in three mobile individuals change jobs at the price of a wage cut (Jolivet et al. 2006). In the U.S., Nosal and Rupert (2007) show that 42% of individuals voluntarily changing jobs suffer wage cuts.

However, individuals who are confronted with thoughts of leaving the employer usually evaluate their alternatives and raise the question whether mobility to a new employer pays off in the future.[3] Individuals, thus, are interested in the medium- or long-term consequences of their decision rather than in an evaluation of immediate success. In fact, it is hard to assess the consequences of mobility because it is problematic to conclude whether the individual within-firm wage path exceeds the between-firm (i.e., mobility) wage path. The main problem is, obviously, that individual wages of the main job are not

[2]Nosal and Rupert (2007) show that about 8% of all workers and approximately 5% of voluntarily mobile workers change jobs to the same wage.

[3]In the data used here, establishments are observed. From now on, firm, employer, and establishment are used interchangeably.

simultaneously observable across firms. The availability of matched employer-employee data and application of the firm-specific estimation approach of Abowd et al. (2006) help to solve this problem. The innovative procedure applied here involves estimation of wage equations for each single firm. In a next step, the results are applied to predict average wage paths at different employers which are used for the simultaneous analysis of individual wage trajectories at different firms. This procedure allows to address the following questions:

1. *In comparison to the old employer, do mobile workers achieve permanently higher wages at the new employer?*

2. *Do workers change to lower wages?*

3. *How many immediate wage cuts pay off in the future?*

This paper, thus, contributes to the immediate wage change of mobility and assesses the long-run consequences. A special focus is on the acceptance of wage cuts in the period of mobility. Connolly and Gottschalk (2008) and Postel-Vinay and Robin (2002) argue that wage cuts are accepted by mobile individuals because of greater wage growth in the new job. Downward mobility, therefore, can be justified as an investment in future wage growth. The results, however, show that workers frequently change to permanently lower wages while only a small share of workers is shown to pay for steeper wage growth at the new employer by wage cuts. This might be explained by differences in non-wage characteristics between the old and the new job which also might enter mobility decisions (Nosal and Rupert 2007).

The paper is structured as follows. Section 2.2 explains the data and the empirical procedure. Section 2.3 presents the results and section 2.4 the conclusions.

2.2 Data and procedure

2.2.1 Data

This analysis utilizes the linked employer-employee data set from the Institute for Employment Research[4] (Alda et al. 2005). This data set includes data from a representative annual establishment survey (the IAB Establishment Panel) and process-produced person-specific data from the IAB. The underlying data set (called LIAB) is an unbalanced panel of cross-sections from 1993 to 2006 at the corresponding record date of June 30^{th}. Hence, 14 periods are available to investigate the working careers of individuals in an unbalanced panel design. In order to obtain efficient estimates, the analysis includes only establishments with at least 50 observations. The main advantage of this data set is that counterfactual wage trajectories can be separately constructed for each firm.

The sample is restricted to German citizens working in full-time jobs in West Germany because the data lacks information on individual working hours. The analysis focuses on males aged between 18 and 60 years who are subject to social insurance contributions. Of foremost interest is the average individual daily wage[5] achieved in the primary occupation surveyed in the data. The consumer price index surveyed by the Statistisches Bundesamt Deutschland is applied to deflate the nominal wages (annual averages, with year 2005 = 100). Wages below the marginal employment ceiling of 400 Euro per month (or 13.33 Euro per day in 2005 wages) are excluded from the analysis.[6] Wages above the upper earnings limit for social security contributions are set to the corresponding ceiling. Wages, thus, are censored in the data. Experience is calculated with respect to the surveyed year of labor market entry.[7] Hence, possible unemployment spells after schooling are considered by construction of experience.

Essentially, the paper focuses on mobile employees who change from one LIAB estab-

[4] Institut für Arbeitsmarkt- und Berufsforschung (IAB).

[5] The data contain the average daily gross wages of individuals, which are calculated via information on "the total wage for the spell of notification" and "the duration of the spell of notification in days". Information on hours worked, thus, is not available.

[6] See Jacobebbinghaus (2008) for an explanation of such implausibly low wages.

[7] In particular, the earliest year of an individual's surveyed labor market entry is considered in order to obtain strictly increasing experience in the data. For East Germans, the surveyed labor market entry is left-censored before 1990, which is the main reason for exclusively focusing on workers in West Germany.

lishment to another LIAB establishment where the establishment identifier in the data is used to determine such transitions. A major concern regarding mobility is the distinction between involuntary and voluntary mobility. Jolivet et al. (2006) consider voluntary mobility as an unconstrained choice of the worker, which is difficult to ascertain using this data set.[8] Instead, I distinguish between direct and indirect transitions. Direct transitions are defined as changes between establishments whereas the individual must be full-time employed at another establishment eight days before entering the new establishment. Workers who are indirectly mobile change from one establishment to another one within a year.[9]

An advantage of this data set is that there are no problems regarding sorting or individual selection into LIAB establishments, because workers are not expected to systematically move from one LIAB establishment to another LIAB establishment. It should be noted that some types of moves cannot be accounted for; that is, transitions from a non-LIAB establishment into the sample or transitions out of the sample are not identified. Therefore, the number of transitions for an individual is unknown if the individual was not observed in the sample for the entire career horizon, as the individual's employment history is not completely captured in the data. Furthermore, this analysis refers to an employer-to-employer transition in which an individual changes establishments within the same employer.

After implementation of all restrictions, 11,340,952 observations on 4,697 firms and 2,622,048 individuals are subject to analysis. Table A2.1 in the appendix displays the descriptive statistics for the entire sample.

2.2.2 Procedure

The analysis focuses on the firm-specific estimation of individual wage trajectories. This procedure becomes necessary as the wage of individual i cannot be observed in two firms simultaneously. In the following, wage trajectories are estimated for each LIAB

[8] Basically, it is impossible to ascertain voluntary mobility of individuals because the data does not provide information on whether a quit occurred or whether the employee was, e.g., laid off.

[9] Individuals are allowed to receive public benefits or, for example, to be employed part-time eight days before entering the new establishment.

establishment at which an employee is employed. The main goal is the prediction of a wage trajectory that displays the individual's firm-specific wage path with respect to experience.

Some studies focus on returns to tenure within a particular firm (Topel 1991), as this relationship can be interpreted as firm-specific human capital. Another strand in the literature argues that wage growth is essentially attributable to general labor market experience (Altonji and Shakotko 1987), industry-specific experience (Parent 2000), or occupational experience (Zangelidis 2008, Kambourov and Manovskii 2009a, 2009b). Schönberg and Gathmann (2010) show that human capital is portable to a large extent when individuals move to similar occupations with similar tasks (also see the concept of task-specific human capital of Gibbons and Waldman 2004). In sum, the literature examines wage growth primarily determined by the accumulation of human capital (see Becker 1964, 1993). This analysis basically focuses on rewards to labor market experience as it can be interpreted as general human capital which is expected to be more portable than firm-specific human capital.

Starting from a simplified framework with one mobility event, the paper continues to obtain a more complex setting by inclusion of multiple transitions. The basic framework considers wage trajectories at different employers simultaneously and consists of four different scenarios which are displayed in Figure 2.1. Workers start their careers at employer 1 and subsequently switch to employer 2; as can be seen, extensive differences in mobility are illustrated. Scenario 1 depicts upward mobility in combination with permanently higher wages. Individuals, thus, are able to achieve permanently higher wages at employer 2 in comparison to employer 1. Mobility, further, is characterized by a wage markup independently of the period of realized mobility.

Scenario 2 illustrates that workers who change to jobs with lower wages might also suffer permanently lower wages. In this case, workers change to lower wages regardless of the period of realized mobility. One possible explanation is involuntary mobility where individuals might accept permanently lower wages because the alternative is nonemployment. In addition, workers may accept a job with a lower wage in order to maintain their labor market insider status (Lindbeck and Snower 2002). Alternatively, this type of mobility might be attributable to non-wage amenities (Nosal and Rupert 2007). The sheer number of studies on the acceptance of wage cuts enforces this analysis, which also informs about the persistence of lower wages. Scenario 3 describes a situation where an individual is employed at firm 1, where the wage is larger in firm 2, and then changes jobs when the wage of employer 1 exceeds the wage of employer 2.

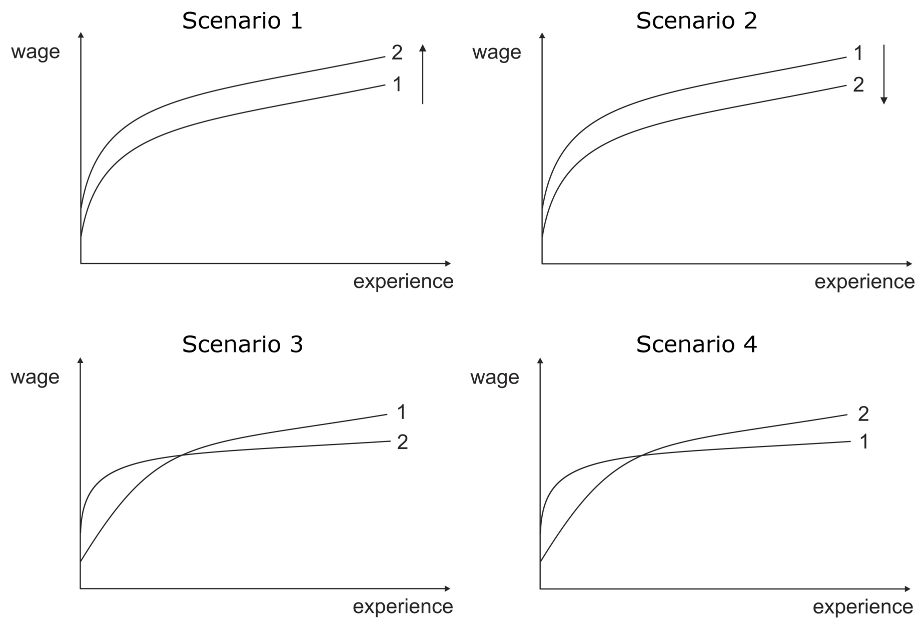

Figure 2.1: Mobility scenarios

Of foremost interest is the last scenario in Figure 2.1, because it provides an explanation for mobility to lower wages, to higher wages, and to equal wages. Changing to equal wages implies mobility at the intersection point of the wage trajectories. Mobility of this type is wage maximizing because the individual achieves the upper wage path of both wage trajectories. Mobility to lower wages is characterized by mobility before the intersection point such that individuals pay for higher future wages by immediate wage cuts. The interpretation of investments in future wage growth as proposed by Connolly and Gottschalk (2008) and Postel-Vinay and Robin (2002), hence, becomes apparent. Finally, scenario 4 illustrates mobility with wage markups for workers who are mobile after the intersection point. Analogously to Bingley and Westergard-Nielsen (2006), the probability of mobility increases synchronously with the wage differential between employers 1 and 2.

In the following, I describe the empirical procedure in more detail. As the LIAB is set up as a panel, I can make efforts to control for unobserved individual heterogeneity. In order to check for the prevalence of individual-specific effects, I apply the Breusch-Pagan Lagrange multiplier test (Breusch and Pagan 1979). The test rejects that there is no individual heterogeneity at the 1% level. Considerations about the structure of unobserved individual characteristics tend to favor fixed-effects estimation because unobserved individual characteristics are expected to be correlated with the regressors. In line with this consideration, the formal test developed by Hausman (1978) reveals that fixed-effects estimation is superior in comparison to random-effects estimation.[10] As a consequence, the fixed-effects estimator is applied.

The specification is given as:

$$w_{ijt} = \beta_{ij} + \beta_{1j}(experience)_{it} + \beta_{2j}(experience)^2_{it} + \delta'_j X_{it} + u_{ijt} \tag{2.1}$$

w_{ijt} is the log wage of individual i in firm j in period t, u_{ijt} is the error term, and X_{it} corresponds to a set of further control variables.

X_{it} consists of tenure (squared), dummy variables for the occupation, and dummy variables for the educational degree which are all shown in Table A2.1. The estimation approach, thus, follows an extended Mincer-type wage equation which includes the main individual characteristics surveyed in the data. Clerks and secondary school level I certificate with completed vocational training provide the usual reference categories. If establishments do not employ any clerks, I choose another group as reference category in order to avoid problems regarding multicollinearity. Analogously, the same holds if a LIAB establishment does not engage any worker with a secondary school level I certificate with completed vocational degree.[11]

In the next step, the establishment-specific estimates are used to predict individual

[10] Note that I perform the Breusch-Pagan Lagrange multiplier test and the Hausman test using the entire sample. The corresponding p-values are <0.001.

[11] This might introduce similar problems as in the Oaxaca-Blinder decomposition which is problematic when it comes to calculation of the contributions from indicator variables because the result depends on the reference group (see Oaxaca (1973), Blinder (1973), and Oaxaca and Ransom (1998)).

wage trajectories in different firms.[12]

$$\hat{w}_{ijt} = \hat{\beta}_{ij} + \hat{\beta}_{1j}(experience)_{it} + \hat{\beta}_{2j}(experience)^2_{it} + \hat{\delta}'_j X_{it} \tag{2.2}$$

The wage differentials between two different firm j and j' can, then, be calculated by application of the counterfactual wage paths in equation (2.2).

$$\begin{aligned} \hat{w}_{ijt} - \hat{w}_{ij't} = & \\ & [\hat{\beta}_{ij} + \hat{\beta}_{1j}(experience)_{it} + \hat{\beta}_{2j}(experience)^2_{it} + \hat{\delta}'_j X_{it}] \\ & - [\hat{\beta}_{ij'} + \hat{\beta}_{1j'}(experience)_{it} + \hat{\beta}_{2j'}(experience)^2_{it} + \hat{\delta}'_{j'} X_{it}], \ \forall\ j \neq j' \end{aligned} \tag{2.3}$$

Table 2.1 presents the calculation of the different mobility patterns illustrated in Figure 2.1. $\hat{w}_{ijt}$ is the predicted log wage based on equation (2.2), while equation (2.3) shows the calculation of the differences in accordance to the wage specifications. Note that the main focus is on workers who are mobile exactly once. For each mobile individual changing from one LIAB establishment (i.e., firm 1) to another LIAB establishment (that is, firm 2), the wage differential between the predicted wages is used to determine the mobility scenario.

[12] I only calculate predictions if establishments are observed in the sample. This assures that workers really are able to work at either establishment 1 or establishment 2. In addition, workers who leave the sample for at least one period are not considered for the analysis of mobility but, however, these individuals are included in the firm-specific estimation procedure. As the wage is censored in the data, the predictions correspond to a lower boundary of the within-firm wage.

Table 2.1

Description of the calculation of the various mobility scenarios

Predicted between-firm wage differential of individual i in t	$t_{1,i}, \ldots, t_{int,i}$	$t_{int+1,i}, \ldots, t_{T,i}$	Scenario (see Figure 2.1)
$\hat{w}_{i1t} - \hat{w}_{i2t}$	< 0	< 0	Scenario 1*
$\hat{w}_{i1t} - \hat{w}_{i2t}$	> 0	> 0	Scenario 2*
$\hat{w}_{i1t} - \hat{w}_{i2t}$	< 0	> 0	Scenario 3**
$\hat{w}_{i1t} - \hat{w}_{i2t}$	> 0	< 0	Scenario 4**

* *persistent wage differential from t=1, ..., T → no intersection point (int) is existent.*

** *wage differential non-uniform in t=1, ..., int and t=int+1, ..., T → intersection point (int) is existent.*

Scenarios 1 and 2 verify steady between-firm wage differentials.[13] Scenarios 3 and 4 exhibit exactly one intersection point. Note, however, that the inclusion of individual-specific time-variant variables in the Mincer-type wage equation might produce saw-blade-shaped counterfactual wage trajectories that may cause multiple intersection points. In this case, wage paths cannot precisely be assigned to one of the scenarios described above. For this reason, I exclude wage trajectories with multiple intersection points from the analysis of the different mobility scenarios.

2.3 Results

2.3.1 Main results

The empirical investigation provides 19,741 individual wage profiles of individuals who are mobile exactly once and in accordance to the scenarios shown above. Application

[13] Scenarios 1 and 2 should be examined in a more differentiated manner. The wage profiles can exhibit an intersection point when enlarging the sample horizon. Hence, mobility as displayed in scenario 1 could change to scenario 3-type mobility, whereas mobility to permanently lower wages could change to scenario 4-type mobility before the intersection point. Here, it is stipulated that wage trajectories do not intersect during the observed working career horizon of individual i.

of the counterfactual wages according to the specification enforces exclusion of 455 individual wage paths which intersect at least twice. The corresponding maximum is four intersection points. The average observations period for the counterfactual wages equals about five years. Figure 2.2 displays the frequencies of the mobility scenarios. Most individual transitions are characterized by a steady between-firm wage differential, as suggested by scenarios 1 and 2. Together, at least seven in ten of the wage paths are characterized by steady between-firm wage differentials. Both procedures show that a low fraction of workers is mobile in accordance with scenarios 3 (12.91%) and 4 (12.77%).

Scenario 1 addresses the first question raised in the introduction and reveals whether workers achieve permanently higher wages when changing employers. The last bar of Figure 2.2 shows that upward mobility (in combination with permanently higher wages) is most common across workers changing establishments once.[14] According to the results, the period of realized mobility does not matter in order to achieve a wage markup for about eight in twenty workers. This finding is in line with other studies which show that workers frequently change to higher wages.

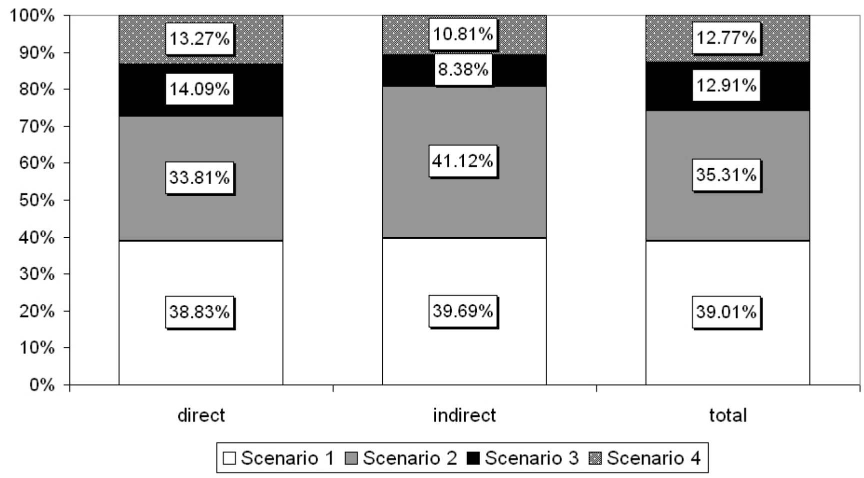

Figure 2.2: Frequencies of the mobility scenarios

Includes individuals who changed firms exactly once during the observed period

Number of observations: N_{all} = *19,741;* N_{direct} = *15,672;* $N_{indirect}$ = *4,069*

[14]Note that establishment-to-establishment transitions within the same employer might affect the results. If workers start at employer-specific low-wage ports of entry and are promoted to a high-wage establishment afterward, I conclude that workers are mobile to permanently higher wages. This paper, thus, is not responsive to this special type of investment in future wage growth.

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

Apart from that, workers are also shown to be mobile to establishments with permanently lower wages. In fact, I find that individuals are frequently mobile in accordance to scenario 2. About seven in twenty workers change to a permanently lower paying establishment. This result might be explained by job-specific amenities in the new establishment. Alternatively, involuntary job reallocations might enforce the acceptance of permanently lower wages when the alternatives are unemployment or non-employment. In addition, mobility toward less working hours in the new job compared to the previous one might introduce permanently lower wages.[15] In brief, Figure 2.2 reveals that a large share of workers changes to permanently lower wages.

Distinguishing between direct and indirect mobility reveals that mobility to permanently lower wages is more common across indirectly mobile workers (41.12%) in comparison to individuals changing directly (33.81%). In other words, workers who are not changing employers within eight days are more likely to suffer permanent wage cuts in comparison to workers who change employers within eight days. This can be explained by involuntary mobility, whereby workers accept lower wages because the only alternative is unemployment. In addition, employers are more likely to hire full-time employees in comparison to unemployed individuals. Workers possibly accept a job with permanently lower wage in order to maintain their labor market insider status (Lindbeck and Snower 2002). Workers who change indirectly, in turn, are said to have failed in keeping their (full-time) employee status and thus must accept lower paying jobs in order to re-enter the full-time workforce.

The large share of mobility to permanently lower wages within eight days is, however, in line with the literature on involuntary job reallocations. In the words of Jolivet et al. (2006, p. 882), "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". This is in line with the paper of Jolivet (2009), according to which workers accept

[15]Remember that the data lacks information on individual working hours. It is to expect that working time is comparable across jobs because of pure consideration of employees in full-time jobs.

lower wages because the only alternative is non-employment (or unemployment). The large share of downward mobility in this data suggests the existence of such involuntary job reallocations.

The third question raised in the introduction can be answered by examination of scenario 4 in combination with scenario 2. Almost all individuals who are mobile in accordance with scenario 4 leave the first employer before the intersection point which implies that the immediate wage cut pays off in the future. In fact, consideration of the baseline predictions reveals that 2,498 workers (direct: 2,063; indirect: 435) pay for steeper wage growth at the new employer by immediate lower wages. This allows for the calculation of the share of investments in future wage growth among all wage cuts. The results clearly suggest that only few wage cuts act as investment in future wage growth (26.38%[16]). The results, thus, suggest that the majority of wage cuts are of permanent nature. A possible explanation is that expectations about future wage growth are hardly to conduct. Workers, however, also might accept lower wages because of job-specific (non-wage) amenities.

The fraction of investments in future wage growth differ by distinction of direct and indirect mobility. Compared to the old employer, steeper wage growth at the new employer is more frequent across direct transitions (28.03%) compared to indirect mobility where the baseline specification finds that 20.64% of immediate wage cuts pay off in the future. This result illustrates that workers who change jobs with virtually no intervening unemployment are more successful in turning immediate wage cuts into higher payoffs at the new employer.

Table 2.2 presents descriptive statistics on the difference between realized mobility and the predicted intersection point in scenario 4. On average, workers are shown to invest for about two years. Some workers seem to invest for a very long time horizon of 13 periods until the wage paths intersect. This finding can be interpreted in the way that investments in future wage growth correspond to individual career planning in the very long run. For individuals changing after the intersection point, the data do

[16] $\frac{\text{early Scenario 4-type mobility}}{\text{Scenario 2 + early Scenario 4-type mobility}} = \frac{2{,}498}{6{,}971+2{,}498} = 0.2638$

not provide evidence for comparable time spans. The corresponding maximum equals 4 years. This result illustrates that the quit probability synchronously increases with the wage differential introduced by mobility after the intersection point. In sum, Table 2.2 shows that the wage choice is interpretable only in the long run. At least for some individuals, it is expected that downward mobility might become an investment in future wage growth under an extended time horizon.

Table 2.2

Deviation of the intersection in case of scenario four

	Observations	Difference between realized mobility and predicted intersection point (mean)	Range
Scenario 4-type mobility	2,520	-1.9048	[-13; 4]
		(1.8131)	

Standard Deviations in parenthesis.

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations).

In order to assess the long-run consequences of mobility, I consider only individuals who are observed for at least eight periods and who are mobile within the first five years. This allows for an observation of wage paths for at least three periods after the establishment change. The number of observations decreased substantially because only 2,712 of the 19,741 initial observations remain. Figure 2.3 shows that mobility to permanently higher wages is substantially higher for this subsample of workers. Mobility with permanent wage cuts remains relatively stable for direct mobility while indirectly mobile workers are significantly less affected by permanent lower wages. Again, more than one in three individuals changes establishments at the price of permanently lower wages. The share of scenario-4 type mobility is very low for this sample. For this reason, an extended time horizon might not necessarily increase the share of investments in future wage growth. However, this subsample does not suggest this pattern because most individuals are mobile with wage markups.

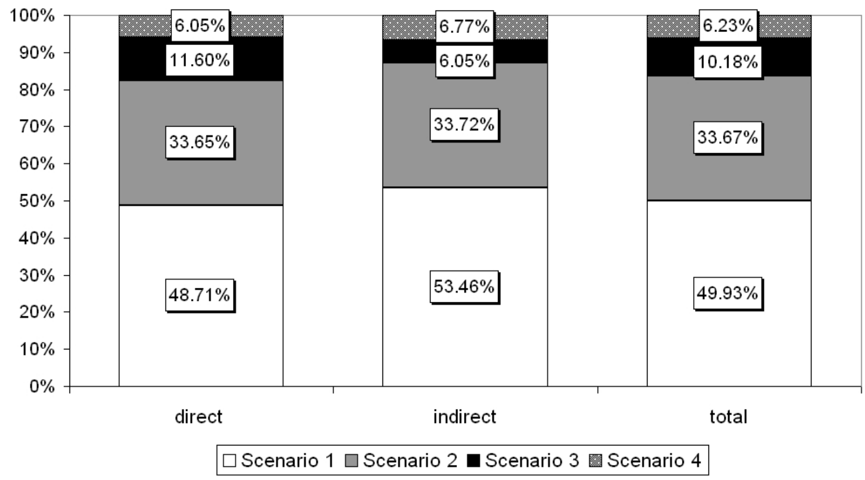

Figure 2.3: Frequencies of the mobility scenarios for an individual panel length of at least eight periods

Includes individuals who changed firms exactly once during in the first five years

Number of observations: N_{all} = *2,712;* N_{direct} = *2,018;* $N_{indirect}$ = *694*

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

So far, this analysis has been restricted to workers who are mobile once during the observed career horizon. Since Germans hold more than two lifetime jobs, this paper examines workers who are mobile more often than once. In addition, I also make effort to include wage paths with multiple intersection points of individuals. The following analysis, then, focuses on the average wage path of all mobile workers. Accordingly, this part of the analysis is adequate to conclude whether the above presented scenarios describe real world mobility to a considerable extent or whether the scenarios above are hardly to justify.

The data illustrate that individuals are mobile up to seven times, which corresponds to eight predicted wage paths. Table 2.3 presents the average predicted wages at different employers. It reveals that the average predicted lifetime wage is economically significantly decreasing when changing to employer 2. Additional transitions help to increase individual lifetime wages again. From this it follows that frequent mobility increases average lifetime earnings. Increasing wages are in line with Borjas (1981), who suggests that job mobility results on average in a wage markup. Note that the counterfactual wage trajectories at employer 1 and 2 vary substantial over the lifetime.

Table 2.3

Average counterfactual wages at different employers

	Observations	$\hat{w}_{ijt}$
Establishment 1	128,659	4.7962
		(11.2656)
Establishment 2	128,659	4.4589
		(12.9566)
Establishment 3	13,394	4.8101
		(1.0157)
Establishment 4	1,587	4.8761
		(0.2473)
Establishment 5	243	4.8827
		(0.1749)
Establishment 6, 7, 8	less than 50 observations available	

Standard Deviations in parenthesis.

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations).

For a more detailed examination, the average predicted wages are plotted in Figure 2.4. More specifically, for each year of experience, I calculate the average predicted wages. I only display four counterfactual wage trajectories for two reasons. First, German workers hold an average of four lifetime jobs. Second, few observations remain after transition three. A basic finding which can be found in all of the following graphs is that the predicted wage paths at employers 1 and 2 vary substantially across years of experience while the counterfactual wages at employers 3 and 4 are less dispersed. The large variation, however, is introduced by the peaks after 25 years of experience.[17] Consistent with the findings above, the graph provides an explanation for the large share of mobility to permanently lower wages. From two to 21 years of experience, the average counterfactual wage at employer 1 exceeds the one at employer 2. This suggests that mobility within

[17]The peaks might be explained by the low number of observations.

that particular period might frequently be characterized by permanent lower wages when not accounting for changes in other wage determinants such as tenure. Note that multiple intersection points are observed. For this reason, the *stylized* scenarios described above only present an approximation of wage mobility.

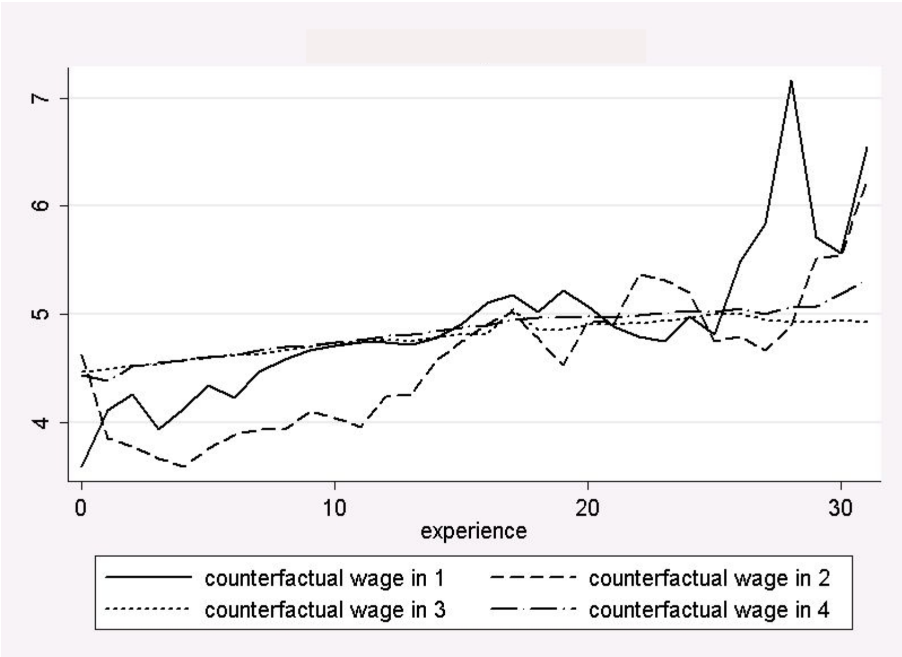

Figure 2.4: Average counterfactual wage paths at different establishments by experience

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

Figure 2.5 turns the attention to the average counterfactual wage paths as a function of tenure. The graph shows that the average predicted wage trajectory in establishment 1 and employer 2 vary substantial. Similar to Figure 2.4, the main variation is introduced by the fluctuations in the late years of tenure. The average predicted wage path at employer 1 suggests a positive trend in the rewards to tenure until 15 to 17 years of tenure. Afterward, the counterfactual returns to tenure exhibit a negative trend. The average counterfactual wage path at employer 2 provides an explanation why so many workers are mobile to permanently lower wages. Tenure-related average predicted wage growth at this employer is not steep enough to compensate for immediate wage cuts. In fact, it is even negative rather than positive. The average counterfactual wage paths at establishments 3 and 4 reveal increasing average returns to tenure in the long run whereas being employed at firm 4 yields higher average returns to tenure.

Figure 2.5: Average counterfactual wage paths at different establishments by tenure

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

To sum up, the results show that only few wage cuts act as investments in future wage growth while the majority of immediate wage cuts are associated with permanently lower wages. Analogously to Rosen (1974, 1987), jobs consist of bundles of various characteristics with implicit prices. In other words, workers might accept even permanently lower wages if the current job is *better* compared to the previous job. This includes non-wage characteristics as well as monetary characteristics. For this reason, the discussion of Nosal and Rupert (2007) about job-specific amenities affecting wage choice over the life cycle becomes relevant. However, further research is needed to assess the reasons why workers change to (permanently) lower wages. Another finding of this paper is that wage markups in the realized period of mobility frequently stem from transitions to a firm which offers higher wages during the observed career horizon. This result is in line with literature on wage markups and upward mobility.

2.3.2 A subgroup analysis

This section presents average counterfactual wage paths at different firms by years of experience and by different groups in order to conclude about the consequences of mobility for heterogeneous groups. Again, I utilize the wage regressions shown above in order to compare wages conditional on experience (squared), tenure (squared), occupation, and educational degree. The main reason for focusing on experience stems from portability of general human capital which might allow for conclusions about a wage maximizing mobility strategy. This methodology might be criticized because I compare individuals with identical experience, but with different tenure. When the period of mobility remains unconsidered, this procedure might introduce outliers because calculation of the mean wage by experience does not ensure that I compare individuals with similar tenure. More specifically, the weighted average of wages by experience might decrease sharply when considering a large share of workers who changed jobs in this period which, then, reduces wages because these individual wage profiles are calculated on the basis of zero tenure. When I consider a period where most workers have some tenure, then the counterfactual wage additionally includes tenure effects, which might increase wages considerably. Note that distinct conclusions are not possible because of the U-shape of tenure. I start with an examination of average counterfactual wage paths by education, whereas I define three skill groups of workers in analogy to Fitzenberger and Garloff (2007): The first category consists of individuals with neither a completed vocational training nor a university degree (low skilled). The second group corresponds to workers with a vocational training degree but without university degree (medium skilled). The third group are persons with a (technical) university degree (high skilled).

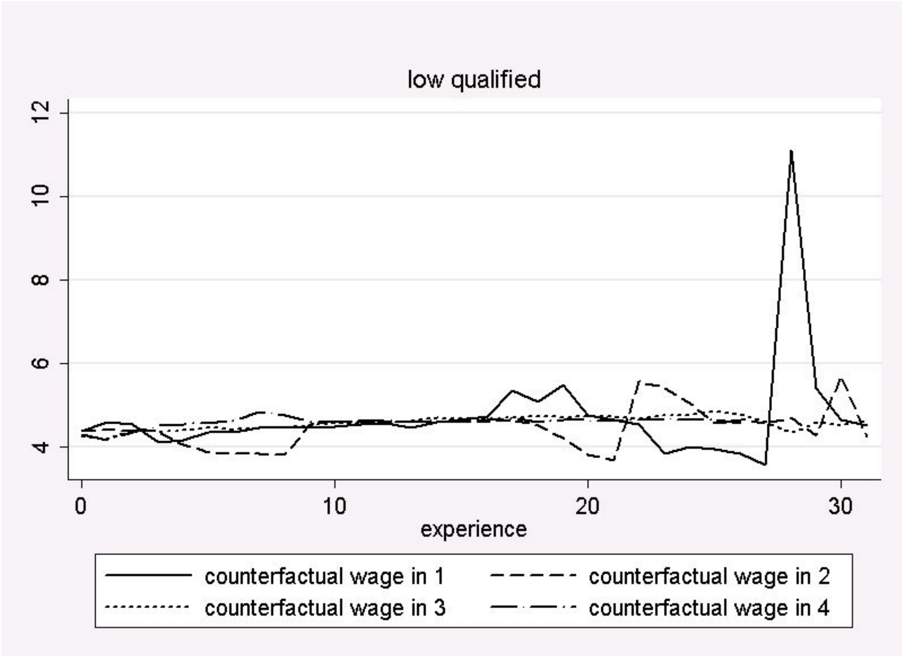

Figure 2.6: Counterfactual wage paths for low skilled workers

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

Figure 2.6 presents four counterfactual wage-experience trajectories for the group of low skilled workers. Note that statements about the fourth wage trajectory are ambiguous because of few observations. The graph reveals three intersection points for the average predicted wage trajectories of employers 1 and 2 within the first ten years. This does not allow for distinct answers regarding a wage-maximizing mobility strategy in the long run for this group of workers. Scenario 3-type and scenario 4-type mobility, however, are suggested to be usual for this group of workers in the short run. More experienced low skilled individuals seem to suffer permanently lower wages within 17 and 22 years of experience. From 22 years to 25 years, the average counterfactual wage trajectory exceeds the one of employer 2. Outliers are suggested to have sizable impact on the calculation of the average counterfactual wage after about 27 years. Figure 2.6 also reveals that the wage paths of all four employers intersect very often. In sum, distinct conclusions about a possible wage-maximizing mobility strategy are hardly to justify.

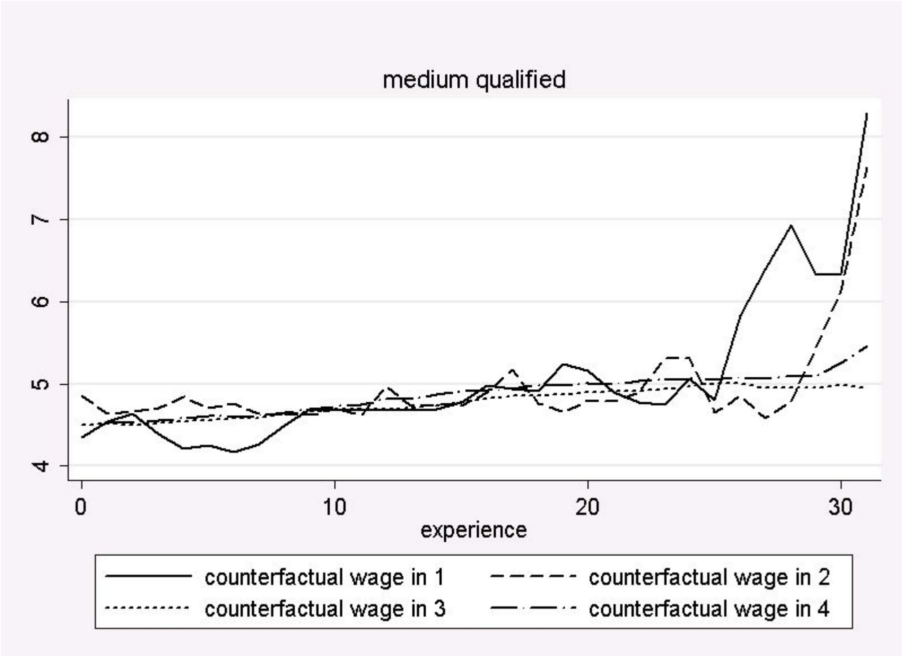

Figure 2.7: Counterfactual wage paths for medium skilled workers

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

Figure 2.7 shows that medium skilled workers are, on average, upwardly mobile after their labor market entry. Up to 17 years of experience, the average predicted wage of firm 2 exceeds the one of employer 1 during almost all years of experience.[18] The average counterfactual wage paths at employers 3 and 4 are, on average, not economically significantly higher than at employer 2 over the career horizon. Individuals with completed vocational training, however, are suggested to change employer once in order to improve their lifetime wages. Inspection of the predicted average wage paths in Figure 2.7 could lead to the conclusion that the expected wage in firm 1 is lower because this particular employer might have paid for vocational training. Literature reasonably suggests that employer and employee share the rent of vocational training after graduation while employers pay somewhat lower wages to self-trained employees. For employers without any costs for training it might be possible to offer higher wages because of no expenses for training. Employers without vocational training programs, then, might try to poach skilled workers from firms that take on trainees by offering higher wages. This, however, must not be true because we do not observe the entire working biography in the data.

[18] An exception is shown at eleven years of experience where the average counterfactual wage at employer 1 exceeds the one of employer 2.

In other words, firm 1 must not be the firm where the vocational training is originally completed.

Finally, Figure 2.8 presents the predicted average wage paths for highly skilled individuals with (technical) university degrees. On average, multiple mobility (at least two transitions) pays off over the observed career horizon. In addition, the graph suggests that multiple mobility should be executed at early stages of the working life which is in line with job shopping (Altonji and Shakotko 1987) as a promising strategy to improve wages over the career horizon. Note that the average predicted wage paths at employers 1 and 2 are significantly lower for high skilled than for medium and low skilled workers at the beginning of the career. In addition, downward mobility is indicated between firm 1 and firm 2 until about 20 years of experience. For this reason, the wage profiles displayed in Figure 2.8 suggest to stay at employer 1 or to change jobs more often than once.

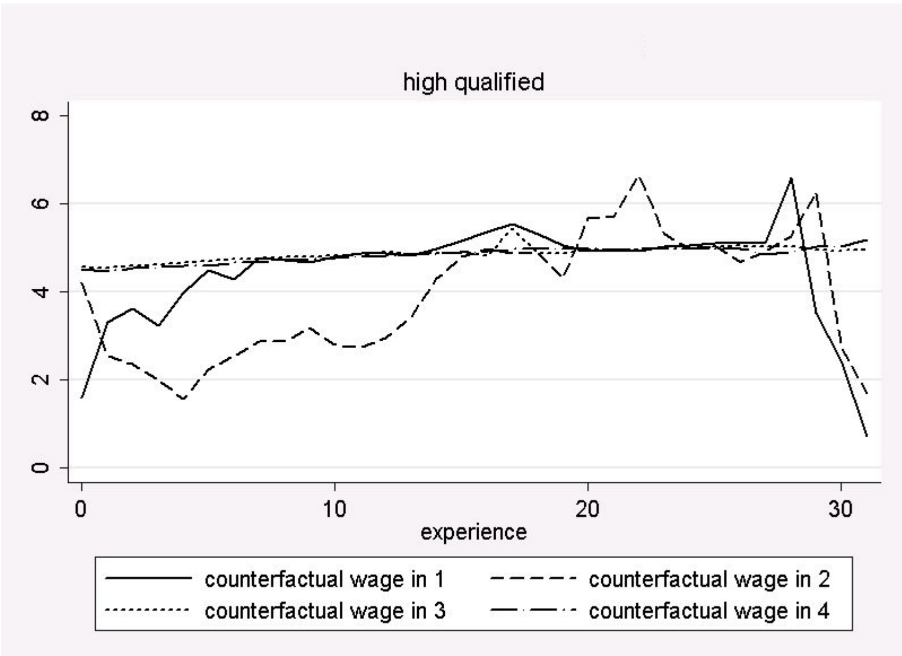

Figure 2.8: Counterfactual wage paths for high skilled workers

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

The average counterfactual wage paths presented in Figures 2.6 to 2.8 might be a first step to provide a guide to wage-maximizing mobility. Unfortunately, the predicted average wages do not reveal a distinct or wage-maximizing mobility strategy. One job change, however, is suggested for medium skilled workers while multiple job changes are suggested for workers with (technical) university degrees. For individuals without completed vocational training and (technical) university degree, a clear wage maximizing strategy cannot be shown because of multiple intersection points of the wage trajectories.[19]

The following figures display the average counterfactual wage paths by occupation while experience is limited to 25 years because of outliers at later stages of experience.[20]

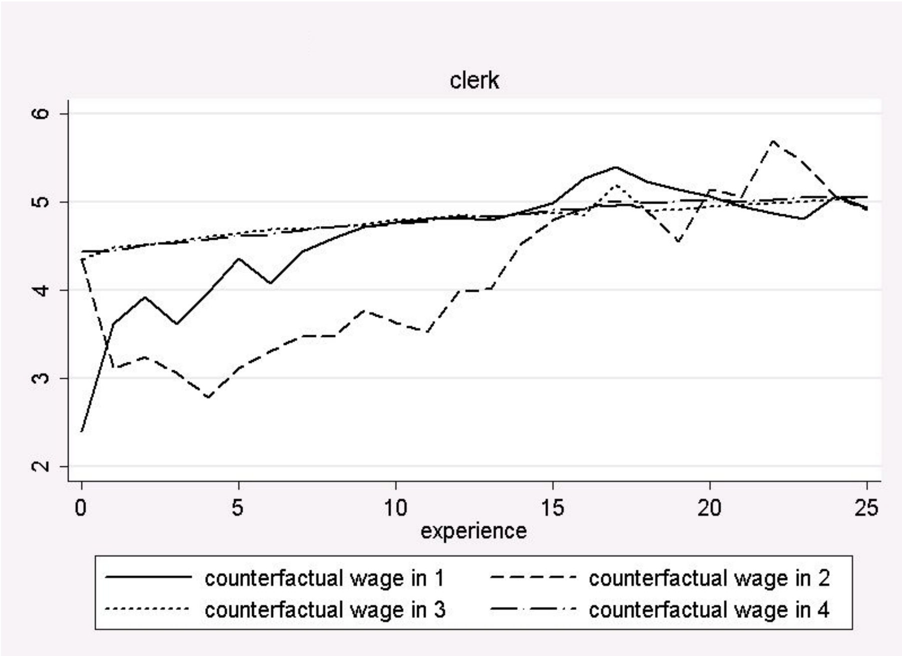

Figure 2.9: Counterfactual wage paths for clerks

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

The low starting wage of the highly skilled workers with academic education might be explained by the average counterfactual wage paths of clerks (see Figure 2.9). Irrespective of the employer, an upward trend in wages is obvious for this particular group of workers over years of experience. It is also shown that the first job change of clerks might frequently be accompanied by permanently lower wages. More specifically, Figure 2.9 indicates scenario 2-type mobility between three years and 20 years of experience. Figure 2.10 shows the predicted average wages at different employers for workers with a completed vocational training in skilled jobs. Multiple intersection points do not allow

[19]As above, the peaks after 25 years of experience stem from low number of observations. I do not display any confidence bands in the figures because of multiple intersection points of the confidence bands which indicate insignificant differences.

[20]See the corresponding Figure A2.1 in the Appendix for the counterfactual wage paths over the complete time horizon. At very late stages of experience, the counterfactual wage paths at establishments 1 and 2 result in a rapid increases or decreases, respectively.

for conclusions about a wage-maximizing career strategy. In contradiction to workers in skilled jobs, workers (without vocational degree) in unskilled jobs are suggested to stay at employer 1 for three years (see Figure 2.11). Afterward, the wage trajectories show that (multiple) job mobility increases the wage until about 15 years of experience. After 15 years of experience, however, peaks and multiple intersection points mitigate this finding. For this reason, unskilled workers are suggested to profit from job shopping at the very beginning of the career. As suggested by Figure 2.12, technicians should stay at employer 1 for about two to three years and then change to another employer in order to maximize the wage.[21] A further transition pays off after about seven years of experience while a third transition is also suggested by the figure. This suggestion for a wage maximizing career path only holds for years of experience below 16 because of multiple intersection points afterward.

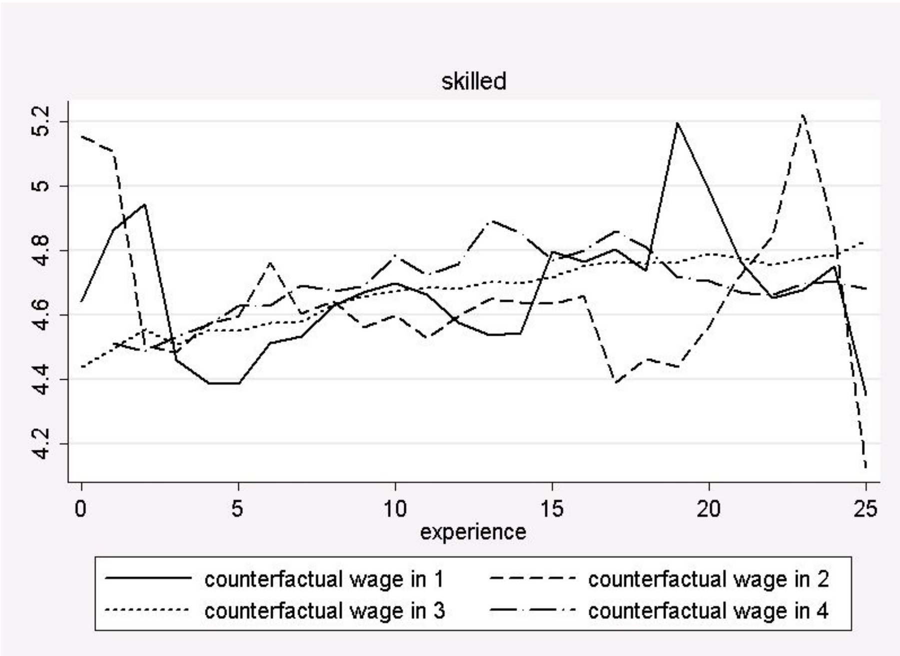

Figure 2.10: Counterfactual wage paths for skilled workers

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

[21] For the group of technicians, only very few observations are available. For this reason, the graphs are to interpret with caution.

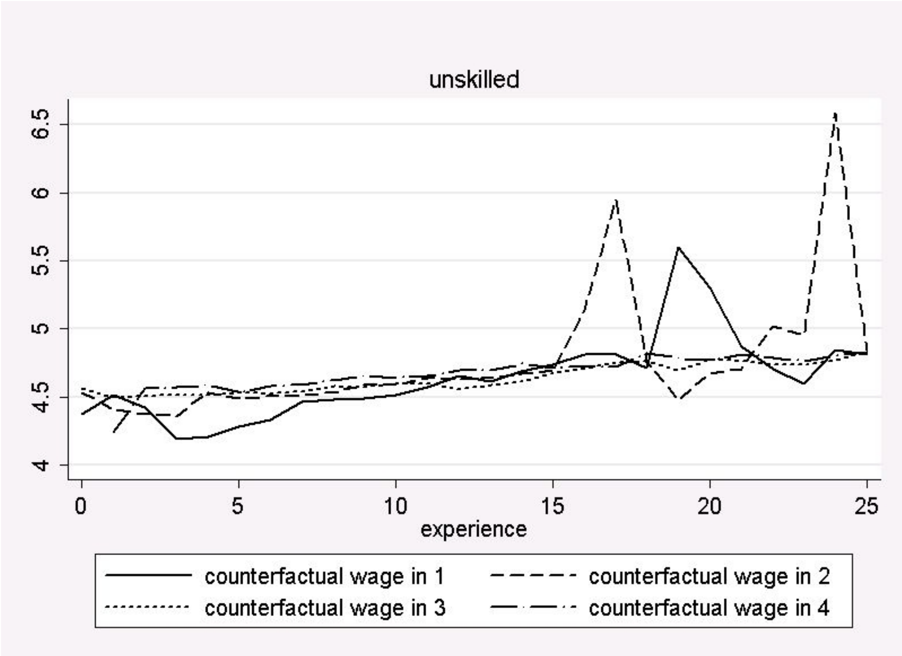

Figure 2.11: Counterfactual wage paths for unskilled workers

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

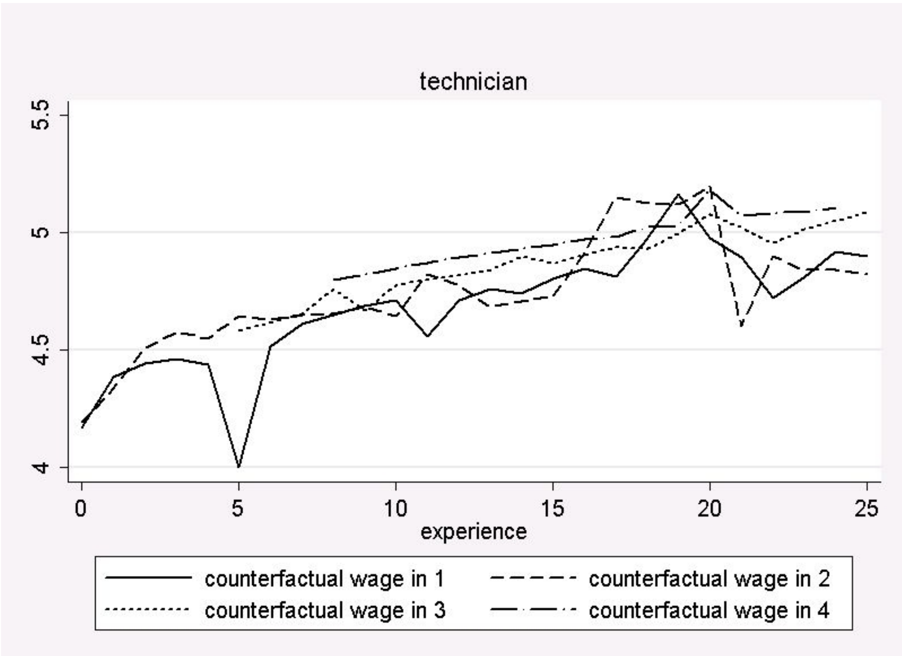

Figure 2.12: Counterfactual wage paths for technicians

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

The figures in this section indicate that mobility to lower wages as well as transitions to higher wages are common and, to some extent, upward and downward mobility depend on the years of experience in the labor market. The long-run consequence of a transition between employers, however, is hardly to assess because of multiple intersection points of the average counterfactual wage profiles at the different employers. More precisely, consideration of a long career horizon (e.g., 10 years of experience or more) leads in most cases to multiple intersection points which are not in line with the *stylized* mobility scenarios shown in section 2.2.2. In addition, workers might change employers more often than once. Hübler (1989) discusses an optimal number of mobility whereas the procedure in this paper is inadequate to conclude about an optimal number of mobility. In sum, mobility and wage growth seem to be very complex and pure consideration of average counterfactual wage paths by years of experience is one strategy to reveal the complexity of mobility.

2.4 Discussion

This paper analyzes the performance of individual employer-to-employer mobility via investigation of individual wage trajectories at multiple firms simultaneously. The four mobility scenarios in the paper contribute to an analysis of the medium-term and long-run consequences of wage cuts and wage markups in the period of mobility.

The most common scenario is mobility to permanently higher wages. In other words, compared to the previous employer, workers change to employers where they earn higher wages in the long run. This result is in line with the literature on wage markups induced by mobility (Borjas 1981, Topel and Ward 1992) and might be explained by on-the-job search for higher wages.

The analysis also shows that individuals change to lower wages. Specifically, it is shown that more than one in three workers suffer permanently lower wages at the current employer in comparison to the previous one. As discussed in Nosal and Rupert (2007), job-specific amenities might affect job mobility and wage choices throughout an individual's life cycle. Unfortunately, the underlying data set is not adequate to check for individual amenities. An alternative interpretation can be derived from the literature on job reallocations, according to which workers accept lower wages because the only alternative is unemployment. The results corroborate this interpretation because workers who are mobile within eight days are less likely to suffer permanently lower wages compared to workers changing jobs with longer intervening periods of time.

A major finding is that immediate wage cuts rarely pay off in the future. For workers who are mobile exactly once, more than 70% of immediate wage cuts result in permanent lower wages at the current employer. According to the different types of mobility (direct or indirect), less than 30% of individuals accept wage cuts in order to improve their future wages, as proposed by Connolly and Gottschalk (2008) and Postel-Vinay and Robin (2002). The results also show that individuals change employers years before the pay-off period. Note that I consider an average observation period of about five years which might be to small for an examination of long-run consequences of wage cuts, although few individuals can be shown to invest in future wage growth in this chapter.

Finally the paper suggests that further research is needed to assess the reasons of mobility to lower wages. Possible explanation arises from considerations about job-specific non-wage amenities which might affect mobility decisions. In line with the results, jobs are not necessarily ranked according to their wages and, therefore, conclusions about transitions based on the wage alone might be misleading. My findings directly advert to the importance to examine the importance of non-wage characteristics in more detail. Inclusion of job-related amenities to conclude about the individual trade-off reasoning between job-specific amenities and wages might be a fruitful avenue for further research.

2.5 Appendix

Table A2.1
Descriptive statistics (entire sample)

Variable	Mean	Standard Deviation
Endogenous Variable		
$\log(w)_{i,f,t}$	4.7218	0.2620
Exogenous Variables		
Experience	16.6203	7.3975
Experience^2	330.9575	238.0060
Tenure	11.4736	7.7798
Tenure^2	192.1686	209.7652
Unskilled^{i}	0.2405	0.4274
Skilled^{ii}	0.2938	0.4555
Technician^{iii}	0.0260	0.1592
Clerk^{iv}	0.4396	0.4963
Secondary school level I certificate	0.1190	0.3238
Secondary school level I certificate & vocational training	0.6812	0.4660
Advanced (technical) college entrance qualification	0.0083	0.0909
Advanced (technical) college entrance qual. & vocational training	0.0406	0.1974
Advanced technical college certificate	0.0648	0.2462
University degree	0.0860	0.2804
Number of observations	11,340,952	
Number of individuals	2,622,048	
Number of establishments	4,697	

Note: German terms:

i) nicht formal qualifiziert.

ii) Facharbeiter.

iii) Meister, Poliere.

iv) Angestellter.

Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations).

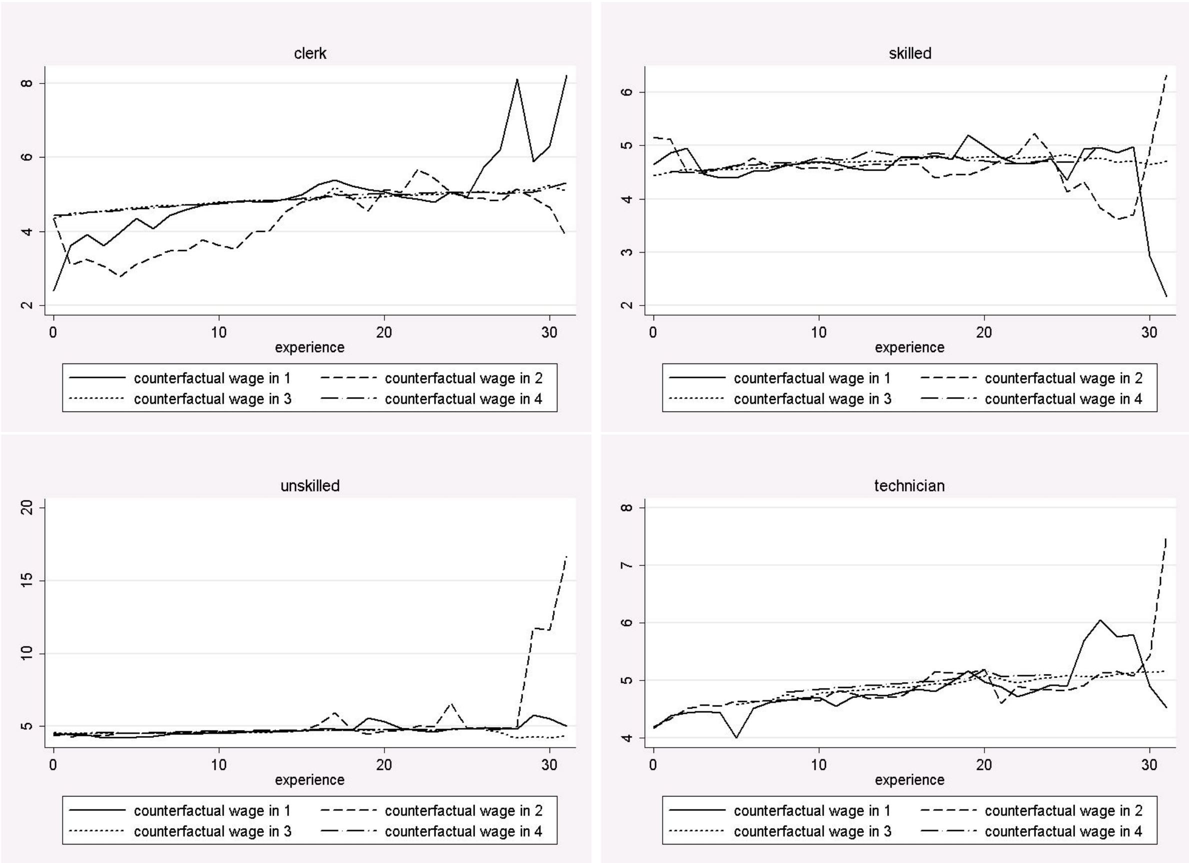

Figure A2.1: Counterfactual wage paths for complete time horizon
Data: Cross-sectional model of the LIAB, Years 1993-2006 (own calculations)

3 Relative Wage Positions and Quit Behavior

This paper utilizes a large linked employer-employee data set in order to analyze the impact of relative wage positions on the individual decision to quit a job. At first, the analysis concentrates on linear effects and finds that a possible signal effect might dominate a possible status effect in quit decisions. In a next step, we establish consideration of nonlinear effects which improves understanding of quits substantially. Finally, we are interested in the relationship between the change in wages and the change in relative wage positions when workers change jobs. The results suggest that workers who experience a loss in their relative wage positions are also more likely to have a wage cut.[1]

[1]This chapter is co-authored with Prof. Dr. Christian Pfeifer, Institute of Economics, Leuphana Universität Lüneburg. The chapter is a revised version of 'Relative Wage Positions and Quit Behavior: New Evidence from Linked Employer-Employee Data', Leibniz University Hannover Discussion Paper, 438, Hannover. This study uses the Cross-sectional model of the Linked-Employer-Employee Data (LIAB) (Years 1996-2006) from the Institute for Employment Research (IAB). Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the IAB and remote data access.

3.1 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, recent labor turnover literature points to the importance of fair wages and status concerns of workers as well as to the fact 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 regard to both streams of the literature, we can contribute empirical findings from German linked employer-employee data, which allows us to compute measures for workers' relative wage positions within their firms and to assess their impact on decisions to quit full-time employment. Our sample contains almost three million annual observations of nearly 850 thousand full-time employed male prime-age workers in more than four thousand West German firms for the period from 1996 to 2005.

Our paper enhances the analysis of relative wage positions on the decision to quit a job in several ways. At first, we are contribute findings on this relationship using German data. Second, we apply several measures for the investigation of relative wage positions in order to strengthen the generality of our results. Third, we include nonlinear effects of relative wage positions on the decision to quit. Finally, we are interested in whether a change in the relative wage position balances out monetary losses. Our findings are in line with other studies which suggest that relative wage positions have a significant impact on the probability to voluntarily quit a job. Regarding possible nonlinearities, we suggest the prevalence of a U-shaped relationship whereas we also come up with a proposal to interpret this result. In addition, the results reveal that losses in the relative wage position are, on average, accompanied by wage cuts. For this reason, individuals who experience decreasing relative standing at the new establishment compared with the previous one are also more likely to change to lower wages.

The paper is structured as follows. The next section illustrates briefly the basic theoretical framework and our research hypotheses. Section 3.3 describes our data set, main variables, and econometric models. In section 3.4, 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 3.5.

3.2 Basic theoretical framework and hypotheses

The relationship between wages and the decision to quit can be incorporated in the broad framework of individual utility maximizing behavior. In equation (3.1), utility U of individual i who works in firm j at time t is a simplified function of the individual absolute wage (w_{ijt}^{abs}), the individual relative wage position within the firm (w_{ijt}^{rel}), and other individual and job characteristics (X_{ijt}).[2] Moreover, we assume that the individual probability to quit a job in firm j during period t is negatively correlated with utility as described in equation (3.2) (Freeman 1978, Akerlof et al. 1988, Clark et al. 1998, Clark 2001, Clark and Georgellis 2006, Lévy-Garboua et al. 2007).

$$U_{ijt} = U_{ijt}(w_{ijt}^{abs}, w_{ijt}^{rel}, X_{ijt}) \tag{3.1}$$

$$Pr(quit = 1|U_{ijt}) = Pr(quit = 1|w_{ijt}^{abs}, w_{ijt}^{rel}, X_{ijt}) \tag{3.2}$$

Standard economic theory (e.g., search models, efficiency wage models) usually accounts explicitly for absolute wages, which should positively affect a worker's utility (Salop and Salop 1976, Salop 1979, Akerlof 1982). Our main focus is 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

[2] Note that this framework only considers the current wage while individuals might also consider wage changes over time. The framework presented here rather considers myopic individuals but, however, measures for wage growth can easily be implemented in the empirical analysis.

(w_{ijt}^{rel}), which includes wages of co-workers as comparison income, is, however, ambiguous. Studies find support for both, either status as well as signal. More specifically, Clark et al. (2008) corroborate status concerns while the more recent study of Clark et al. (2009) favors the signal effect which are described in the following.

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. A lower relative wage might be perceived as unfair and of low social status (Adams 1965, Garner 1986, Akerlof and Yellen 1990, Clark et al. 2008, Frank 1984a, 1984b), which consequently decreases utility and increases the quit probability, ceteris paribus. This is called the 'status effect'. The relative wage position within a firm, in turn, 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 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.[3] This illustrates that the effect of relative wage positions might be ambiguous because good fortune of co-workers could introduce either jealousy (low own status) or it might provide information about own career opportunities (signal). Note that status and signal are contradictory in that the workers in low relative wage positions have large career prospects but low status, while in the case of workers in high relative wages the reverse is true.

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

[3]See Clark et al. (2009) for an extensive discussion of status and signal effects. The study of Carporale et al. (2009) who analyze the impact of comparison income on happiness is a good example to explain both effects. The authors find that higher reference income lowers satisfaction in Western European countries while in Eastern Europe countries, reference income is likely to increase satisfaction. The results for East Europe are consistent with the signal effect because reference income is likely to be a source which provides information about own future economic possibilities. In other words, the higher the income of the comparison group, the higher my own level of satisfaction because of prosperous future perspectives. The results for West European countries correspond to the status effect which predicts that higher reference income lowers my own level of satisfaction because of low own social status within the reference group.

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

Many studies find a significant relationship between relative wage positions, satisfaction, and labor turnover (see, e.g., Galizzi and Lang 1998, Clark et al. 2008, Clark et al. 2009). One question, however, remains unanswered: Is the relationship between relative wage positions and individual mobility decisions simply linear or does it contain a nonlinear component? An examination of nonlinear effects becomes necessary because Clark and Oswald (1996) show in their descriptive analysis that, on average, the most satisfied male workers are in the highest and lowest income quintiles and the least satisfied in the middle income quintiles. Consideration of possible nonlinearities also might shed light on the relationship between relative wage positions and the quit propensity since it is problematic to separate status from signal by application of one single measure for the relative standing.

Consider a worker i who obtains utility from status as well as from career advancement opportunities within his firm j in period t. Empirical studies usually apply a single measure for the relative wage position, but this renders separation of status from signal problematic, because status is a byproduct of signal and vice versa.[4] In the following, the wage rank of a worker within his firm is applied to measure the relative standing within a particular firm so that U_{ijt}^{status} and U_{ijt}^{signal} are functions of rank. To calculate the rank, workers in each firm are sorted in the ascending order of their wages in each period t. When the workers are sorted, those with the lowest wage occupy rank zero and those with the highest wage hold the highest rank. Workers in rank zero are not allowed to obtain utility from status because they occupy the lowest rank. As a consequence, it is assumed that these workers have the lowest status within a particular firm. Workers in the highest rank, in turn, are imposed to achieve no utility from signal because of the lack of future career advancement opportunities. Apart from that, workers with the highest rank achieve all their utility from status while workers in rank zero only accumulate utility

[4]See Clark et al. (2009) for a discussion of this particular problem.

from future career prospects. A worker's total utility in equation (3.3), given below, is the sum of the utilities from status, signal, and other job characteristics X, which include, for example, own absolute wages and tenure.

$$
\begin{aligned}
U_{ijt} = U_{ijt}^{status}(rank) + U_{ijt}^{signal}(rank) + U_{ijt}^{other}(X) \\
with \quad \frac{\partial U_{ijt}^{status}}{\partial rank} > 0; \quad \frac{\partial^2 U_{ijt}^{status}}{\partial rank^2} < 0 \\
\frac{\partial U_{ijt}^{signal}}{\partial rank} < 0; \quad \frac{\partial^2 U_{ijt}^{signal}}{\partial rank^2} < 0
\end{aligned}
\tag{3.3}
$$

It is assumed that individual utility from status increases with rank and workers in higher positions gain lower additional utility from climbing up the hierarchy. The underlying reasoning stems from individual downward comparisons, that is, workers compare themselves to colleagues below their own position to make themselves feel better. Workers in the lowest rank, then, are constrained in terms of such comparisons because all colleagues occupy higher relative wage positions. Perceptions of low own status might introduce jealousy or subjective thoughts of unfairness and, hence, decrease satisfaction. Being in the middle of the hierarchy allows for individual downward comparisons which significantly increases the utility obtained from status. But still there are co-workers who have higher status. At the top of the hierarchy, individuals are expected to be generally satisfied with their own status and are assumed not to gain much additional (status-related) utility from ascensions to the highest positions. As a result, marginal utility from status is positive but at decreasing marginal rates. Individual utility from signal is described as a decreasing function of rank, which implies that workers in low relative wage positions derive larger utility as compared to those in high relative wage positions. Utility from signal is likely to be a concave function of rank. Workers in low and medium ranks have more career advancement opportunities because of the availability of further positions in the top ranks of the hierarchy. Being at the top position reduces the promotion possibilities considerably because of diminishing career advancement possibilities.

Given the individual position within a particular firm, the amount of total individual utility U_{ijt} depends on the individual valuation of status and signal. Addition of the functions U_{ijt}^{status} and U_{ijt}^{signal}, leads to an inverse U-shaped path of U_{ijt} with respect to rank.[5] This implies that workers in the highest position achieve all utility from status, while workers in the lowest rank solely obtain utility from future career prospects. However, workers in positions between the highest and the lowest rank add up utility from status and utility from signal which leads to a higher total utility compared with sole utility from either status or signal.

As discussed above, individual quit decisions are described as an inverse measure of job satisfaction. It is assumed that workers stay with their firms if total individual utility in those firms exceeds an alternative utility U_{iat} that can be obtained in another firm. Equation (3.4) shows that the probability of quitting a job is described by the probability that total utility in another firm (U_{iat}) plus an idiosyncratic error term ϵ exceeds the utility in the current firm (U_{ijt}). As related to our framework, it should be more likely that workers quit their jobs if they are in very high or very low relative wage positions, because utility is almost solely determined by either status or signal.

$$\begin{aligned} Pr(quit_{ijt} = 1) &= Pr(U_{ijt} < U_{iat} + \epsilon) \\ &= Pr(U_{ijt}^{status} + U_{ijt}^{signal} + U_{ijt}^{other} < U_{iat}^{status} + U_{iat}^{signal} + U_{iat}^{other} + \epsilon) \end{aligned} \tag{3.4}$$

$$Pr(quit_{ijt} = 1) = \Phi(\alpha_0 + \alpha_1 rank_{ijt} + \alpha_2 rank_{ijt}^2 + X_{ijt}\beta) \tag{3.5}$$

We test our hypothesis of a U-shaped relationship between status effect and signal effect by application of the probability model in equation (3.5) in which Φ is the cumulative density function of the standard normal distribution. X_{ijt} denotes a large set of control variables which are described in the following section. From the considerations about status and signal effects described above, we derive hypothesis 2.

Hypothesis 2: Workers in high relative wage ranks are more likely to change jobs because of few career advancement opportunities while workers in low relative wage ranks

[5] Linear or convex functions of U_{ijt}^{signal} also allow for inverse U-shaped utility functions. Convex increasing U_{ijt}^{status} only allows for an inverse U-shaped utility function if signal is defined as a convex decreasing function. We decided for, at least in our view, the most reasonable utility functions. We also impose that the maximum of U_{ijt}^{status} equals the maximum of U_{ijt}^{signal}.

change jobs because of their low status.

In a further step, our paper aims to shed some more light on the empirical observation that many mobile workers experience wage cuts. Table 3.1 outlines some results of recent studies on wage cuts induced by mobility. Mobility to a lower wage is common in the U.S. and in Germany. 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% of job-to-job transitions in Germany and 23% of transitions in the U.S. are accompanied by wage cuts.[6] Nosal and Rupert (2007) provide evidence for the U.S. that about two out of five (voluntarily) mobile individuals change to lower wages.

Table 3.1

Recent studies about job mobility and wage changes

Authors	Country	Mobility (in percent)		
	(data set)	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. 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.

Theoretical approaches explain voluntary and immediate wage cuts commonly as investments in future wage growth (Postel-Vinay and Robin 2002, Connolly and Gottschalk 2009). These approaches have in common that individual decisions are simplified to a monetary maximization problem in which workers maximize the long-run value of job opportunities and wages. Nosal and Rupert (2007) contribute to this literature by inclusion

[6] Jolivet et al. (2006) report the shares of mobility to lower wages also for a number of further European countries.

of job-specific (non-wage) amenities which affect the job choice and consequently individual mobility decisions. We expect that the relative wage position is such an 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 opposing signal and status effects. On the one hand, workers might accept wage cuts if they can improve their status in the new firm, which is measured as a higher relative wage 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, that is, 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' dominates 'signal effect').

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

3.3 Data and methodological remarks

3.3.1 Data set

As the interest of this study is in the relative wage positions within firms, the estimation framework requires information about workers, co-workers, and their firms. It is also desirable that the data relate to as many workers as possible in each firm 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 requirements (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 an annual survey that includes a random sample of establishments 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 1996 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. The 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. The main interest is on the daily wage which is surveyed in the data. The consumer price index surveyed by the Statistisches Bundesamt Deutschland is applied to deflate the nominal wages (annual averages, with year 2005 = 100). 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 Euro per month (or 13.33 Euro per day in 2005 wages). Another problem of the data is that wages above the upper earnings limit for social security contributions are set to the corresponding value of the ceiling. This implies that all wages above are censored in the data. To reduce the impact of the censoring, the sample is restricted to workers who do not have more than a high-school degree with completed vocational training.[7]

The sample is restricted to male, full-time employees of West German firms, who are aged between 25 and 55 years, and who do not have any academic qualification. We focus on workers exceeding 25 years of age because, at least for our sample, schooling and apprenticeship degrees are basically completed. Analogously to Galizzi and Lang (1998),

[7]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 whether regular workers compare themselves with employees with academic degrees who typically perform other tasks. This choice, however, might also be a critical one because possible frustration about low own status within an organization can be a bigger problem for those individuals who had made the effort of investing in their education. For this reason, we might exclude the group of workers for whom status might matter the most.

we limit the sample to men who are less than 55 years old as older 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 in East Germany. The data consequently report only a lower boundary for tenure and experience in East Germany. The focus on low and medium educated workers also alleviates possible criticism of Brown et al. (2008, p. 380) who state that "[...] it seems important to understand exactly how a person chooses a reference group. [...] We are forced in our econometric specification simply to assume that the workplace is the comparison set." It seems to be reasonable that medium and low educated workers compare themselves more likely with the coworkers within a particular establishment in comparison to the highly qualified workers. For example, consider highly qualified managers or highly qualified specialists who meet other specialists all around the world at conferences and meetings. These individuals might be expected to be more likely to compare themselves with specialists in other firms all around the world. Workers with completed vocational training or workers in factories, in turn, are expected to be more often in contact with a set of more similar coworkers which might be an explanation why they compare themselves to colleagues within a firm.[8]

Establishments with less than ten workers under the above restrictions are excluded from the analysis because we need to estimate earnings functions for single firms and need sufficient wage variance within firms for our analysis of relative wage positions.[9] The final

[8]See Selezneva (2010) for a summary on reference groups and income comparisons with a special focus on economies in transition. Empirical evidence for Germany finds that gender-specific comparisons are the most important, followed by comparisons within certain professions (Mayraz et al. 2009). Especially for males, income comparisons are better predictors of subjective well-being than for females. Our data allows for the identification of a more distinctive reference group which enables comparison between co-workers within the same establishment. Clark and Senik (2010) state that the most frequently cited comparison group is that of the colleagues. Accordingly, we focus on males in combination with co-workers as comparison group.

[9]Consideration of more than ten annual observations per establishment decreases the number mobility events which are already a rare event. Table A3.1 in the appendix presents number of mobility

sample for analysis, covering the period 1996 to 2005, contains 2,902,724 observations from 833,359 workers in 4,260 different firms in an unbalanced panel design. Only few observations (3.8%) comply with the upper earnings limit for social security contributions in our sample so that wage censoring is not of much concern.

The present analysis of quit behavior relies on the assumption that individuals leave their employer voluntarily. In our data, we do not have any information on the reason for the separation of employer and employee, but we expect to approximate quits via the following conditions which need to be met. The worker has been a full-time employee for two successive periods in two different establishments. Identification is possible via observation of changes in the establishment identifier which is associated with the worker. Additional information is gathered on the individual's employment relationship eight days prior to the commencement of his new job because direct transitions without intervening unemployment is more likely to be of voluntary nature. If the worker was a full-time employee at another establishment eight days before entering the new establishment, he is expected to have voluntarily quit the previous job, because he switched establishments with virtually no unemployment.[10] In sum, 5,464 mobility events are observed whereas 5,280 individuals have been mobile up to four times.

3.3.2 Wage measures

We have introduced different 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 straightforwardly as the log mean daily wage in Euro of individual i in firm j in year t surveyed

events and number of observations for different samples with respect to annual observations per establishment. In addition, it is shown that most quits occur in establishments with more than 100 workers.

[10]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." This very low share stems from 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 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.

in the data. 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.[11] 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 for every firm in every year[12] which include the worker characteristics schooling, experience[13], squared experience, and occupation.[14] A summary of all calculated (relative) wage measure is provided in Table 3.2.

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)[15], we construct the wage rank (w_{ijt}^{rank}) as well as the wage range (w_{ijt}^{range}) of a worker within his firm so that both variables lie in the unit interval (0, 1). Value one indicates that the individual is at the top of the pay scale.[16] In general, the higher the rank, the higher the worker is up the pay scale of his firm. The wage range measures the normalized distance of individual i's wage in firm j in period t to the lowest wage in his firm as proportion of the wage spread between the highest and the lowest wage in the firm. The impact of both variables can be compared to assess if the ordinal rank

[11] The data reports a lower boundary of the average wage within the establishment because of the censoring. This problem can be neglected in our analysis because of the low frequency of censoring in the underlying data.

[12] 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.

[13] Experience is calculated with respect to the surveyed individual labor market entry. Hence, possible unemployment spells or apprenticeships directly following after school are accounted for.

[14] We do not include tenure in these estimates because the comparison group should also include comparable workers at later career stages (career prospects). Another reason is related to problems involving the degrees of freedom because we observe establishments with at least ten annual observations. Descriptive statistics are presented in Table A3.2 in the Appendix.

[15] 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.

[16] 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. See Pfeifer and Schneck (2010) for a slightly different calculation of rank. The following results, however, are robust to both calculations of rank.

(wage rank) or the cardinal rank (wage range) is more important to workers (Fields and Fei 1978, Brown et al. 2008). The individual wage rank indicates, in an ordinal sense, a worker's position in his firm's wage hierarchy. To analyze nonlinearity, the squared wage rank of workers is used. 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).

Table 3.2

Definition of wage measures

w_{ijt}^{abs}	Log wage of individual i in period t in establishment j	$w_{ijt}^{abs} = ln(wage_{ijt})$
$\bar{w}_{jt}$	Average log wage paid in establishment j in period t	$\bar{w}_{jt} = \frac{1}{N_{jt}} \sum_{i_{jt}}^{N_{jt}} w_{ijt}^{abs}$ with N_{jt}: number of employees in firm j in period t
$\hat{w}_{ijt}^{inside}$	Predicted comparison wage in own establishment (given individual characteristics) in period t	Annual regression for establishment j: $\hat{w}_{ijt}^{inside} = \hat{\alpha}_j + \hat{\iota}_j' X_{ijt}$ with X: experience (squared), dummies for occupation, and schooling
w_{ijt}^{rank}	Ordinal relative wage position of individual i in establishment j in period t	$w_{ijt}^{rank} = \frac{\text{wage rank}_{ijt} - 1}{\text{wage rank}_{jt}^{max} - 1}$ Workers with equal wages within establishment j have the same rank.*
w_{ijt}^{range}	Cardinal relative wage position of individual i in establishment j in period t	$w_{ijt}^{range} = \frac{w_{ijt} - w_{ijt}^{min}}{w_{jt}^{max} - w_{jt}^{min}}$
w_{ijt}^{CDF}	Empirical cumulative distribution function (CDF) of w_{ijt} in establishment j in period t	$w_{ijt}^{CDF} = Pr(W_{jt} \leq w_{ijt})$ with 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

* *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.*

Table 3.2 summarizes the definitions of our wage measures. Table A3.2 presents means and standard deviations of the constructed relative wage measures. As already noted by Brown et al. (2008, p. 372), the "different measures of pay are [...] somewhat correlated. Nevertheless, the large number of observations makes it possible, in practice, to estimate the separate variables' effects". As the measures also contain quite similar information, we only account for one measure of the relative wage position 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 because both measures differ especially at the tails of the distribution. Note that we do not say that workers actually know co-worker's wages: "All we can say is that people act as though they are able to form a reasonable estimate of where, as individuals, they lie in the pay ordering and the range" (Brown et al. 2008, p. 379).

3.3.3 Econometric models

Our basic estimation framework looks as in equation (3.6), in which Quit* denotes the latent individual quit probability, α the constant, β the coefficients of the relative wage position w_{ijt}^{r} for which we incorporate the different relative wage measures discussed in the previous section ($\bar{w}_{jt}$, $\hat{w}_{ijt}^{inside}$, w_{ijt}^{rank}, w_{ijt}^{range}, w_{ijt}^{CDF}), κ the coefficients for the absolute wage[17], γ the coefficients of worker characteristics X (schooling degree, tenure, squared tenure, experience, squared experience, professional status), δ the coefficients of firm characteristics Y (share of unskilled workers within the establishment, establishment size classes, firm-reported number of layoffs, firm-reported number of quits, works council, collective bargaining, sector, and federal state), λ time fixed effects, and ϵ the remaining residual term. For descriptive statistics of the variables see Table A3.2 in the Appendix.

$$Quit_{ijt}^{*} = \alpha + \beta' w_{ijt}^{r} + \kappa' w_{ijt}^{abs} + \gamma' X_{it} + \delta' Y_{jt} + \lambda_t + \epsilon_{ijt} \tag{3.6}$$

[17] Analogously to Galizzi and Lang (1998), we include information on the lagged absolute wage. This specification can be reparameterized in order to allow for either wage growth and present wage because the decision to quit is not expected to be a myopic one. Note that we utilize the individual wage information of the year 1995 as lagged wage information in year 1996.

As the quit probability cannot directly be observed, we concentrate on the actual quit behavior.[18] The dependent variable here is binary and that is why binary choice models are preferable. In addition, Table A3.1 shows that quits are a rare event and, therefore, linear models should not be applied. As the data are set up as a panel, effort can be made to control for unobserved heterogeneity. Note that individual fixed-effects probit models are inappropriate because these models are affected by the incidental parameter problem. This problem decreases when the individual observations increase (see Heckman 1981) but, here, the panel is too short with an average individual panel length of 3.5 years. For these reasons, the individual random-effects probit model is utilized here in which Φ is the cumulative density function of the standard normal distribution and ν_i is the individual random effect. Our explanatory variable of interest is the discussed measure of a worker's relative wage position within his firm. To ensure that the relative wage variable does not reflect the absolute wage, the individual's absolute wage also is included in addition to the individual's relative wage position. Because of our interest in average comparison wages within firms, which do not vary across workers in one firm, we do not control for firm fixed effects.

$$Quit_{ijt} = \begin{cases} 1 & \text{if worker } i \text{ quits his job in firm } j \text{ in period } t \\ 0 & \text{if worker } i \text{ stays in firm } j \text{ in period } t \end{cases} \tag{3.7}$$

$$Pr(Quit_{ijt} = 1) = \Phi(\alpha + \beta' w^{r}_{ijt} + \kappa' w^{abs}_{ijt} + \gamma' X_{it} + \delta' Y_{jt} + \lambda_t + \nu_i) \tag{3.8}$$

β is a scalar when testing hypotheses 1a and 1b because the main interest is on linear effects of relative wage positions while β becomes a vector by consideration of hypothesis 2 because we test the hypothesis of a U-shaped relationship between rank and quit probability by application of w^{rank}_{ijt} and squared w^{rank}_{ijt}. Evidence in favor of hypothesis 2 is provided in case of a negative estimate for the coefficient of w^{rank}_{ijt} in combination with a positive estimate for squared w^{rank}_{ijt}.

In addition to the determinants of individual quit behavior, we also analyze the conse-

[18] Each worker in the consecutive empirical investigation is observed in at least two consecutive periods. Calculation of the wage measures also includes individuals who are observed only once.

quences 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). 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 3.2, utility and quit probability 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 absolute 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.

In addition, it is possible to regress the absolute wage difference on the difference in relative wage positions to assess possible trade-offs in the utility function. Equation (3.9) presents the estimation framework, in which $(w^{new}\text{-}\ w^{old})_{it}$ is the difference of individual absolute wages between the new and the previous 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 previous firm, γ the coefficients of worker characteristics X (schooling degree, change in establishment size class, experience, squared experience), λ time fixed effects, and ϵ the remaining residual term. This estimation framework further allows us to investigate which socio-demographic groups (e.g., workers without completed vocational training) are more affected by wage cuts.

$$(w^{new} - w^{old})_{it} = \alpha + \eta(w^{r,new} - w^{r,old})_{it} + \gamma' X_{it} + \lambda_t + \epsilon_{it} \tag{3.9}$$

If workers accept lower absolute wages in the new firm because they are compensated with higher status, that is, higher relative wage positions, we would expect the coefficients η's to be negative. 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 (3.10). Equation (3.11) displays that the dependent variable takes the value one in case of a wage cut and zero otherwise.

$$Pr(\text{wage cut}_{it} = 1) = \Phi(\alpha + \eta(w^{r,new} - w^{r,old})_{it} + \gamma' X_{it} + \lambda_t) \tag{3.10}$$

$$\text{wage cut}_{it} = \begin{cases} 1 & (w^{new} - w^{old})_{it} < 0 \\ 0 & (w^{new} - w^{old})_{it} \geq 0 \end{cases} \tag{3.11}$$

3.4 Results

This section is to reveal which of the hypotheses derived above meet the data. At first, we discuss the results on whether a linear status effects dominates a possible linear signal effect or whether it is vice versa. In a next step, we present estimates related to possible nonlinearities. Finally, results on the consequences of quits are addressed.

In the following, we present the results of the individual random effects probit model as discussed in equation (3.8) in section 3.3.3. Likelihood-ratio tests reject the null hypothesis of no individual unobserved heterogeneity in all specifications (see Table A3.3), which indicates that the random effects probit model is more appropriate compared to a pooled probit model. Note that we primarily discuss marginal effects of our wage variables at the means of all covariates and under the assumption that the individual error term is zero. The estimated absolute marginal effects which are presented in Table 3.3 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 sizable.[19] The complete estimation output and the corresponding coefficients are presented in Table A3.3 in the Appendix.

Before presenting the estimation results of the relative wage measures, we discuss the estimated effects of the absolute wage shown in Table A3.3. The first specification in Table

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

A3.3 only focuses on the effect of the absolute wage. We can conclude that individuals are somewhat myopic when deciding to leave their employer. To be more precise, the current absolute wage exhibits significant impact on the decision to quit while the lagged absolute wage is found to be insignificant and considerably smaller in size. This finding is confirmed when we include the measures for the relative wage position. At first glance, this result is counter-intuitive as we would expect that a worker's utility depends positively on his current wage. For this reason, a higher wage should decrease the individual quit propensity. Galizzi and Lang (1998) provide an explanation for this result and suggest that better paid workers might have better outside job opportunities because differences in wages might reflect some degree of unobserved productivity differences.

We now turn the focus on the relative wage measures. The first specification in Table 3.3 presents the marginal effect of the average within-establishment wage as relative wage measure which is obtained via specification (2) in Table A3.3. The marginal effect indeed seems to be quite small but, however, the relative marginal effect is quite sizable. The marginal effect presented below can be interpreted in the way that an increase of the average wage within a firm j at time t of one log unit decreases the quit probability by 0.0000439 percentage points or by 12.30%, respectively. Note that the coefficient is not statistically significant which does not allow for distinct conclusions about hypothesis 1. From specification (2) in Table 3.3, we can learn that the coefficient of $\hat{w}_{ijt}^{inside}$ is also statistically insignificant. In brief, the presented coefficients in specifications (1) and (2) of Table 3.3 are not adequate to provide a distinct answer in favor or against hypotheses 1a and 1b.

Specifications (3) to (5) in Table 3.3 present the results for the rank measure, the range, and CDF which are obtained via specifications (4) to (6) in Table A3.3. The corresponding coefficients are shown to be significant but, however, the estimated coefficients advert to different effects because the effects of w_{ijt}^{rank} and w_{ijt}^{CDF} are positive while the coefficient for w_{ijt}^{range} is negative. This implies that the estimated effects for w_{ijt}^{rank} and w_{ijt}^{CDF} support hypothesis 1b. In other words, workers in higher relative wage positions quit their jobs more likely compared with workers in lower relative wage positions because of the lack of future career prospects. The results for the range measure, however, are reverse because the lower the worker's position, the higher the probability to change employers. Accordingly, this implies that the status effect dominates the signal effect which is in line with hypothesis 1a. To conclude, the results in Table 3.3 do not clearly support hypothesis 1a or hypothesis 1b because it depends on the definition of the measure for the relative wage position.

Table 3.3

Random-effects probit results for quit probability

	$\bar{x}$	(1)	(2)	(3)	(4)	(5)
$\bar{w}_{jt}$	4.6921	-0.0000439				
		(0.00007)				
		[-0.1230]				
$\hat{w}_{ijt}^{inside}$	4.6987		0.0000459			
			(0.00008)			
			[0.1286]			
w_{ijt}^{rank}	0.5114			0.0000946**		
				(0.00004)		
				[0.2657]		
w_{ijt}^{range}	0.6500				-0.000106*	
					(0.00005)	
					[-0.2969]	
w_{ijt}^{CDF}	0.5212					0.0000873**
						(0.00004)
						[0.2452]
Pr(quit=1| $\bar{x}$)		0.000357	0.000357	0.000356	0.000357	0.000356
Number of observations				2,902,724		
Number of individuals				833,359		

Note: Random-effects probit estimation. The table presents marginal effects at $\bar{x}$, standard errors in parentheses, and relative marginal effects in brackets.

**** p<0.01, ** p<0.05, * p<0.1.*

The corresponding coefficients of the probit estimates and the complete results are presented in Table A3.3. Table A3.2 contains descriptive statistics.

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

The individual-specific and establishment-specific control variables which are included into the above specifications are robust across the different specifications. The higher the educational level, the higher the probability to quit a job. This effect is in line with expectations because highly productive employees are more likely to find a new job. The effect of individual tenure on the decision to quit is U-shaped which reveals that individuals are most likely to change jobs very early (or after a very long tenure in the same firm). The effect of general human capital (experience) is not significant. The effects of the firm characteristics are also highly robust across specifications.

When turning the focus on an examination of hypothesis 2 and possible nonlinearities, we estimate the individual random-effects probit model as discussed in equation (3.8) in section 3.3.3 and include a linear as well as quadratic term of w_{ijt}^{rank}. Significant coefficients for either the linear and the quadratic term for the relative wage, then, suggest the presence of an (inverted) U-shape relationship regarding quits and relative wage positions. Evidence in favor of hypothesis 2 can be found if the estimated coefficient for w_{ijt}^{rank} is negative while w_{ijt}^{rank} squared is estimated to be positive.

Figure 3.1 summarizes the predicted quit probabilities for an average worker (at $\bar{x}$) as a function of of w_{ijt}^{rank} which are obtained from the linear specification (see specification (4) in Table A3.3) and the quadratic specification (see specification (1) in Table A3.4). The linear specification reveals a linear increasing quit probability for the average worker. Starting from 0.032% in lowest relative wage positions, it rises until about 0.041% in the highest ranks. Figure 3.1 illustrates that workers in higher relative wage positions are more likely to change employers voluntarily. In addition, the predicted quit probability of the quadratic specification is presented. The U-shape in Figure 3.1 directly supports hypothesis 2. Specification (1) in Table A3.4 presents the results in detail and shows that both coefficients of interest are highly significant and have the expected signs which confirms our expectations derived in the theoretical framework. This effect is interpreted in the way that workers in low relative positions quit their jobs because of concerns regarding their low status. Workers in high relative wage positions, in turn, are suggested to be more likely to change jobs because of few career advancement opportunities within the firm. An interpretation which is closer to the theoretical part above is that workers change employers more likely if utility is obtained purely by status or solely by signal in comparison to workers who are able to add up utility from both, signal as well as status.

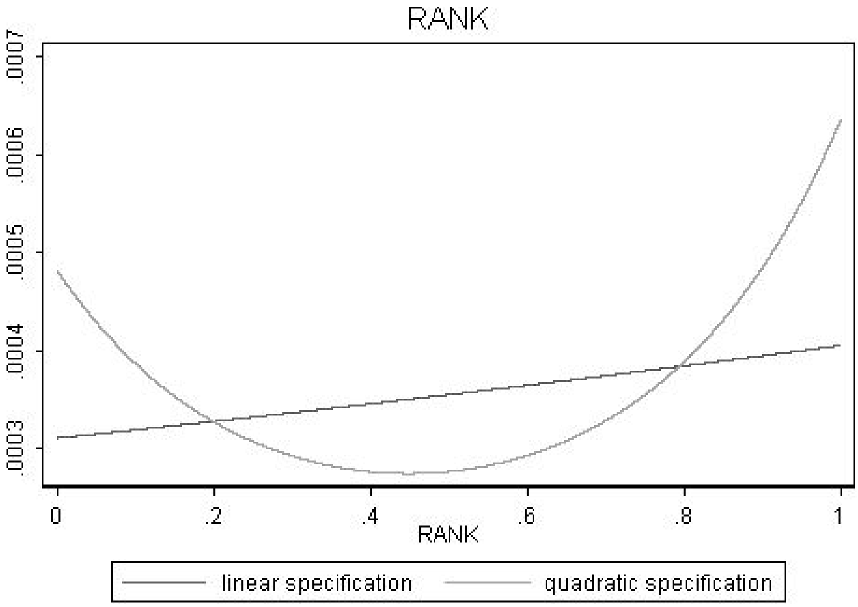

Figure 3.1: Predicted quit probabilities and relative wage positions (at $\bar{x}$) for w_{ijt}^{rank}

We conducted several robustness checks in order to check the sensitivity of the nonlinear effect. At first, we checked the effect of a cubic term in specification (2) of Table A3.4. Again, the linear and the quadratic coefficient advert to a U-shaped relationship while the cubic term is found to be insignificant. This suggests that the quadratic specification is satisfactory. Specification (3) in Table A3.4 considers the relative standing as dummy variables which indicate the relative wage position of workers within a firm's wage distribution in each year. As reference category, we decided to apply the middle wage quintile because this allows for direct evidence regarding a possible U-shape. Precisely, if the coefficients for the remaining wage quintiles are positive, then workers in the third wage quintile have the lowest probability to quit. The results show that mobility is significantly more common in the first and the fifth quintile, compared to workers in the third quintile. This directly illustrates a U-shaped quit probability because workers at the left and at the right margin of the firm's wage distribution are significantly more likely to quit jobs in comparison to workers in the middle of the firm's wage distribution. Also the workers in the fourth wage quintile are shown to be more likely to quit when compared with workers in the third quintile. Note that workers in the second wage quintile are less likely to change jobs voluntarily compared with workers in the middle wage quintile whereas this effect is not statistically significant. In sum, the robustness checks shown in Table A3.4 confirm the presence of a U-shape relationship of the rank in quit decisions.

Table A3.5 displays further robustness checks for the nonlinear effects. Application of w_{ijt}^{range} in specification (1) does not support the results obtained for w_{ijt}^{rank}. In fact, an inverse U-shape relationship is estimated but, however, the coefficients are insignificant. As a result, application of w_{ijt}^{range} does not allow for distinct conclusions against or in favor of the theoretical expectations. The quadratic specification of w_{ijt}^{CDF} in specification (2) of Table A3.5 confirms the significant U-shape as shown in Table A3.4, Figure 3.1, and the theoretical part. The graph (see Figure A3.1 in the appendix) generally confirms the relationship as shown in Figure 3.1 but reveals some differences in the levels of the predicted quit probability of an average worker as a function of w_{ijt}^{CDF}. To sum up, the results basically are in line with a U-shaped relationship and are robust to several relative wage measures (w_{ijt}^{rank}, w_{ijt}^{CDF}, and within-firm wage quintiles) with exception of w_{ijt}^{range}.

In a last step, we turn the focus on hypothesis 3 where we analyze the consequences of quits with respect to absolute wages and relative wage positions within a firm. More precisely, we are interested in trade-off reasoning between absolute wages and relative wage positions after a voluntary job change and the question whether mobile workers are compensated for wage cuts by an increase in their relative wage positions, that is, by a gain in status. Previous studies have found most voluntary job mobility to be associated with wage gains but also that a substantial share of quits is accompanied with wage cuts. When only focusing on mobile workers, 42.86% of individuals experience a wage cut when changing the firm (see Table A3.6). This number exceeds the size of previous studies for Germany (see Jolivet et al. 2006, Fitzenberger and Garloff 2007).

Table 3.4

Descriptive statistics for consequences of quits

		Mean	Std. Deviation
$(w^{new} - w^{old})_{it}$	all	0.0170	0.1426
	wage markup	0.0934	0.1031
	wage cut	-0.0847	0.1230
$(rank^{new} - rank^{old})_{it}$	all	-0.0244	0.2390
	wage markup	0.0576	0.2145
	wage cut	-0.1336	0.2261
$(range^{new} - range^{old})_{it}$	all	0.0109	0.1977
	wage markup	0.0597	0.1934
	wage cut	-0.0542	0.1843
$(CDF^{new} - CDF^{old})_{it}$	all	-0.0203	0.2386
	wage markup	0.0648	0.2147
	wage cut	-0.1338	0.2210

Note: 5,464 transitions of 5,280 individuals are observed. 3,122 moves are executed to higher wages and 2,342 moves to lower wages (wage cut).

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table 3.4 presents descriptive statistics about changes in absolute wages $(w^{new} - w^{old})_{it}$ and relative wage positions within the new and the old firm $((w_{ijt}^{rank,new} - w_{ijt}^{rank,old})_{it}$, $(w_{ijt}^{range,new} - w_{ijt}^{range,old})_{it}$, and $(w_{ijt}^{CDF,new} - w_{ijt}^{CDF,old})_{it})$. Workers gain, on average, 0.0170 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, 0.0847 log points lower wages, while workers with a wage markup receive an average of 0.0934 log points higher wages. For the wage position 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 compared to the previous one. The cardinal wage range measure (w_{ijt}^{range}), in turn, is slightly positive and indicates that workers change to higher relative wage positions. One might be tended to misleadingly conclude that the average gain in absolute wages and the loss in ordinal relative wage positions directly supports Hypotheses 3a which states that workers trade off absolute wages and relative wage positions when changing firms (status effect). If we look at workers with wage cuts, we see that those workers also suffer 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 might have more chances for career advancements (signal effect).

In the following, we present Epanechnikov kernel density estimators for changes in relative wage positions to shed some more light into the heterogeneous consequences of quits. Figure 3.2 distinguishes between mobile workers with wage cuts and wage markups in order to visualize the changes in relative wage positions by wage change. It displays the distributions of changes in the wage positions and in the relative wage positions. All the results for the different measures of relative wage positions are quite similar. Most workers who suffer lower wages additionally lose in relative wage positions. In contradiction, most mobile workers with wage markups do not experience losses in their relative wage positions. Their relative wage positions, however, remain relatively stable. Nevertheless, it can be seen that a somewhat larger fraction of workers with wage markups rather gain than lose with respect to relative wage positions. Only few workers seem to accept wage cuts in order to improve their relative wage positions and to gain additional status by inspection of Figure 3.2.

The kernel density estimators shown below do not account for further determinants of consequences of quits such as educational level or labor market experience. Thus, we use linear regressions in order to regress changes in absolute wages on changes in relative wage positions and a set of control variables (see equation (3.9) for the econometric model and Table A3.6 in the Appendix for descriptive statistics).

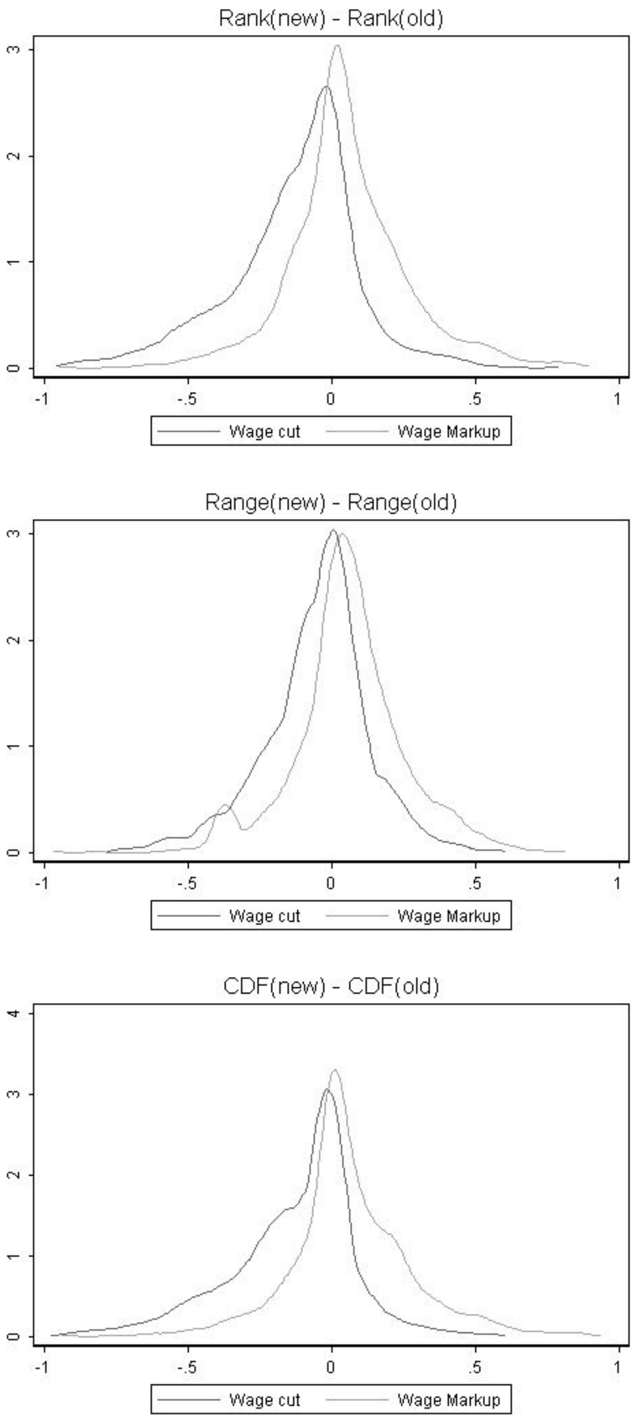

Figure 3.2: Kernel density estimator for changes in relative wage positions

Table 3.5

Linear regression results for wage change

$(w^{new} - w^{old})_{it}$	(1)	(2)	(3)
$(rank^{new} - rank^{old})_{it}$	0.302***		
	(0.0114)		
$(range^{new} - range^{old})_{it}$		0.340***	
		(0.0145)	
$(CDF^{new} - CDF^{old})_{it}$			0.309***
			(0.0114)
Mobility to larger establishment	0.0358***	0.00691	0.0368***
(Dummy variable)	(0.00540)	(0.00496)	(0.00534)
Secondary school leaving certificate		reference	
Secondary school leaving certificate	0.0174**	0.00787	0.0160**
and apprenticeship	(0.00718)	(0.00764)	(0.00718)
(Technical) college entrance qualification	0.0117	-0.00331	0.00949
	(0.0124)	(0.0117)	(0.0120)
(Technical) college entrance qualification	0.0255***	0.0143*	0.0197**
and apprenticeship	(0.00815)	(0.00850)	(0.00813)
Experience	-0.00460***	-0.00385***	-0.00445***
	(0.00139)	(0.00144)	(0.00138)
Experience^2	0.0000662	0.0000527	0.0000605
	(0.0000405)	(0.0000420)	(0.0000404)
Firm-reported quits	0.0000557**	0.000108***	0.0000617**
	(0.0000279)	(0.0000287)	(0.0000278)
Firm-reported layoffs	0.0000499	0.000446***	0.0000596
	(0.0000762)	(0.0000821)	(0.0000772)
Constant	0.0354***	0.0175	0.0389***
	(0.0138)	(0.0148)	(0.0137)
Annual Dummy variables		included	
R^2	0.2981	0.2575	0.3091
Number of observations		5,464	

Note: OLS coefficients are presented. Robust standard errors clustered for 5,280 individuals in parentheses.

**** p<0.01, ** p<0.05, * p<0.10.*

Table A3.6 contains descriptive statistics.

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

The results in Table 3.5 support our previous findings that changes in absolute wages and changes in relative wage positions are positively correlated. This implies that workers who are able to improve their relative wage position when changing establishments are more likely to achieve higher wages. We, further, estimate probit models for the determinants of accepting wage cuts (see equation (3.10)). Table 3.6 presents the estimated marginal effects at $\bar{x}$. The results show that workers who improve their relative wage positions are less likely to experience a wage cut. Note that the effects of the control variables vary somewhat across the different specifications and estimation approaches. The effect of education, for example, is significant in the linear regression while it is insignificant in the probit model on wage cuts.

The results in Tables 3.5 and 3.6 also shed light on the question which groups are more likely to suffer wage cuts. In the linear models on the wage change, specifications (1) and (3) suggest that workers who are mobile to larger establishments[20] gain by mobility. Also in the probit model, specifications (1) and (3) corroborate that individuals who change to larger establishments are significantly less likely to suffer 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). In addition, the linear regression results as well as the probit estimates suggest that in times of increasing quits, workers are more likely to achieve wage markups. This might be explained by poaching strategies where firms try to attract highly skilled and experienced workers via higher wages.

[20]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 than the previous firm. The establishment size classes can be found in Table A3.2 in the Appendix.

Table 3.6

Probit estimation results for wage cuts

$wage\ cut_{it}$	(1)	(2)	(3)
$(rank^{new} - rank^{old})_{it}$	-1.0294***		
	(0.0414)		
$(range^{new} - range^{old})_{it}$		-0.805***	
		(0.0402)	
$(CDF^{new} - CDF^{old})_{it}$			-1.104***
			(0.0429)
Mobility to larger establishment	-0.0999***	-0.0178	-0.1047***
(Dummy variable)	(0.0199)	(0.0190)	(0.0199)
Secondary school leaving certificate		reference	
Secondary school leaving certificate	-0.0483	-0.0264	-0.0450
and apprenticeship	(0.0301)	(0.0304)	(0.0304)
(Technical) college entrance qualification	0.0668	0.0896	0.0738
	(0.0741)	(0.0712)	(0.0742)
(Technical) college entrance qualification	-0.0565	-0.0414	-0.0379
and apprenticeship	(0.0347)	(0.0351)	(0.0355)
Experience	0.00887	0.00754	0.00872
	(0.00542)	(0.00521)	(0.00547)
Experience^2	0.0000251	0.0000210	0.0000398
	(0.000158)	(0.000153)	(0.000160)
Firm-reported quits	-0.000278*	-0.000393***	-0.000300**
	(0.000146)	(0.000147)	(0.000146)
Firm-reported layoffs	0.000501	-0.000591	0.000501
	(0.000384)	(0.000373)	(0.000391)
Annual Dummy variables		included	
Pseudo R^2	0.1704	0.1029	0.1840
Number of observations		5,464	

Note: Marginal effects at $\bar{x}$ are presented. Robust standard errors clustered for 5,280 individuals in parentheses.

**** p<0.01, ** p<0.05, * p<0.10.*

Table A3.6 contains descriptive statistics.

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

The data do not allow for distinct definitions regarding voluntary mobility. For this reason, we conduct a robustness check, where we relax the assumption that individuals have to change jobs within eight days. We, now, refer to job-to-job mobility within one year and relax the assumption of mobility with virtually no intervening unemployment.[21] Application of this definition for mobility reveals 6,835 transitions of 6,403 individuals who are mobile up to 4 times. Tables A3.7 and A3.8 present the main results. Specification (1) in Table A3.7 reveals that the coefficient of $\bar{w}_{jt}$ is positive. This particular effect is shown to be non-robust when compared to the corresponding effect presented in Tables 3.3 and A3.3 because of different signs. The effects, however, are statistically insignificant in all the discussed tables which suggests that the effects are imprecisely measured. An increase of $\hat{w}_{ijt}^{inside}$ increases the quit probability, confirming the results presented in Tables 3.3 and A3.3. Table A3.7, further, shows that the main results for w_{ijt}^{rank} and w_{ijt}^{CDF} hold because the effects are comparable to the ones shown above (see specifications (3), (4), (7), and (8)). Similar to the results for mobile individuals within eight days, a positive coefficient in the linear model and a U-shaped relation in the quadratic model, respectively, are indicated. The effect of w_{ijt}^{range} is robust in the linear case (see specification (5)). Specification (6) in Table A3.7 shows that a quadratic relationship cannot be confirmed. Precisely, the results suggest that increasing w_{ijt}^{range} decreases the quit probability linearly because of the insignificant effect for w_{ijt}^{range} squared. Table A3.8 refers to the consequences of mobility and applies the definition of job-to-job mobility within one year. Similar to Table 3.5, the linear models in Table A3.8 reveal that transitions to higher relative wage positions, on average, lead to wage markups (see specifications (1) to (3)). Specifications (4) to (6) refer to the probit models and show that mobility to a higher relative wage position decreases the probability for wage cuts. The effects, thus, are robust to the ones presented for mobility between two establishments within eight days.

In sum, our analysis does not provide evidence that wage cuts are accepted in exchange for an increase in status which is associated with higher relative wage positions. Workers who suffer decreasing relative wage positions are, in fact, also more likely to suffer lower wages when changing establishments and vice versa. On the one hand, this finding can

[21] The only requirement for individual mobility is a change in the establishment identifier in consecutive periods.

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 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.

3.5 Conclusions

Our main results are that relative wage positions have a significant impact on the probability to voluntarily quit a job. For w_{ijt}^{rank} and w_{ijt}^{CDF}, 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, in turn, have 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. In contradiction, w_{ijt}^{range} suggests that the reverse status effect is dominant which implies that workers quit their jobs more often because of low own status. In brief, the presented results are mixed. For this reason it is hard to draw distinct conclusions about whether the status effect dominates the signal effect in quit decisions or whether it is vice versa.

More detailed insights into the relationship between relative wage positions and the quit behavior can be derived by examination of nonlinear effects. This enhancement to the literature basically reveals U-shaped relation between the individual rank and the decision to quit. This finding is consistent with the expectations outlined in the theoretical part of this chapter. We also come up with an interpretation of this effect. Specifically, we interpret this U-shaped relation in the way that workers at the bottom of the within-establishment pay scale are more sensitive to status considerations while those at the top are suggested to be more responsive to signal considerations. In other words, workers in low relative wage positions seem to care more about their low status than about future career advancement opportunities while for individuals in high relative wage positions the reverse is true because these workers seem to be more sensitive to their low future career prospects than to their high status. The present study of nonlinear effects, thus, provides detailed insights into the complex relationship between comparison income and economic behavior.

In the last part of the analysis, we find that relative wage positions significantly affect the probability to accept a wage cut when changing firms. Workers who are able to improve their relative wage position are, on average, less likely to pay for mobility by lower wages. Workers changing to lower relative wage positions, in turn, are found to be confronted with an additional loss in earnings. This is in line with the hypothesis that across mobile individuals lower absolute wages and lower relative wage positions go hand in hand. From the theory about signal effects, we suggest that workers accept short-term wage cuts and lower relative wage positions in order to climb up the career ladder at the new employer in the long run.

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. 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 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).

3.6 Appendix

Table A3.1

Quits by annual observations per establishment

	Annual observations per establishment			
	$N \geq 10$	$N \geq 15$	$N \geq 50$	$N \geq 100$
Number of quits	5,464	5,434	5,198	4,791
Number of observations	2,902,724	2,897,110	2,812,568	2,671,678

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.2
Descriptive statistics

Variable	Mean	Standard Deviation
Quit = 1	0.0019	0.0433
w_{ijt}^{abs}	4.7017	0.2288
w_{ijt-1}^{abs}	4.6893	0.2277
within-firm wage quintile 1	0.1788	0.3832
within-firm wage quintile 2	0.2041	0.4031
within-firm wage quintile 3	0.2127	0.4092
within-firm wage quintile 4	0.2167	0.4120
within-firm wage quintile 5	0.1877	0.3904
$\bar{w}_{jt}$	4.6921	0.1416
$\hat{w}_{ijt}^{inside}$	4.6987	0.1936
w_{ijt}^{rank}	0.5114	0.2796
w_{ijt}^{rank} squared	0.3398	0.2912
w_{ijt}^{range}	0.6500	0.2280
w_{ijt}^{CDF}	0.5212	0.2901
Tenure	13.1255	7.4479
Tenure^2	227.7509	217.0195
Experience	18.4650	6.6272
Experience^2	384.8737	233.8787
Professional status (dummies)		
Unskilled worker^{i}	0.2777	0.4478
Skilled worker/ Craftsman^{ii}	0.3561	0.4788
Technician/ Foreman^{iii}	0.0310	0.1733
Clerk^{iv}	0.3353	0.4721
Highest schooling degree (dummies)		
Secondary school leaving certificate^{v}	0.1200	0.3250
Secondary school leaving certificate and apprenticeship^{vi}	0.8169	0.3867
(Technical) college entrance qualification^{vii}	0.0088	0.0932
(Technical) college entrance qualification and apprenticeship^{viii}	0.0543	0.2265
Establishment size class (dummies)		
Workforce of establishment in [10;49]	0.0039	0.062
Workforce of establishment in [50;199]	0.0445	0.2062
Workforce of establishment in [200;999]	0.2475	0.4315
Workforce of establishment in $\geq$ 1000	0.7041	0.4564
Share of unqualified workers within the establishment	0.2719	0.2455
Works council	0.9793	0.1425
Collective bargaining	0.9627	0.1895
Firm-reported quits	20.077	38.7673
Firm-reported layoffs	7.5058	16.3732
Sector (dummies)		
Agriculture	0.0000	0.0000
Mining	0.0502	0.2184
Building	0.0108	0.1034
Credit	0.0611	0.2395
Traffic	0.0513	0.2205
Retail	0.0235	0.1516

Continued on next page

Table A3.2 (continued)

Variable	Mean	Standard Deviation
Hotel	0.0051	0.0713
Education	0.0118	0.1080
Service	0.0254	0.1574
Welfare	0.0262	0.1597
Public utility	0.0570	0.2319
Production	0.6775	0.4674
Federal region (dummies)		
Schleswig Holstein	0.0250	0.1562
Hamburg	0.0723	0.2589
Lower Saxony[ix]	0.1243	0.3299
Bremen	0.0181	0.1331
North Rhine-Westphalia[x]	0.2931	0.4552
Hesse (Hessen)	0.0811	0.2730
Baden-Wuerttemberg	0.1324	0.3389
Bavaria[xi]	0.1593	0.3660
Rhineland-Palatinate and Saarland[xii]	0.0944	0.2924
annual dummy variables		
1996	0.1225	0.3278
1997	0.1059	0.3078
1998	0.0911	0.2877
1999	0.0931	0.2905
2000	0.0873	0.2823
2001	0.1007	0.3009
2002	0.1013	0.3018
2003	0.0994	0.2993
2004	0.1011	0.3014
2005	0.0976	0.2967
Number of observations	2,902,724	
Number of individuals	833,359	
Number of establishments	4,260	

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.

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.3
Random-effects probit results for quit probability

Quit=1	(1)	(2)	(3)	(4)	(5)	(6)
$\tilde{w}_{jt}$		-0.0338				
		(0.0533)				
$\hat{w}_{ijt}^{inside}$			0.0353			
			(0.0583)			
w_{ijt}^{rank}				0.0729**		
				(0.0293)		
w_{ijt}^{range}					-0.0814**	
					(0.0414)	
w_{ijt}^{CDF}						0.0672**
						(0.0289)
w_{ijt}^{abs}	0.204***	0.2136***	0.192***	0.132*	0.267***	0.136*
	(0.0716)	(0.0731)	(0.0744)	(0.0770)	(0.0787)	(0.0770)
w_{ijt-1}^{abs}	0.0163	0.0173	0.0129	0.0193	0.0211	0.0157
	(0.0709)	(0.0709)	(0.0710)	(0.0707)	(0 .0712)	(0.0707)
Tenure	-0.0370***	-0.0370***	-0.0372***	-0.0372***	-0.0370***	-0.0371***
	(0.00274)	(0.00274)	(0.00275)	(0.00275)	(0.00274)	(0.00274)
Tenure^2	0.000585***	0.000586***	0.000588***	0.000590***	0.000585***	0.000587***
	(0.000104)	(0.000104)	(0.000104)	(0.000104)	(0.000104)	(0.000104)
Experience	-0.00223	-0.00236	-0.00261	-0.00275	-0.00211	-0.00276
	(0.00387)	(0.00388)	(0.00392)	(0.00388)	(0.00387)	(0.00388)
Experience^2	-0.0000860	-0.0000837	-0.0000780	-0.0000765	-0.0000884	-0.0000767
	(0.000114)	(0.000114)	(0.000115)	(0.000114)	(0.000114)	(0.000114)
Unskilled worker				reference		
Skilled worker/ Craftsman	-0.0254*	-0.0258*	-0.0265*	-0.0283*	-0.0254*	-0.0277*
	(0.0149)	(0.0149)	(0.0150)	(0.0150)	(0.0149)	(0.0150)
Technician/ Foreman	-0.0895**	-0.0914**	-0.0953***	-0.0983***	-0.0882**	-0.0971***
	(0.0359)	(0.0360)*	(0.0371)	(0.0360)	(0.0359)	(0.0360)
Clerk	0.0444***	0.0432**	0.0395**	0.0395**	0.0452***	0.0396**
	(0.0170)	(0.0171)	(0.0188)	(0.0171)	(0.0170)	(0.0171)
Secondary school leaving certificate				reference		
Secondary school leaving certificate and apprenticeship	0.128***	0.128***	0.126***	0.130***	0.128***	0.130***
	(0.0207)	(0.0207)	(0.0209)	(0.0207)	(0.0207)	(0.0207)
(Technical) college entrance qualification	0.0905*	0.0913*	0.0881*	0.0939*	0.0910*	0.0920*
	(0.0518)	(0.0518)	(0.0519)*	(0.0518)	(0.0518)	(0.0518)
(Technical) college entrance qualification and apprenticeship	0.262***	0.262***	0.259***	0.264***	0.262***	0.264***
	(0.0275)	(0.0275)	(0.0279)	(0.0275)	(0.0275)	(0.0275)
Workforce of establishment in [10;49]				reference		
Workforce of establishment in [50;199]	0.0550	0.0560	0.0545	0.0583	0.0544	0.0597
	(0.0661)	(0.0661)	(0.0661)	(0.0661)	(0.0661)	(0.0661)
Workforce of establishment in [200;999]	-0.0665	-0.0647	-0.0675	-0.0603	-0.0647	-0.0592
	(0.0652)	(0.0653)	(0.0653)	(0.0653)	(0.0653)	(0.0653)
Workforce of establishment in $\geq$ 1000	-0.174***	-0.172***	-0.176***	-0.164**	-0.169***	-0.164**
	(0.0656)	(0.0658)	(0.0657)	(0.0657)	(0.0657)	(0.0658)
Share of unqualified workers within the establishment	-0.139***	-0.143***	-0.137***	-0.154***	-0.136***	-0.152***
	(0.0247)	(0.0258)	(0.0248)	(0.0255)	(0.0247)	(0.0253)
Works council	-0.147***	-0.145***	-0.149***	-0.139***	-0.152***	-0.140***
	(0.0318)	(0.0321)	(0.0319)	(0.0320)	(0.0319)	(0.0320)
Collective bargaining	0.245***	0.245***	0.245***	0.245***	0.245***	0.245***
	(0.0312)	(0.0312)	(0.0312)	(0.0312)	(0.0312)	(0.0312)
Firm-reported quits	0.00163***	0.00164***	0.00162***	0.00166***	0.00161***	0.00165***
	(0.000110)	(0.000111)	(0.000111)	(0.000111)	(0.000110)	(0.000111)
Firm-reported layoffs	0.0000672	0.0000747	0.0000616	0.0000864	0.000111	0.0000875
	(0.000315)	(0.000315)	(0.000315)	(0.000314)	(0.000315)	(0.000314)
Sector dummy variables				included		
Regional dummy variables				included		
Annual dummy variables				included		
Constant	-4.107***	-4.000***	-4.187***	-3.824***	-4.374***	-3.827***
	(0.150)	(0.226)	(0.201)	(0.188)	(0.203)	(0.192)
LR test of rho=0 (p-value)	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001
Log-Likelihood	-37269.05	-37268.85	-37268.86	-37265.94	-37267.12	-37266.34
Number of observations				2,902,724		
Number of individuals				833,359		
Number of establishments				4,260		

Note: Random-effects probit (coefficients). Standard errors in parentheses. 'LR' denotes likelihood-ratio.
**** p<0.01, ** p<0.05, * p<0.1.*
Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.4
Random-effects probit results for quit probability

Quit=1	(1)	(2)	(3)
w_{ijt}^{rank}	-0.687***	-0.709***	
	(0.0743)	(0.182)	
w_{ijt}^{rank} squared	0.767***	0.820**	
	(0.0692)	(0.404)	
w_{ijt}^{rank} cubic		-0.0346	
		(0.261)	
within-firm wage quintile 1			0.0645***
			(0.0184)
within-firm wage quintile 2			-0.00319
			(0.0168)
within-firm wage quintile 3			reference
within-firm wage quintile 4			0.0293*
			(0.0165)
within-firm wage quintile 5			0.165***
			(0.0187)
w_{ijt}^{abs}	0.145*	0.146*	0.115
	(0.0761)	(0.0766)	(0.0754)
w_{ijt-1}^{abs}	-0.00464	-0.00479	0.00551
	(0.0699)	(0.0699)	(0.0701)
Tenure	-0.0354***	-0.0354***	-0.0360***
	(0.00276)	(0.00276)	(0.00275)
Tenure^2	0.000530***	0.000530***	0.000545***
	(0.000105)	(0.000105)	(0.000105)
Experience	-0.000629	-0.000633	-0.00123
	(0.00389)	(0.00389)	(0.00388)
Experience^2	-0.000146	-0.000146	-0.000128
	(0.000114)	(0.000114)	(0.000114)
Unskilled worker		reference	
Skilled worker/ Craftsman	-0.0147	-0.0147	-0.0192
	(0.0151)	(0.0151)	(0.0150)
Technician/ Foreman	-0.130***	-0.130***	-0.133***
	(0.0362)	(0.0362)	(0.0362)
Clerk	0.0240	0.0239	0.0228
	(0.0172)	(0.0172)	(0.0172)
Secondary school leaving certificate		reference	
Secondary school leaving certificate	0.133***	0.133***	0.132***
and apprenticeship	(0.0208)	(0.0208)	(0.0207)
(Technical) college entrance qualification	0.100*	0.100*	0.101**
	(0.0519)	(0.0519)	(0.0519)
(Technical) college entrance qualification	0.267***	0.267***	0.266***
and apprenticeship	(0.0276)	(0.0276)	(0.0275)
Workforce of establishment in [10;49]		reference	
Workforce of establishment in [50;199]	0.0640	0.0640	0.0605
	(0.0663)	(0.0663)	(0.0662)
Workforce of establishment in [200;999]	-0.0512	-0.0512	-0.0586
	(0.0655)	(0.0655)	(0.0654)
Workforce of establishment in $\geq$ 1000	-0.154**	-0.154*	-0.164**
	(0.0659)	(0.0659)	(0.0658)
Share of unqualified workers within the	-0.162***	-0.162***	-0.167***
establishment	(0.0255)	(0.0255)	(0.0254)
Works council	-0.136***	-0.136***	-0.1345137***
	(0.0321)	(0.0321)	(0.0320257)
Collective bargaining	0.244***	0.244***	0.2415764**
	(0.0313)	(0.0313)	(0.0312386)
Firm-reported quits	0.00171***	0.00171***	0.0017156***
	(0.000111)	(0.000111)	(0.0001111)
Firm-reported layoffs	0.0000934	0.0000930	0.0000535
	(0.000315)	(0.000315)	(0.0003157)
Sector dummy variables		included	
Regional dummy variables		included	
Annual dummy variables		included	
Constant	-3.695***	-3.697***	-3.715***
	(0.189)	(0.189)	(0.191)
LR test of rho=0 (p-value)	<0.001	<0.001	<0.001
Log-Likelihood	-37204.80	-37204.79	-37209.35
Number of observations		2,902,724	
Number of individuals		833,359	
Number of establishments		4,260	

Note: Random-effects probit (coefficients). Standard errors in parentheses. 'LR' denotes likelihood-ratio.
**** p<0.01, ** p<0.05, * p<0.1.*
Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.5

Random-effects probit results for quit probability: Robustness checks

Quit=1	(1)		(2)
w_{ijt}^{range}	0.0317		
	(0.1115)		
w_{ijt}^{range} squared	-0.1005		
	(0.0919)		
w_{ijt}^{CDF}			-0.619***
			(0.0724)
w_{ijt}^{CDF} squared			0.663***
			(0.0644)
w_{ijt}^{abs}	0.273***		0.165**
	(0.0791)		(0.0764)
w_{ijt-1}^{abs}	0.0255		-0.0258
	(0.0714)		(0.0700)
Tenure	-0.0372***		-0.0350***
	(0.00275)		(0.00276)
Tenure^2	0.000592***		0.000522***
	(0.000104)		(0.000105)
Experience	-0.00217		-0.00108
	(0.003872)		(0.00388)
Experience^2	-0.0000862		-0.000136
	(0.000114)		(0.000114)
Unskilled worker		reference	
Skilled worker/ Craftsman	-0.0266*		-0.0126
	(0.0150)		(0.0151)
Technician/ Foreman	-0.0868**		-0.119***
	(0.0359)		(0.0362)
Clerk	0.0466***		0.0261
	(0.0170)		(0.0173)
Secondary school leaving certificate		reference	
Secondary school leaving certificate	0.127***		0.135***
and apprenticeship	(0.0207)		(0.0208)
(Technical) college entrance qualification	0.0913*		0.0898*
	(0.0518)		(0.0518)
(Technical) college entrance qualification	0.262***		0.263***
and apprenticeship	(0.0275)		(0.0275)
Workforce of establishment in [10;49]		reference	
Workforce of establishment in [50;199]	0.0521		0.0627
	(0.0661)		(0.0662)
Workforce of establishment in [200;999]	-0.0684		-0.0583
	(0.0654)		(0.0654)
Workforce of establishment in $\geq$ 1000	-0.173***		-0.166**
	(0.0658)		(0.0659)
Share of unqualified workers within the	-0.136***		-0.149***
establishment	(0.0247)		(0.0253)
Works council	-0.152***		-0.139***
	(0.0319)		(0.0321)
Collective bargaining	0.244***		0.246***
	(0.0312)		(0.0312)
Firm-reported quits	0.00161***		0.00166***
	(0.000110)		(0.000111)
Firm-reported layoffs	0.000126		0.0000962
	(0.000315)		(0.000315)
Sector dummy variables		included	
Regional dummy variables		included	
Annual dummy variables		included	
Constant	-4.441***		-3.684***
	(0.213)		(0.192)
LR test of rho=0 (p-value)	<0.001		<0.001
Log-Likelihood	-37266.52		-37213.51
Number of observations		2,902,724	
Number of individuals		833,359	
Number of establishments		4,260	

Note: Random-effects probit (coefficients). Standard errors in parentheses. 'LR' denotes likelihood-ratio.
**** p<0.01, ** p<0.05, * p<0.1.*
Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.6

Descriptive statistics

Variable	Mean	Standard Deviation
$(w^{new} - w^{old})_{it}$	0.0170	0.1426
$wage\ cut_{it}$	0.4286	0.4949
$(rank^{new} - rank^{old})_{it}$	-0.0244	0.2390
$(range^{new} - range^{old})_{it}$	0.0109	0.1977
$(CDF^{new} - CDF^{old})_{it}$	-0.0203	0.2386
1996	0.0897	0.2857
1997	0.0866	0.2812
1998	0.0650	0.2465
1999	0.1279	0.3340
2000	0.1288	0.3351
2001	0.1340	0.3406
2002	0.0653	0.2471
2003	0.0738	0.2614
2004	0.0882	0.2836
2005	0.1407	0.3478
Mobility to larger establishment (Dummy variable)	0.1827	0.3864
Secondary school leaving certificate	0.0631	0.2432
Secondary school leaving certificate and apprenticeship	0.7981	0.4014
(Technical) college entrance qualification	0.0124	0.1109
(Technical) college entrance qualification and apprenticeship	0.1263	0.3322
Experience	16.8494	6.8814
Experience^2	331.2462	234.0323
Firm-reported quits	30.9771	59.9449
Firm-reported layoffs	7.0494	19.5800
Number of observations	5,464	
Number of individuals	5,280	

Table A3.7

Random-effects probit results for quit probability (job-to-job mobility within one year)

Quit=1	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\bar{w}_{jt}$	0.0756							
	(0.0518)							
$\hat{w}_{ijt}^{inside}$		0.148***						
		(0.0569)						
w_{ijt}^{rank}			0.0589**	-0.806***				
			(0.0284)	(0.0722)				
w_{ijt}^{rank} squared				0.878***				
				(0.0676)				
w_{ijt}^{range}					-0.167***	-0.280***		
					(0.0396)	(0.105)		
w_{ijt}^{range} squared						0.102		
						(0.0875)		
w_{ijt}^{CDF}							0.0632**	-0.765***
							(0.0281)	(0.0703)
w_{ijt}^{CDF} squared								0.806***
								(0.0629)
Further control variables				included				
LR test of rho=0 (p-value)	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001	<0.001
Log-Likelihood	-44490.86	-44488.55	-44489.77	-44405.59	-44483.07	-44482.39	-44489.40	-44407.59
Pr(quit=1| $\bar{x}$)	0.000231	0.000230	0.000230	0.000223	0.000230	0.000230	0.000230	0.000225
Number of observations				2,904,095				
Number of individuals				833,635				
Number of establishments				4,279				

Note: Random-effects probit (coefficients). Standard errors in parentheses. 'LR' denotes likelihood-ratio.
Further control variables as in Table A3.3.
**** p<0.01, ** p<0.05, * p<0.1.*
Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

Table A3.8

Estimation results for consequences of mobility (job-to-job mobility within one year)

	(1)	(2)	(3)	(4)	(5)	(6)
		linear regression			probit estimation	
		$(w^{new} - w^{old})_{it}$			$wage\ cut_{it}$	
$(rank^{new} - rank^{old})_{it}$	0.295***			-1.005***		
	(0.0103)			(0.0371)		
$(range^{new} - range^{old})_{it}$		0.345***			-0.822***	
		(0.0129)			(0.0366)	
$(CDF^{new} - CDF^{old})_{it}$			0.302***			-1.070***
			(0.0104)			(0.0386)
Further control variables				included		
(Pseudo) R^2	0.2707	0.2472	0.2791	0.1636	0.1014	0.1754
Number of observations				6,835		

Note: Marginal effects at $\bar{x}$ are presented for the probit estimates. Robust standard errors clustered for 6,403 individuals in parentheses.

Further control variables as in Tables 3.5 and 3.6.

**** p<0.01, ** p<0.05, * p<0.10.*

Data: Cross-sectional model of the LIAB. Years 1995-2005 (own calculations).

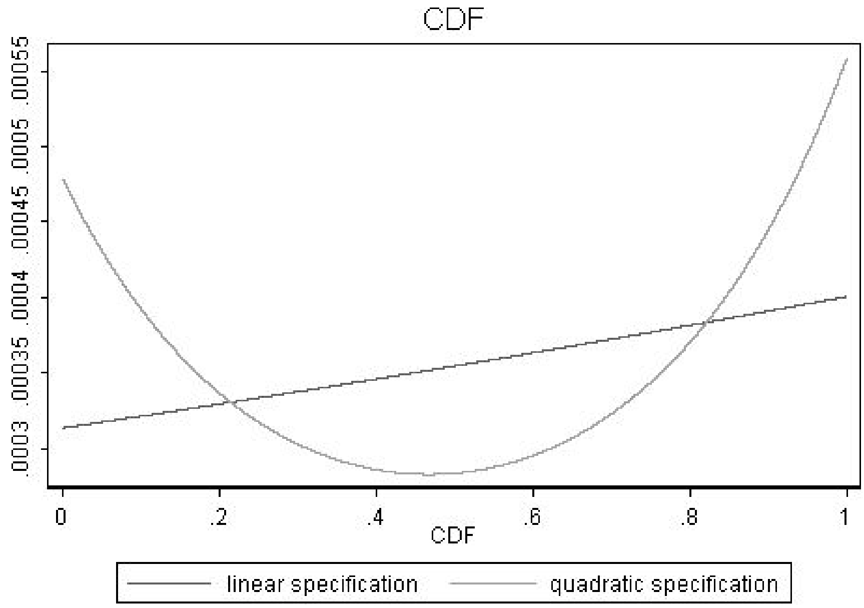

Figure A3.1: Predicted quit probabilities and relative wage positions (at $\bar{x}$) for w_{ijt}^{CDF}

4 The Acceptance of Earnings Losses After Voluntary Mobility

Because rational individuals know that they cannot always get what they want, they are assumed to make appropriate adjustments. However, little is known about trade-off reasoning in labor market mobility decision making. The objective of this paper is to analyze the effect of job-specific amenities on the decision to voluntarily accept wage cuts. Application of German household data reveals that voluntarily mobile workers are more likely to accept lower wages when strain can be improved. In other words, the considered mobile workers trade off amenities and monetary rewards when changing employers.[1]

[1]The chapter is a revised version of "The Acceptance of Earnings Losses After Voluntary Mobility", *Economics: The Open-Access, Open-Assessment E-Journal*, Vol. 5, 2011-2. http://dx.doi.org/10.5018/economics-ejournal.ja.2011-2

4.1 Introduction

Today's labor markets are characterized by a large degree of flexibility. Among a variety of aspects, labor market mobility contributes to this flexibility (OECD 1997). In recent times, a growing strand of literature corroborates that a considerable fraction of workers are changing jobs at the cost of wage cuts. In Germany, a large number of workers are shown to be mobile toward lower wages. Fitzenberger and Garloff (2007) refer to establishment-to-establishment transitions during two successive years and show that more than one in five individuals are mobile with wage cuts. Jolivet et al. (2006) apply data from the European Community Household Panel Survey to reveal that 36.3% of job-to-job transitions in Germany are accompanied by wage cuts. The authors define job-to-job mobility as transitions without noticeable unemployment spells of less than one month.

Transitions to lower wages are not a typical German phenomenon. In their cross-country analysis, Jolivet et al. (2006) show that almost one in five individuals is mobile to lower wages in Portugal and Belgium. The largest shares of wage cuts are observed in Denmark, France, and Germany. In these countries, more than 34% of mobile individuals suffer wage cuts in the period of mobility. In line with this result, Postel-Vinay and Robin (2002) show that more than one in three workers changing jobs directly did so at the cost of a wage cut.[2] For the United States, Jolivet et al. (2006) indicate that 23% of job-to-job transitions are to lower wages.[3] Nosal and Rupert (2007) utilize the Panel Study of Income Dynamics and show that about two in five individuals (voluntarily) change to lower wages. The results of these studies for different countries indicate that scientists should turn their attention to the reasons for mobility with wage cuts.

This paper sets forth an analysis of the reasons for job-to-job mobility to lower wages with a special focus on changes in different non-pecuniary job characteristics after the transition. It utilizes the German Socio-Economic Panel (GSOEP in the following; see

[2]Using French data, Postel-Vinay and Robin (2002) refer to direct mobility as job-to-job mobility with a maximum intervening unemployment spell of 15 days.

[3]Using the Panel Study of Income Dynamics, the authors refer to job-to-job mobility when intervening unemployment, if any, does not exceed three weeks.

Wagner et al. 2007), which includes questions on the reasons for job termination at the previous employer and surveys comparisons between both jobs. This is a major enhancement to previous papers because it allows one to determine whether workers voluntarily accept wage cuts in order to improve job-specific non-wage amenities.

This chapter is structured as follows. The next section illustrates briefly the basic framework and the research hypotheses. Section 4.3 describes the data set, main variables, and econometric models. I present the econometric results for the impact of subjective improvements in different job-specific characteristics on the decision to accept wage cuts in section 4.4. A conclusion is presented in section 4.5.

4.2 Framework

Recent literature considers wage cuts a result of job termination. In Jolivet (2009), workers are allowed to change jobs directly to lower wages because their only alternative is non-employment. These transitions are referred to as job reallocations and are also mentioned in other studies (e.g., Jolivet et al. 2006). Other theoretic approaches explain wage cuts as an investment in future wage growth (Connolly and Gottschalk 2008, Postel-Vinay and Robin 2002). It is also reasonable to change to a new employer offering lower wages if the wage cut at the current employer had been larger (see, e.g., Shi 2009 or Mortensen and Pissarides 1994). Schneck (2010) empirically suggests the prevalence of investments in future wage growth but also revealed that a substantial fraction of workers are mobile to permanently lower wages. Because workers are shown to accept lower wages on a permanent basis, other determinants are hypothesized to affect mobility decisions. For example, it is suggested that job-specific (non-wage) amenities affect the job choice (Nosal and Rupert 2007).

Economic and psychological literature, however, lack detailed information about the reasons for accepting lower wages. The basic idea of this paper proposes that differences in wages between two jobs might be balanced out by differences in (non-wage) job characteristics. Analogously to Rosen (1974, 1986) one could hypothesize that jobs consist of bundles of various characteristics with implicit, or hedonic, prices. Competent and self-supporting individuals, however, know that they cannot always get what they want, and that is the reason why they are expected to make appropriate adjustments. More specifically, individuals are expected to know that it is unlikely to find a *better* job with a higher wage, more flexible work time arrangements, and more job security right at their front door. It is important to analyze the extent of trade-off reasoning in the context of labor market mobility because ”Trade-off reasoning should be so pervasive and so well rehearsed as to be virtually automatic for the vast majority of the [...] population” (Tetlock 2000, p. 239). For this reason, this analysis refers to utility maximizing individuals who maximize their utility U.

Here, I assume that workers only change jobs if the utility U of worker i at employer j in time t exceeds the utility at the previous employer:[4]

$$U_{ijt} > U_{i,j-1,t-1} \tag{4.1}$$

Workers are confronted with job offers which contain information on the wage and a set of various job-specific amenities. Wage offers of employer j are offered to worker i in t independently of the worker's marginal willingness to avoid disamenities or to pay for amenities. Utility maximization implies that the worker changes employer if:

$$U_{ijt}(wage, amenities) > U_{i,j-1,t-1}(wage, amenities) \tag{4.2}$$

This paper concentrates on whether voluntary mobile workers accept a decrease in wages in exchange of an improvement in amenities. For this reason, the article mainly focuses on the theory of trade-off reasoning. The exclusive concentration on voluntary quits in the paper is assumed to assure that the drop in wages is compensated for by improvements

[4]Mobility costs are ignored. In addition, this paper is only responsive to short-term wage cuts which might pay off in the long-run.

in amenities.

$$\underbrace{wage_{ijt} - wage_{i,j-1,t-1}}_{-} = \underbrace{U(amenities_{ijt} - amenities_{i,j-1,t-1})}_{+} \tag{4.3}$$

$$\begin{aligned} Pr(Wage\ Cut_{ijt} = 1) &= U_{ijt}(\underbrace{improvement\ in\ amenities}_{+}, S) \\ Pr(Wage\ Cut_{ijt} = 1) &= \Phi(\beta_0 + \beta' improvement\ in\ amenities_{ijt} + \delta' S_{ijt}) \end{aligned} \tag{4.4}$$

The hypothesis about trade-off reasoning is summarized in equation (4.3). Worker i balances out improvements in job-specific amenities between two jobs and the wage decline when changing employer in period t. The probability to accept wage cuts, then, is expected to be positively affected by certain job-specific amenities. S summarizes further determinants which might affect the decision to accept lower wages. I tested the hypothesis by application of the probability model in which Φ is the cumulative density function of the standard normal distribution. Evidence in favor of trade-off reasoning in mobility decisions is provided in case of a positive estimate for β. The following strategy to estimate the willingness to pay for amenities (β) exploits the preferences about wages and amenities that are revealed when workers change jobs voluntarily.[5] Precisely, utility-maximizing workers only change employers if job-specific amenities compensate them for the loss in wages. In the following, I describe the effects of the job-specific amenities 'flexible work schedules', 'subjective job security against job loss', 'commuting', 'promotion possibilities', and 'strain' on the probability to accept a wage cut.

The paper addresses whether workers trade off improvements in strain and wages. Strain is shown to negatively affect individual satisfaction (see, e.g., Loscocco and Spitze 1990). Cornelissen (2009) finds a negative effect of hard manual labor and stress (which are dimensions of job strain) on job satisfaction. According to Mobley (1977), dissatisfaction with a job is translated into thoughts of leaving the employer, evaluation of

[5] Note that selection of workers might bias the estimates. In regression analysis, biased estimates are obtained when unobserved determinants of the outcome and unobserved determinants of selection into the the sample are correlated. The correlation between unobservables, however, cannot be directly evaluated. I expect that my estimates rather provide an upper bound for the acceptance for wage cuts since the workers in my sample are indeed compensated for the loss in wages by amenities. Note that some workers might also be compensated for the improvement in job-specific amenities by sacrificing (large) wage markups.

alternatives, and mobility because starting a new job is expected to result in a higher satisfaction. In fact, Judge (1993) shows that dissatisfied workers are more likely to quit than other individuals. Literature, however, lacks information on whether mobile individuals are willing to accept wage cuts in order to leave the dissatisfying job. This paper assesses whether individuals who expect decreasing strain when changing jobs are willing to accept lower wages. Analogous argumentation is expected to hold for improved job security by wage cuts because Cornelissen (2009) shows that satisfaction with the job is negatively affected by worries about (perceived) job security.

Based on the question of Altonji and Paxson (1988) on whether workers are willing to sacrifice wage gains for better working hours when quitting a job, I ask whether workers are even willing to accept wage cuts for an improvement of work time regulations. The main reason for a special focus on the latter hypothesis is that individuals face a trade-off between time constraints and monetary rewards. To be more precise, if the current employer offers few possibilities for flexible leisure, then, working at a new employer with more flexible working schedules might be preferred despite lower wages. In other words, workers know that it is very problematic (almost impossible) to achieve the highest flexibility without paying a price for it.

Do improvements in commuting expenses affect the acceptance of lower wages? Van Ophem (1991) shows that the commuting distance exhibits a significant impact on the search probability. Specifically, results show that the higher the distance from home to the workplace (in minutes), the higher the search propensity. This result suggests that commuting is an important determinant for labor market mobility because job search is a good predictor of actual mobility (see Cornelissen 2009). However, little is known about the relationship between commuting expenses and mobility to lower wages. For this reason, the paper asks whether subjective improvements in commuting are paid for by wage cuts. Note that commuting might also be expensive in monetary terms because larger distances are assumed to increase the price to get from home to work which directly reduces profitability of the job. Possible critique arises because commuting might be an economic outcome variable because commuting can be seen part of the wage. More specifically, economic theory predicts that commuters are compensated for commuting by higher wages. Stutzer and Frey (2008, p. 339), however, find that ”people spend a lot of time commuting and often find it a burden” because subjective well-being decreases with longer commuting time. For this reason, individuals also reveal trade-off reasoning when deciding to quit a job in order to change to a job with shorter or longer commuting time.

In addition, the possibilities for promotions at the new employer might affect the decision to accept wage cuts. Pfeifer and Schneck (2010) show that workers who change to higher relative wage positions compared to the previous establishment have, on average, a lower probability to change to lower wages. Workers who change to lower relative wage positions, in turn, likely suffer more wage cuts. For this reason, the authors do not present evidence in favor of trade-off reasoning in relative wage positions and wages. However, it is suggested that workers who change with wage cuts to a lower relative wage position might benefit from better chances for future promotions within the new firm. For this reason, it is argued that workers might pay for future promotion opportunities by wage cuts.

Usually workers evaluate these job-specific amenities before the transition. The data, in turn, refers to realized transitions with completed trade-off reasoning (a more detailed description of the data follows in the next section). For this reason, individual answers on the questions about subjective improvements in the new job might involve problems regarding cognitive dissonance reduction theory (Festinger 1957). This particular theory describes that unpleasant arousal drives people to resolve the cognitive inconsistency. In other words, if two cognitions are discrepant, individuals simply change one to make it consistent with the other. Here, workers might act contrary to their attitude because of mobility to lower wages. As a consequence, these workers adjust their cognition about the job in a positive way to balance out this effect. In the underlying case, workers might change their attitude toward the new job in a positive way as a consequence from the decision to be mobile to lower wages. As a result, workers who accept wage cuts report to be more satisfied with the new job compared to workers changing without wage cuts. If this is true, the estimated coefficients on subjective comparisons (improvements) between the previous and the current job would be upwardly biased. A direct test of this possible critique cannot be conducted by application of the GSOEP.

4.3 Data and procedure

4.3.1 Data

This study utilizes the GSOEP household survey to examine the impact of job-specific amenities on the probability of being mobile with wage cuts. The main advantage of this data set stems from the fact that it includes subjective comparisons between the previous and current jobs. I restricted the analysis to German citizens who are employed full-time in two successive years during the period 1994–2007. The sample considers private sector employees with permanent contracts aged between 20 and 60 years. The lower age boundary is chosen because the school degrees are usually achieved before 20 years of age.[6]

The data include annual information on the last monthly gross wage of individual i in period t (measured in Euro) which is applied in the consecutive analysis. I apply the consumer price index provided by the Statistisches Bundesamt Deutschland (annual averages, with year 2005 = 100) to deflate the wages. In addition, the questionnaire asks the "How many hours are stipulated in your contract (excluding overtime)?" The corresponding information is utilized to calculate the hourly wage of individuals. The hourly wage (w_{ijt}) as well as the real hourly wage (w_{ijt}^{real}) of individual i in period t at employer j are defined as follows:

$$\begin{aligned} w_{ijt} &= \frac{monthly\ wage_{ijt}}{4.33 * contractual\ weekly\ working\ time_{ijt}} \\ w_{ijt}^{real} &= \frac{deflated\ monthly\ wage_{ijt}}{4.33 * contractual\ weekly\ working\ time_{ijt}} \end{aligned} \tag{4.5}$$

Note that the GSOEP also includes information on overtime or the actual hours worked.

[6] I consider the years of education which is based on information provided by the GSOEP.

I decided to concentrate on the contractual working hours because this measure is less affected by (cyclical or employer-specific) fluctuations. As the data are set up as a panel, information about the wage in the previous year is utilized to determine wage cuts and wage improvements. To examine the probability of wage cuts, a binary variable is constructed to illustrate whether individuals are mobile to lower wages or not:[7]

$$
\begin{aligned}
Wage\ Cut_{ijt} &= \begin{cases} 1 & \text{mobility to lower wages} \quad (w_{it} - w_{i,j-1,t-1} < 0) \\ 0 & \text{mobility to higher wages} \quad (w_{it} - w_{i,j-1,t-1} \geq 0) \end{cases} \\
Wage\ Cut_{ijt}^{real} &= \begin{cases} 1 & \text{mobility to lower wages} \quad (w_{it}^{real} - w_{i,j-1,t-1}^{real} < 0) \\ 0 & \text{mobility to higher wages} \quad (w_{it}^{real} - w_{i,j-1,t-1}^{real} \geq 0) \end{cases}
\end{aligned}
\tag{4.6}
$$

In order to account for the individual trade-off reasoning appropriately, the analysis attempts to identify voluntary mobility, which is defined as an unconstrained decision of the individual. The underlying GSOEP includes detailed retrospective information about labor mobility. Each year, the questionnaire asks whether a new job was started at a new employer.[8] Individuals who reported an employer change, then, are asked whether they resigned on their own initiative. In the subsequent analysis, only those reporting a resignation on their own initiative are considered. In addition, I focus on mobile workers who changed employer within one month. This criterion was instituted to meet the definition of job-to-job mobility where individuals have to be mobile within one month (Jolivet et al. 2006, Royalty 1998). In sum, 800 voluntary employer-to-employer transitions of 670 individuals who quit their jobs up to four times are considered. Note that the sample size of the entire GSOEP data is considerably reduced by implementation of the restrictions but the sample size is comparable to the one reported in Villanueva (2007).

A diversity of subjective improvements of different job characteristics are surveyed in

[7] Wage information of the year 1993 is utilized to calculate the wage growth of mobile workers in 1994. I drop reported wages of zero.

[8] The analysis excludes workers starting their first job or have a new job after a break. Individuals who report a job change within a firm and individuals who become self-employed are also not subject of the underlying analysis. The paper, hence, focuses on transitions between different employers. Unique information about this special pattern is available from 1994 onwards.

the data. More specifically, the data set includes information about comparisons between the previous and current jobs if individuals reported a job change. The corresponding question read as follows: "How would you judge your present position compared to your last one? In what ways has it improved, stayed the same, or worsened?" This particular question considers the following characteristics:

1. wages
2. job type
3. chances for promotion
4. work load (strain)
5. length of commute to and from work
6. work schedule regulations (work time)
7. fringe benefits
8. security against job loss[9]

In the subsequent analysis, the answers to the question on strain, job security, commuting, work time, and promotions are applied to analyze the impact of trade-off reasoning on the decision to be mobile to lower wages. Another question asks whether the individual uses his or her knowledge and skills more, the same, or less than in the previous job. This variable is to describe whether the worker's skills meet the required ones in the new job and can be interpreted as a match quality indicator. Table 4.1 shows the descriptive statistics and reveals that only very few transitions (9.5%) are accompanied by a subjective worsening of wages. In the following, the paper concentrates on dummy variables which describe improvements or worsenings of the subjective comparisons. The corresponding frequencies of the subjective comparisons are shown in Table A4.1 while means and standard deviations are presented in Table A4.2.

[9] The questions and potential answer categories differ slightly over the years. No information is available in the 2008 wave of the GSOEP. Regarding fringe benefits, the German questionnaire refers to "betriebliche Sozialleistungen" while the English questionnaire refers to "benefits".

4.3.2 Descriptive statistics

Table 4.1 presents descriptive statistics on wage changes induced by voluntary job-to-job mobility. Application of nominal wages reveals that 24.38% transitions are executed to lower wages. These numbers are comparable to the ones reported in Fitzenberger and Garloff (2007). For real wages, however, the results are closer to the ones reported in Jolivet et al. (2006), where about one in three transitions are to lower wages. On average, all directly mobile workers generate a wage markup of about 15.40% (nominal) and 13.68% (real), respectively. This average wage premium for mobility is another reason for the conventional hypothesis that employer-to-employer mobility is voluntary.

Table 4.1
Descriptive statistics on wage changes after mobility

	share of wage cuts	10% percentile	mean	90% percentile	Number of observations
$\frac{w_{ijt}}{w_{i,j-1,t-1}}$	0.2438	0.8756	1.1540	1.4463	800
	(0.4296)		(0.3440)		
	0.0000	1.0163	1.2488	1.5484	605
	(—)		(0.3380)		
	1.0000	0.6857	0.8599	0.9811	195
	(—)		(0.1296)		
$\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$	0.3175	0.8635	1.1368	1.4246	800
	(0.4658)		(0.3381)		
	0.0000	1.0309	1.2563	1.5498	546
	(—)		(0.3393)		
	1.0000	0.7164	0.8801	0.9910	254
	(—)		(0.1267)		
subjective worsening in wages	0.0950				800
	(0.2934)				

Standard deviations in parentheses.
$\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$: *Deflated gross wages (year 2005 = 100).*

The share of workers who are mobile with wage markups, however, is very different from the workers who are mobile to lower wages. The average wage markup amounts to 24.88% (nominal) and 25.63% (real) for upwardly mobile individuals. Downwardly mobile workers, in turn, suffer an average wage cut of more than 10%. The wage markups and wage cuts presented here are comparable to the ones presented in Fitzenberger and Garloff (2007). Note that the subjective perception of declines (worsenings) in wages is, by far, smaller than the number of wage cuts. Precisely, 9.50% of transitions are accompanied by subjective wage cuts while more than two in five transitions are to lower hourly wages. This suggests that the disutility introduced by monetary losses might be offset by other dimensions of the current job which directly adverts to trade-off reasoning in job mobility.

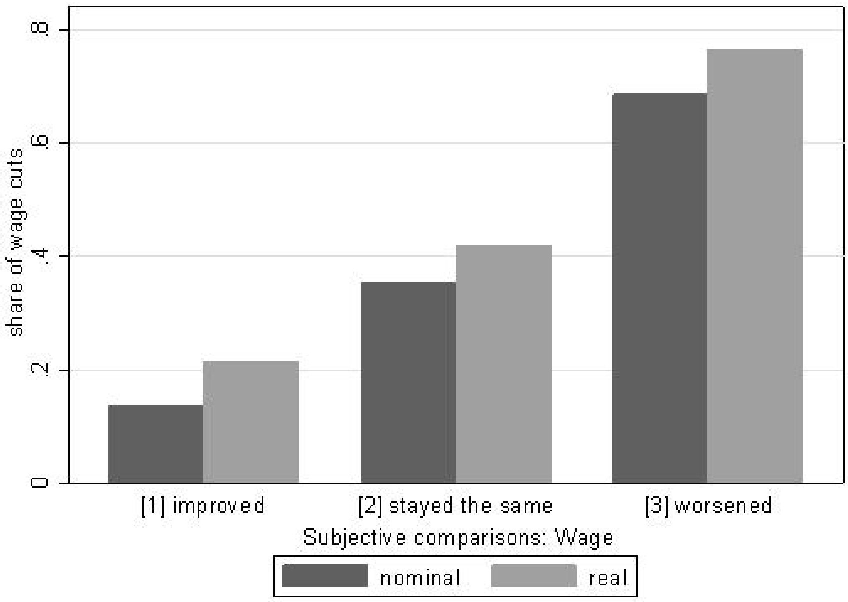

Figure 4.1: Share of wage cuts by subjective cognition about wage change

Number of observations: $N_{improved} = 521,\ N_{stayed\ the\ same} = 203,\ N_{worsened} = 76$

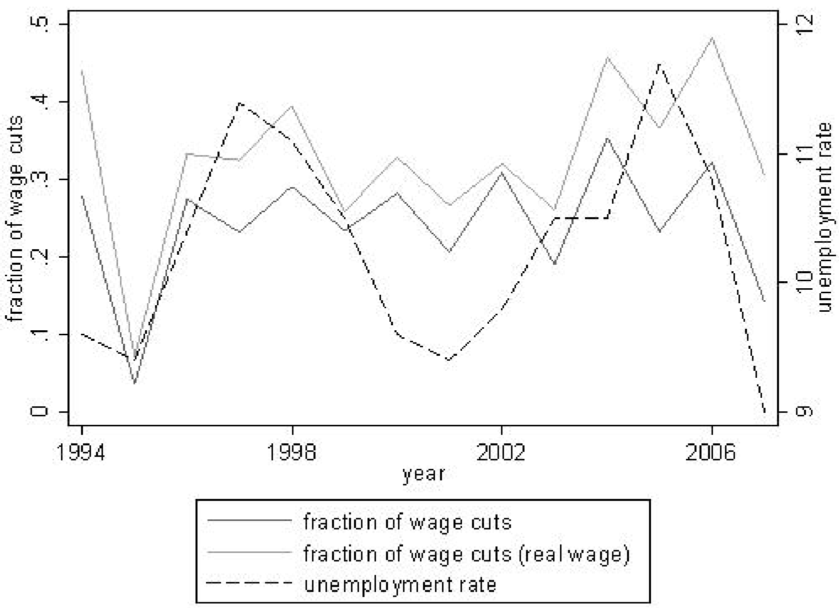

Figure 4.2: Share of wage cuts by year

Figure 4.1 presents the share of wage cuts by the categories of subjective comparison of wages between two jobs. As expected, the share of workers with realized wage cuts increases with increasing subjective worsenings about the wage change. In other words, 68.42% (nominal) and 76.32% (real) of individuals who report subjective worsenings in wages indeed suffer wage cuts, whereas only 13.63% and 21.31% of individuals who report a subjective improvement in wages actually experience wage cuts. Figure 4.2 shows that the share of mobility with wage cuts within a certain period are rather unaffected by the business cycle. To be more precise, the period between 1996 and 2006 was especially characterized by a relatively stable share of transition to lower wages. Note that this does not imply that mobility is equally common across the different phases of the business cycle (see footnote 11). In sum, the descriptive statistics show that mobility to lower wages is frequent across different phases of the business cycle, which accentuates the importance of an analysis of the reasons for the acceptance of lower wages.

4.3.3 Methods and procedure

The research question on whether workers accept wage cuts in exchange for improvements in job amenities can directly be addressed in a probit model because the dependent variable on whether a wage cut was accepted or not is binary by construction. Literature recommends the analysis of binary dependent variables by application of binary choice models. Here, a probit model that relates to equation (4.4) was utilized. Equation (4.7) shows the applied probit model, where X_{ijt} stands for dummy variables which describe improvements or worsenings between the previous and the current jobs, whereas S_{ijt} describes sociodemographic information and other determinants affecting mobility to lower wages. The corresponding descriptive statistics are presented in Table A4.2.

$$Pr(Wage\ Cut_{ijt} = 1) = \Phi(\alpha + \beta' X_{ijt} + \delta' S_{ijt}) \qquad (4.7)$$

Individual characteristics include gender, age, education (in years), and whether or not individuals live with a partner. Regional mobility is included in the analysis because

Yankow (2003) shows that changing locale affects wages. More specifically, I accounted for the federal state (Bundesland) in which an individual is working. If a worker changes to a job in a different federal state compared to the previous one, the corresponding dummy variable for regional mobility equals one. In addition, transitions from blue-collar to white-collar jobs are accounted for by a dummy variable. I also account for the economic environment in different years. Precisely, I include the growth of unemployment rate into the analysis.[10] The number of individual mobility describes the calculated number of quits on own initiative between 1985 and the year of the interview. Note that the minimum is one because the current quit is included.[11]

In a next step, the marginal willingness to pay for different amenities is estimated via application of OLS regression. The dependent variable describes the wage change while the set of control variables is identical to the one in the probit model discussed previously.

$$\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} = a + b'X_{ijt} + d'S_{ijt} + u_{ijt} \tag{4.8}$$

Finally, the corner solution (tobit) estimation approach is applied which combines aspects of the binomial probit for the distinction of $\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} \geq 1$ and $\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} < 1$ and the

[10]I apply the unemployment rate provided by the Sachverständigenrat zur Begutachtung der gesamtwirtschaftlichen Entwicklung (Table 090; with respect to share of civilian labor force). Unemployment growth is defined as $unemp_t - unemp_{t-1}$.

[11]Table A4.2 presents descriptive statistics for the control variables which are included in the subsequent multivariate analysis. Workers who voluntarily change jobs are, on average, about 35 years old. This finding can be interpreted with the hypothesis that middle-aged workers assess their own aspiration levels best (Clark et al. 1996). More than two in three mobile individuals are renters. The average education in years is between 12 and 13 years. Regional mobility plays a minor role by simple consideration of its frequency, since workers are shown to leave their federal state for a new job rarely. Only 4.75% of individuals perform cross-border transitions between federal states in Germany. A minority of mobile individuals life together with a partner (21.25%). About one in 20 transitions are from a blue-collar job to a white-collar job. It is also necessary to account for the workforce in the previous and the current firm (see, e.g., Brown and Medoff 1989). 13.25% of individuals are leaving a firm with more than 2,000 employees while 17.13% of mobile workers are employed at a new firm with more than 2,000 employees. The following cross-table illustrates the number of observations by firm-size categories.

Number of observations by firm-size

Dummy variable for workforce$_{i,j-1,t-1}$ > 2,000	Dummy variable for workforce$_{i,j,t}$ > 2,000		
	0	1	Total
0	604	90	694
1	59	47	106
Total	663	137	800

regression model for $E[\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}}|X_{ijt}, S_{ijt}, \frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} < 1]$. The setting is adequate for Tobit because individuals decide on how much they are willing to pay for a *better* job rather than first deciding on whether to accept a wage cut and then, if this first decision is affirmative, decide on how much to pay for the new job.

$$\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}}^{*} = e + f'X_{ijt} + g'S_{ijt} + v_{ijt} \tag{4.9}$$

$$\frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} = \begin{cases} \frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}}^{*} & \text{if } \frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} < 1 \\ 0 & \text{if } \frac{w_{ijt}^{(real)}}{w_{i,j-1,t-1}^{(real)}} \geq 1 \end{cases} \tag{4.10}$$

Estimation of the corner solution model, then, allows to compute the marginal willingness to pay for different amenities by wage cuts, given that the individual changes to lower wages.[12] As the data are set up as a panel, I am able to make effort to control for unobserved individual heterogeneity. All the tests do not reject the null hypothesis of no individual heterogeneity.[13]

Note that the analysis of this particular trade-off reasoning might be characterized by simultaneity in the acceptance of wage cuts and improvements in the new job. This problem might introduce problems regarding endogeneity. One way to deal with this type of problem is to utilize a two-stage least square estimator, where I need to identify

[12] The tobit approach can be viewed as a special case of the so-called Heckman sample selection model (Heckman 1979) when the selection equation and the regression equation are identical. One reason to refer to the tobit model is that it is problematic to define a reasonable selection equation because of a lack of literature on the acceptance of wage cuts in voluntary mobility decisions.

[13] Precisely, different tests were conducted for the entire sample. A likelihood-ratio test was conducted in order to assess whether individual random-effects were evident in the probit model which explains whether a wage cut was accepted or not. The Breusch-Pagan Lagrange multiplier test (Breusch and Pagan 1979) was applied to test for unobserved individual heterogeneity in the linear model on the wage change. Finally, a likelihood-ratio test was applied for the tobit model. The null hypothesis cannot be rejected in all cases. Note that the Chow test type for poolability of the data over time (see Baltagi 2008, chapter 4) also rejects the null hypothesis. For this reason, it is suggested that the data should not be pooled across all periods. Few observations in certain periods, however, enforce pooling of the data (see the number of observation per years below). In order to address this problem, I conduct robustness checks for different time periods and discuss potential changes in the effects.

Year	1994	1995	1996	1997	1998	1999	2000
Number of observations	25	28	51	43	48	81	85
Year	2001	2002	2003	2004	2005	2006	2007
Number of observations	116	81	84	48	30	31	49
Total number of observations:	800						

instrument variables. However, it is hard to find any variable which is partially correlated with subjective improvements between two jobs and exogenous in the decision to accept wage cuts. Given any simultaneity in trade-off reasoning, the following coefficients do not have a causal interpretation.

4.3.4 Specification

This section concentrates on the choice of specification. As mentioned above, the data include a large set of dummy variables for subjective comparisons which can be included in X_{ijt}. Note that some of the dummy variables of subjective comparisons between jobs are highly correlated. Table A4.3 presents the correlation coefficients where Spearman's correlation and Tetrachoric correlations for binary variables are applied. Obvious problems regarding multicollinearity, however, are not revealed because of a maximum correlation coefficient of 0.6501 for a worsening in fringe benefits and a worsening in job security. Note that I abstract from Tetrachoric correlations of -1.000 between improvements and worsenings in job-specific amenities which are plausible because an improvement can never be associated of the same subjective comparison measure. Regarding the choice of specification, the match-specific component (comparison of use of skills) is included in all specifications because of its importance on the wage determination in economic literature. As discussed in the framework above, individual preferences about trade-off reasoning are also revealed when comparing flexible work schedules, strain, commuting, promotion chances, and perceived job security between the previous job and the current job. For this reason, these determinants are subject to the first ("preferred") specification.

In a further step, I extended the preferred specification by inclusion of dummy variables for subjective improvements and worsenings of fringe benefits and of the general job type. This specification, then, might be referred to as the full specification because all subjective comparisons (with exception of the subjective comparison of wages) are considered. Please note that subjective perceptions about the general job type and the use of skills are significantly correlated.[14] This suggests that both variables might describe

[14]The corresponding Tetrachoric correlation equals 0.6173 for an improvement in the general job type and and better use of skills and is the third highest correlation coefficient in Table A4.3. For worse

the subjective change in the match quality when comparing the current job to the previous job. Fringe benefits might be monetary amenities which are paid by the firm. For this reason, this measure might reflect some redeployment of wages rather than trade-off reasoning. The full specification, however, is expected to provide a valuable robustness check of the results obtained by the preferred specification.

In a next step, factor analysis is utilized in order to reduce the dimension from the multitude of dummy variables of subjective comparisons to a lower number of factors. Precisely, principal component factor analysis with orthogonal varimax rotation is conducted. The obtained factors are a set of independent and mutually orthogonal linear combinations of all of the subjective comparisons between the jobs. Because the choice of the number of factors is complex, one can rely on information criteria or one can search for solutions which are to be interpreted in an economically meaningful way. The Bayesian information criterion suggests considering six factors wherein the factor loadings can be meaningfully interpreted. The corresponding results are shown in Table 4.2.

Table 4.2 presents the factor loadings which are used for interpretation of the six factors, where the bold numbers describe the highest loadings for the different factors. One can learn from the table that factor 1 is highly affected by subjective comparisons in strain and work time regulations. For this reason, these variables are used to assign the label to factor 1 because workload and work schedules are dimensions of job-specific working conditions. Analogously, factor 2 can be interpreted as an improvement in 'job amenities', as fringe benefits and the perceived job security against job loss exhibit the highest factor loadings. Note that, for example, the factor loadings for an improvement in work time also loads high on the factor two. For this reason, a more flexible work schedule is suggested to affect factor 2 as well but is not directly included in the following interpretation of this particular factor. The remaining factors are defined in a similar way.

jobs in general and less use of skills, the correlation is similar (0.6066) and significant.

Table 4.2

Factor analysis with six factors

Rotated factor loadings (pattern matrix) and unique variances

Variable	Factor 1	Factor 2	Factor 3	Factor 4	Factor 5	Factor 6	Uniqueness
Interpretation of the factor	*working conditions*	*job amenities improved*	*match improved*	*commuting*	*match worsened*	*job amenities worsened*	
Strain ↓	**0.7224**	0.0181	0.0862	0.0226	0.0176	0.0932	0.4609
Strain ↑	**-0.6747**	0.0571	0.2635	-0.0341	0.3162	-0.0612	0.3671
Work time ↓	**0.6259**	-0.1730	0.2254	0.1308	0.3259	0.0014	0.4042
Work time ↑	**-0.5616**	0.4256	0.1224	-0.1563	0.0467	0.1743	0.4315
Fringe benefits ↑	-0.1447	**0.7672**	0.0955	0.0173	0.0550	0.0103	0.3780
Job security ↑	-0.0057	**0.7061**	0.1845	-0.0010	0.0774	-0.0601	0.4578
General job ↑	-0.1250	0.0378	**0.7325**	-0.0086	-0.1662	0.0269	0.4179
Use of skills ↑	0.0808	0.1743	**0.6448**	0.0061	-0.4248	0.0557	0.3638
Promotion chances ↑	0.0718	0.1430	**0.6134**	0.0895	0.1598	-0.4008	0.4040
Commuting ↑	-0.0590	0.0495	0.0760	**-0.8640**	0.0749	0.0420	0.2345
Commuting ↓	0.0506	0.0533	0.1102	**0.8457**	0.0553	0.0812	0.2576
Use of skills ↓	-0.1225	0.0290	-0.1259	-0.0536	**0.7662**	0.0595	0.3748
General job ↓	0.2454	0.2133	-0.2175	0.0325	**0.5831**	0.2444	0.4462
Promotion chances ↓	0.0350	0.1279	-0.1654	0.0354	0.0489	**0.7731**	0.3537
Job security ↓	-0.0256	-0.3762	0.1573	0.0809	0.1791	**0.5987**	0.4359
Fringe benefits ↓	0.2331	-0.4429	0.2563	-0.0766	0.1395	**0.4473**	0.4584

Method: principal component factors with orthogonal varimax rotation.

Number of observations: 800.

↑ *describes improvements,* ↓ *refers to subjective worsenings.*

The next step calculates the factor scores as proposed by Thomson (1951). Workers who report improvements in strain and work time are likely to have a negative factor score for 'working condition', whereas workers who change to jobs with worse strain and worse work schedules are more likely to be associated with positive scores for factor 1. Workers who report an improvement in commuting never obtain a positive factor score for the factor 'commuting', whereas workers reporting a worsening never obtain a negative value. Interpretation of the remaining factors is straightforward. Note that the determination of a set of factors allows a reduction in the dimensionality of the analysis but it can also hide what is going on at the disaggregated level. Therefore, inclusion of the factor scores into the estimation framework should only be viewed as a further robustness check.[15]

In sum, the analysis concentrates on three different sets of variables included in X_{ijt}.

[15] The following table shows the factor scores which have a mean close to zero and a standard deviation of one.

The first specification refers to the variables mentioned in the framework. Precisely, subjective comparisons of the use of skills, flexible work schedules, strain, commuting, promotion chances, and perceived job security are subject to the first specification. The second specification additionally accounts for comparisons in fringe benefits and the general job type. The set of variables in third specification contains the factor scores which are described above.

4.4 Results

This section presents the results of the multivariate analysis. At first, I accounted for the dummy variables for subjective improvements and worsenings in commuting, strain, work time, promotion chances, security against job loss, and the match indicator (preferred specification). Table 4.3 presents the results for the probit estimation framework on whether workers accepted a wage cut when changing jobs. Note that the endogenous variable varies over specifications. Precisely, specification (1) explains mobility to lower wages when accounting for gross wages, specification (2) corresponds to deflated gross wage cuts, and specification (3) presents the probit estimates for the subjective decline in wages. The link test for the corresponding probit models shows that the following specifications are satisfactory because $\hat{y}^2$ is insignificant in all the test equations (see Ramsey (1969) for a comparable test).

Regarding the above hypothesis of trade-off reasoning between subjective improvements in amenities and mobility decisions to lower wages, different specifications in Table 4.3 provide distinct insights. One can learn from specification (1) that individuals pay for an improvement in strain by lower wages. The coefficient is significant and positive, which

Descriptive statistics: Factor scores

Variable	Mean	Std. Deviation	Minimum	Maximum
Factor score 1	1.39e-09	1	-2.314253	2.998188
Factor score 2	6.73e-10	1	-2.629714	3.765187
Factor score 3	7.42e-11	1	-1.978934	2.801363
Factor score 4	3.31e-10	1	-1.594935	1.68852
Factor score 5	-1.41e-09	1	-1.771901	4.142451
Factor score 6	-1.38e-10	1	-1.640562	5.140013

Number of observations: 800.

implies that an improvement in strain compared to the previous job increases the probability for voluntary mobility to lower wages. As a result, trade-off reasoning is evident. The estimated coefficient in specification (1) equals 0.0638, which can be interpreted in that an improvement in strain increases the probability for mobility to lower wages by 0.0638 percentage points. This result reveals that trade-off reasoning between improvements in strain and wages is evident. The coefficient of subjectively improved commuting is comparable and is significant at the 10% level. Again, trade-off reasoning, as hypothesized above, becomes evident. The subjective evaluation of promotion opportunities also have sizable impact by considering the size as well as significance of the coefficients. Workers changing to a job with subjectively improved opportunities to climb up the hierarchy are less likely to accept wage cuts, whereas workers who change to worse future career prospects are more likely to suffer earnings losses. This result contradicts the ones obtained by Pfeifer and Schneck (2010), who report that a change in relative wage positions is positively correlated to a change in wages. Accordingly, transitions to lower relative wage positions which, in turn, increase future career prospects are accompanied by lower wages. The results obtained here, however, suggest that workers who change to jobs with better promotion opportunities are less likely to change to lower wages. The differences might stem from the definition of the variables in both studies: This study utilizes subjective comparisons between jobs which can be evaluated in a completely different manner when compared to an objective measure (the change in the relative wage position) as utilized in Pfeifer and Schneck (2010).

When accounting for real wages, the coefficient for improvements in work time becomes significant and indicates that an improvement in work schedules reduces the probability of wage cuts (see specification (2) in Table 4.3). This result contradicts trade-off reasoning between mobility decisions to lower wages and job-specific amenities. Note that specification (2) confirms the estimated signs obtained in the first specification for strain, work time, promotion opportunities, and the match indicator. The signs for the estimated coefficients for a worsening in security against a job loss and for the subjective worsenings in commuting differ across specifications. These effects, thus, are not only insignificant but also non-robust across specifications. The probability that an individual reports a subjective worsening in wages is described in specification (3). Note that some of the coefficients for the subjective wage cut contradict the ones obtained for the objective measures for wage cuts. An explanation for the effect that the sign of an improvement in work time is (insignificantly) positive in specification (3) but negative in specifications (1) and (2) might be attributed to cognitive dissonance reduction (as mentioned by Festinger 1957), which resolves the perception of a wage cut. In other words, workers might change their attitude toward the current job in a positive way (more flexible work schedules) as a consequence of the decision to accept perceived wage cuts. An interpretation for the different signs of the coefficient for a worsening in strain across specifications is that workers who are less satisfied with the current job have a higher probability to feel to be subjectively worse off in wages. A somewhat surprising result is that the match indicator variable does not contribute any significant on the probability to accept wage cuts. Individuals, however, are less likely to suffer wage cuts when better use of skills is achieved in the new job compared to the previous one. For less use of skills, negative effects are found in specifications (1) and (2), while a positive effect is revealed in specification (3). An explanation for this result might be that workers who are not able to use all of their skills might feel bored, which possibly introduces dissatisfaction with wages or perceptions of earnings losses. The negative coefficients in specifications (1) and (2) are hardly to explain. It might be hypothesized that workers change to jobs where they are not able to use all of their skills, but instead apply one very special and highly paid skill. Thus, especially for highly qualified specialists, less use of skills also might reduce the probability of wage cuts. The remaining control variables are, to the largest extent, insignificant. For the growth in the unemployment rate, I do not find any significant impact that confirms the considerations above. Cyclical fluctuations only have low impact on the acceptance of voluntary wage cuts.

Table 4.3
Probit model on whether workers accepted a wage cut

Variables	(1) mobility to lower wages	(2) mobility to lower wages^{real}	(3) subjective worsening of wages
Strain improved	0.0638*	0.0911**	0.0628**
	(0.0373)	(0.0406)	(0.0244)
Strain worsened	-0.0311	-0.0603	0.0203
	(0.0424)	(0.0468)	(0.0309)
Work time improved	-0.0263	-0.0695*	0.0183
	(0.0347)	(0.0381)	(0.0217)
Work time worsened	0.0275	0.0340	0.0289
	(0.0523)	(0.0549)	(0.0349)
Security against job loss improved	0.0277	0.0222	-0.00447
	(0.0352)	(0.0386)	(0.0195)
Security against job loss worsened	0.0577	-0.0162	0.0103
	(0.0655)	(0.0648)	(0.0337)
Use of skills improved	-0.0369	-0.0408	-0.0172
	(0.0336)	(0.0374)	(0.0182)
Use of skills worsened	-0.0469	-0.0412	0.000916
	(0.0453)	(0.0521)	(0.0242)
Commuting improved	0.0669*	0.0588	0.0110
	(0.0386)	(0.0421)	(0.0226)
Commuting worsened	-0.00235	0.00342	0.00761
	(0.0390)	(0.0420)	(0.0228)
Chances for promotion improved	-0.0765**	-0.0715**	-0.0850***
	(0.0321)	(0.0362)	(0.0189)
Chances for promotion worsened	0.199***	0.256***	0.163***
	(0.0754)	(0.0764)	(0.0604)
Homeowner	0.0617*	0.0422	0.0108
	(0.0334)	(0.0365)	(0.0200)
Number of previous individual mobility	-0.00341	0.0116	0.00156
	(0.0167)	(0.0188)	(0.00870)
Age	0.0335**	0.00784	0.00860
	(0.0152)	(0.0166)	(0.00803)
Age^2	-0.000335*	-5.56e-06	-8.12e-05
	(0.000202)	(0.000220)	(0.000105)
Education in years	-0.0187***	-0.0109	-0.00485
	(0.00663)	(0.00733)	(0.00382)
Blue-collar to white-collar transition	0.000722	-0.0233	0.0265
	(0.0645)	(0.0704)	(0.0411)
Male	-0.00425	-0.0250	-0.0112
	(0.0337)	(0.0378)	(0.0183)
Partner	0.0154	0.0128	0.0226
	(0.0400)	(0.0437)	(0.0262)
Firm more than 2,000 workers	0.0423	0.0515	-0.0135
	(0.0442)	(0.0475)	(0.0224)
Previous firm more than 2,000 workers	-0.0185	0.00117	0.0634*
	(0.0461)	(0.0523)	(0.0355)
Regional mobility	-0.0683	-0.117	0.00173
	(0.0641)	(0.0713)	(0.0425)
Growth in unemployment rate	0.0187	0.00228	-0.00481
	(0.0190)	(0.0218)	(0.0107)
Number of observations		800	
Pseudo R^2	0.074	0.054	0.145
Predicted $\Pr(y = 1 \mid \bar{x})$	0.226	0.309	0.066

Marginal effects at $\bar{x}$ are presented.
Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

In order to check the robustness of the results displayed in Table 4.3, I conducted probit estimation using the full sample of subjective comparisons between the current and previous jobs and the factors discussed in section 4.3.4. Consideration of all subjective comparison measures in the data (with exception of the subjective change in wages) confirms the estimated effects for strain. Table 4.4 shows that workers who change to a job with a better level of strain compared to the previous job are significantly more likely to accept wage cuts, where the estimated coefficients are larger in size compared to the ones presented in Table 4.3. Although most of the effects for the objective wage cut are robust to the effects presented in Table 4.3, the estimated coefficients for security against job loss are non-robust in the specification for the perceived wage cut. A similarity to the results shown in Table 4.3 is that most effects are statistically insignificant. This also holds for the effects of the subjective comparison of the job in general. More fringe benefits in the current job compared to the previous one significantly decrease the probability that workers suffer wage cuts. Note that, however, fringe benefits can also be included in the monthly payments, and thus, might be interpreted as monetary job-specific amenities. Workers who change to less fringe benefits perceive significant wage losses. This might be explained by habit-persistence, where workers get used to different amenities and react with strong negative perceptions in case amenities disappear.

Before turning the focus on the absolute and relative wage change, the six factors obtained via factor analysis described above are applied to check the robustness of the results. Table 4.5 shows that a better match quality significantly reduces the probability of the acceptance of earnings losses in all specifications. This might be explained by economic literature where the match quality is a main factor of wage determination. A worsening in job amenities is not suggested to be compensated for by higher wages. In fact, the reverse is true because individuals are significantly more likely to suffer lower wages at the new employer compared to the previous one. An interesting result is that the factors 'working conditions' and 'improved job amenities' significantly affect the subjective perception of wage cuts while insignificantly affecting the probability of an objective wage cut. The size of the coefficients, however, is comparable across specifications. It seems plausible that workers with improved job amenities are significantly less likely to perceive worsenings in wages because of general satisfaction with the job which also might result in more satisfaction with wages. Remember that it is not straightforward to interpret the factor score for working conditions because it includes subjective worsenings and improvements of strain and work time regulations. For this reason, I omit interpretation of this factor.

In sum, Tables 4.3 to 4.5 reveal that workers accept lower wages for improved strain. The remaining coefficients are, to the largest extent, imprecisely measured or are not consistent with the hypothesis of trade-off reasoning in mobility decisions. Promotion opportunities are shown to have a robust and highly significant effect on the probability of mobility to lower wages. The estimates, however, reveal no evidence in favor of trade-off reasoning as hypothesized above. The results, furthermore, contradict the ones presented in Pfeifer and Schneck (2010), which might be reasoned by different definitions of the measures for future career prospects. The authors use an objective measure for the change in promotion opportunities, whereas this study applies a subjective measure which depends on individual perceptions. Evidence on the basis of the factor scores (which potentially hide the mechanisms on the less aggregated level) do not support the hypothesis of trade-off reasoning between wages and job amenities in mobility decisions as well. In fact, the coefficients for this particular factor score are estimated to be negative, which suggests that a worker who changes to a new job with better amenities compared to the previous one is less likely to accept wage cuts.

Table 4.4
Probit model on whether workers accepted a wage cut (all subjective comparisons included)

Variables	(1) mobility to lower wages	(2) mobility to lower wages^{real}	(3) subjective worsening of wages
Strain improved	0.0702*	0.0991**	0.0733***
	(0.0375)	(0.0405)	(0.0240)
Strain worsened	-0.0312	-0.0624	0.0132
	(0.0423)	(0.0466)	(0.0262)
Work time improved	-0.0102	-0.0514	0.0341
	(0.0365)	(0.0398)	(0.0210)
Work time worsened	0.0251	0.0252	0.00716
	(0.0534)	(0.0563)	(0.0267)
Security against job loss improved	0.0489	0.0455	0.0138
	(0.0377)	(0.0409)	(0.0186)
Security against job loss worsened	0.0535	-0.0398	-0.0281
	(0.0701)	(0.0671)	(0.0202)
Use of skills improved	-0.0233	-0.0341	-0.0120
	(0.0357)	(0.0393)	(0.0172)
Use of skills worsened	-0.0476	-0.0419	0.00150
	(0.0455)	(0.0521)	(0.0228)
Commuting improved	0.0659*	0.0563	0.0101
	(0.0385)	(0.0422)	(0.0200)
Commuting worsened	-0.00415	0.00282	0.0128
	(0.0389)	(0.0420)	(0.0212)
Chances for promotion improved	-0.0700**	-0.0658*	-0.0751***
	(0.0327)	(0.0369)	(0.0174)
Chances for promotion worsened	0.199**	0.254***	0.141**
	(0.0775)	(0.0777)	(0.0582)
Fringe benefits improved	-0.0591*	-0.0688*	-0.0402**
	(0.0354)	(0.0402)	(0.0169)
Fringe benefits worsened	-0.000826	0.0490	0.156***
	(0.0546)	(0.0624)	(0.0530)
Job in general improved	-0.0322	-0.0142	-0.0202
	(0.0354)	(0.0388)	(0.0178)
Job in general worsened	0.0209	0.0165	-0.00512
	(0.0872)	(0.0945)	(0.0349)
Additional control variables (as in Table 4.3)	yes	yes	yes
Number of observations		800	
Pseudo R^2	0.078	0.059	0.194
Predicted $\Pr(y = 1 \mid \bar{x})$	0.225	0.308	0.056

Marginal effects at $\bar{x}$ are presented.
Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*
All specifications are satisfactory by consideration of the link test because $\hat{y}^2$ is insignificant in all test equations.

Table 4.5
Probit model on whether workers accepted a wage cut (factors included)

Variables	(1) mobility to lower wages	(2) mobility to lower wages^{real}	(3) subjective worsening of wages
Factor score 1	-0.0180	-0.0198	-0.0165*
(working conditions)	(0.0149)	(0.0168)	(0.00978)
Factor score 2	-0.0121	-0.0211	-0.0171**
(job amenities improved)	(0.0155)	(0.0177)	(0.00865)
Factor score 3	-0.0312**	-0.0288*	-0.0204**
(match improved)	(0.0158)	(0.0171)	(0.00940)
Factor score 4	-0.0262*	-0.0194	-0.0102
(commuting)	(0.0149)	(0.0165)	(0.00900)
Factor score 5	0.00930	0.0141	0.0113
(match worsened)	(0.0152)	(0.0170)	(0.00782)
Factor score 6	0.0491***	0.0470***	0.0444***
(job amenities worsened)	(0.0142)	(0.0165)	(0.00767)
Additional control variables (as in Table 4.3)	yes	yes	yes
Number of observations		800	
Pseudo R^2	0.064	0.038	0.127
Predicted $\Pr(y = 1 \mid \bar{x})$	0.229	0.312	0.072

Marginal effects at $\bar{x}$ are presented.
Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*
Specifications (1) and (3) are satisfactory because $\hat{y}^2$ is insignificant in the link test equations. The link test associates a p-value of 0.052 to the coefficient of $\hat{y}^2$ in specification (2).

As the probit model does not tell us something about the magnitude of a possible voluntary earnings loss, I conducted OLS estimation. Similar to the probit approaches, different dependent variables are applied in order to quantify the willingness to pay for amenities. Specifications (1) and (2) in Table 4.6 explain the relative (real) wage change, whereas specifications (3) and (4) refer to the absolute (real) wage change in Euros. The effects of the subjective comparisons in the first specification are measured somewhat imprecisely. With the exception of an improvement in commuting, none of the remaining subjective variables are significant at the 10% level. Examination of Tables 4.3 and 4.6 shows that some results obtained in the probit model on the objective wage cut are consistent. Precisely, Table 4.3 shows that workers who change to a job with subjectively better strain are, on average, more likely to accept a wage cut, and Table 4.6 reveals that workers, on average, pay for better strain by about 1.50% of the hourly wage or by about 45 to 51 Cents when changing jobs. This effect, however, is not statistically significant in the OLS regression because the standard errors are quite sizeable. Note that comparison of Tables 4.3 and 4.6 also averts to some inconsistencies. An example is the worsening in strain. Although workers are shown to be, on average, less likely to pay for worse strain by objective wage cuts (see specifications (1) and (2) in Table 4.3), OLS estimation suggests that workers, on average, suffer wage cuts when changing to a job with worse strain. The only statistically significant effect for the (real) relative wage change in Table 4.6 is estimated for subjective improvements in commuting. The negative sign is in line with the results obtained in the probit model where an improvement in commuting increases the probability for the acceptance of a wage cut. The OLS estimates suggest that better length of commuting to and from work is paid for by, on average, about 5% lower wages compared to the previous job. Trade-off reasoning in the decision to accept earnings losses and improvements in commuting is thus evident. Compensating wage differentials might be indicated by the positive effects of a worsening in security against a job loss and a worsening in work time. In other words, workers are compensated for disamenities such as less job security and less flexible work schedules by higher wages. Note, however, that both effects are statistically insignificant but economically sizable. Some of the effects of the subjective comparisons contradict each other when comparing absolute and the relative wage changes. For example, more flexible work schedules are suggested to be paid for by a wage cut when looking at the relative wage change, whereas a wage markup is indicated for the absolute wage change. Analogously, the problem arises for the coefficients for more security against job loss and a worsening in commuting which are imprecisely measured by consideration of the corresponding standard errors.

Table 4.6 also refers to the absolute wage change of voluntarily mobile individuals in Euros. The effect of the match indicator is significant (see specifications (3) and (4)), whereas an improvement as well as a worsening of the match have a positive effect on the wage by consideration of the (real) absolute wage change as dependent variable. The positive coefficient for an improvement in the match quality can be explained by economic theory, where an improvement in the match quality leads to an increase in wages. The positive effect of a worsening might be reasoned in mobility from a multi-task job to a highly specialized job where workers perceive less use of their skills. Specialists, however, are to the largest extent paid for application of specific knowledge although specialists could also make use of a several different skills. Specification (4), moreover, shows that workers pay for less career advancement opportunities by a significant lower wage of more than one Euro. This effect is statistically insignificant in case of non-deflated wages but economically significant when considering hourly wage changes. The control variables in Table 4.6 reveal that homeowners are more likely to change to lower wages compared to renters. Precisely, homeowners, ceteris paribus, accept an average wage cut of at least 4.20% or a minimum of 63 Cents when compared to renters. The growth in unemployment does not significantly affect the wage change.[16]

It might be argued that outliers affect the results presented in Table 4.6. For this reason, I excluded wage changes above the 90% percentile and below the 10% percentile. The exact values for the cutoff of the relative wage change are shown in Table 4.1.[17] The results for the trimmed sample are presented in Table 4.7. Note that the specifications differ in the number of individuals. Application of the trimmed sample reveals that the only non-robust coefficient across specifications is the dummy variable for improved strain because specification (3) suggests a positive effect, whereas the other specifications reveal a negative impact while the remaining coefficients have the same sign for either (real) relative wage changes as well as (real) absolute wage changes.

[16] Note that regression with annual dummies basically confirms the results shown in Table 4.6 in terms of significance and sign.

[17] Descriptive statistics for the (real) absolute wage change are presented in this footnote. On average, workers gain about 1.41 to 1.54 Euro when changing jobs voluntarily.

Descriptive statistics for absolute wage change

Variable	10% Percentile	Mean	90% Percentile	Observations
$w_{ijt} - w_{i,j-1,t-1}$	-1.7827	1.5411	5.4003	800
		(4.5613)		
$w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$	-2.2389	1.4123	5.5858	800
		(4.8488)		

Standard deviations in parentheses

Table 4.6
OLS estimation results for wage change

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$
Strain improved	-0.0149	-0.0152	-0.446	-0.506
	(0.0250)	(0.0246)	(0.328)	(0.350)
Strain worsened	-0.00483	-0.00597	-0.264	-0.343
	(0.0361)	(0.0353)	(0.518)	(0.547)
Work time improved	-0.0196	-0.0182	0.219	0.251
	(0.0255)	(0.0250)	(0.350)	(0.370)
Work time worsened	0.0264	0.0269	0.740	0.851
	(0.0461)	(0.0450)	(0.541)	(0.583)
Security against job loss improved	0.0144	0.0141	-0.125	-0.133
	(0.0246)	(0.0242)	(0.311)	(0.329)
Security against job loss worsened	0.0666	0.0661	0.643	0.698
	(0.0522)	(0.0514)	(0.628)	(0.666)
Use of skills improved	0.0403	0.0393	0.700*	0.759*
	(0.0268)	(0.0263)	(0.374)	(0.396)
Use of skills worsened	0.0331	0.0327	0.687*	0.728*
	(0.0330)	(0.0325)	(0.401)	(0.420)
Commuting improved	-0.0509*	-0.0497*	-0.832*	-0.846*
	(0.0278)	(0.0273)	(0.445)	(0.468)
Commuting worsened	0.00482	0.00486	-0.0664	-0.0556
	(0.0342)	(0.0336)	(0.423)	(0.457)
Chances for promotion improved	0.00804	0.00792	0.256	0.215
	(0.0264)	(0.0259)	(0.385)	(0.406)
Chances for promotion worsened	-0.0712	-0.0703	-1.015	-1.121*
	(0.0517)	(0.0510)	(0.618)	(0.658)
Homeowner	-0.0435*	-0.0420*	-0.635*	-0.709*
	(0.0223)	(0.0219)	(0.366)	(0.384)
Number of previous individual mobility	-0.0188*	-0.0187*	-0.0781	-0.0951
	(0.0107)	(0.0105)	(0.164)	(0.172)
Age	-0.0185	-0.0179	-0.102	-0.102
	(0.0133)	(0.0130)	(0.173)	(0.181)
Age^2	0.000187	0.000179	0.000967	0.000887
	(0.000172)	(0.000169)	(0.00246)	(0.00256)
Education in years	0.00837	0.00819	0.256***	0.258***
	(0.00510)	(0.00501)	(0.0703)	(0.0732)
Blue-collar to white-collar transition	-0.0315	-0.0299	-0.424	-0.368
	(0.0371)	(0.0365)	(0.422)	(0.447)
Male	0.00480	0.00486	0.357	0.344
	(0.0312)	(0.0306)	(0.339)	(0.368)
Partner	-0.000611	-0.000110	0.00415	0.0191
	(0.0385)	(0.0377)	(0.406)	(0.440)
Firm more than 2,000 workers	-0.00558	-0.00549	-0.0766	-0.133
	(0.0313)	(0.0309)	(0.553)	(0.578)
Previous firm more than 2,000 workers	-0.00363	-0.00384	-0.271	-0.362
	(0.0415)	(0.0410)	(0.599)	(0.637)
Regional mobility	0.0934	0.0923	0.397	0.415
	(0.0755)	(0.0740)	(0.852)	(0.896)
Growth in unemployment rate	-0.0120	-0.0113	-0.244	-0.219
	(0.0144)	(0.0141)	(0.189)	(0.196)
Constant	1.493***	1.466***	0.612	0.652
	(0.245)	(0.240)	(2.802)	(2.948)
R^2	0.0500	0.0500	0.0582	0.0561
Number of observations	800			

Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Table 4.7
OLS estimation results for wage change (trimmed sample)

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Strain improved	-0.00513	-0.00761	0.00541	-0.0444
	(0.0125)	(0.0123)	(0.146)	(0.157)
Strain worsened	0.00907	0.0125	0.128	0.153
	(0.0154)	(0.0155)	(0.175)	(0.189)
Work time improved	0.0191	0.0182	0.000151	0.129
	(0.0122)	(0.0121)	(0.148)	(0.157)
Work time worsened	-0.00588	-0.00761	-0.180	-0.0929
	(0.0159)	(0.0159)	(0.190)	(0.201)
Security against job loss improved	0.00857	0.00689	0.158	0.125
	(0.0120)	(0.0118)	(0.143)	(0.152)
Security against job loss worsened	0.0183	0.0165	0.545*	0.447
	(0.0254)	(0.0249)	(0.288)	(0.312)
Use of skills improved	0.0311***	0.0276**	0.350**	0.348**
	(0.0118)	(0.0116)	(0.139)	(0.150)
Use of skills worsened	0.0131	0.0126	0.0729	0.0742
	(0.0175)	(0.0173)	(0.200)	(0.215)
Commuting improved	-0.0346**	-0.0313**	-0.325**	-0.335*
	(0.0140)	(0.0138)	(0.160)	(0.171)
Commuting worsened	-0.0188	-0.0167	-0.248	-0.218
	(0.0138)	(0.0136)	(0.159)	(0.172)
Chances for promotion improved	0.0151	0.0174	0.164	0.0989
	(0.0123)	(0.0120)	(0.145)	(0.155)
Chances for promotion worsened	-0.0559**	-0.0565**	-0.656**	-0.757**
	(0.0225)	(0.0221)	(0.301)	(0.321)
Homeowner	-0.00716	-0.00556	-0.0933	-0.0557
	(0.0118)	(0.0117)	(0.135)	(0.149)
Number of previous individual mobility	-0.000903	0.000695	0.0191	0.000632
	(0.00546)	(0.00531)	(0.0728)	(0.0777)
Age	-0.00699	-0.00731	-0.0390	-0.0432
	(0.00549)	(0.00544)	(0.0646)	(0.0690)
Age^2	7.53e-05	8.18e-05	0.000419	0.000433
	(7.44e-05)	(7.37e-05)	(0.000889)	(0.000951)
Education in years	0.000542	0.000791	0.0384	0.0291
	(0.00244)	(0.00241)	(0.0293)	(0.0315)
Blue-collar to white-collar transition	0.00784	0.00995	0.229	0.302
	(0.0226)	(0.0221)	(0.263)	(0.285)
Male	0.0171	0.0164	0.345***	0.410***
	(0.0115)	(0.0113)	(0.129)	(0.139)
Partner	-0.00364	-0.00194	-0.0348	-0.0345
	(0.0136)	(0.0135)	(0.156)	(0.167)
Firm more than 2,000 workers	0.0113	0.00710	0.267	0.278
	(0.0164)	(0.0160)	(0.199)	(0.218)
Previous firm more than 2,000 workers	0.00189	0.00200	-0.00536	0.0363
	(0.0171)	(0.0168)	(0.200)	(0.219)
Regional mobility	0.0435	0.0431	1.109***	1.135***
	(0.0293)	(0.0293)	(0.370)	(0.386)
Growth in unemployment rate	-0.00605	-0.00550	-0.0649	-0.0304
	(0.00676)	(0.00674)	(0.0810)	(0.0867)
Constant	1.235***	1.215***	1.308	1.387
	(0.0959)	(0.0950)	(1.114)	(1.178)
R^2	0.0606	0.0562	0.0848	0.0766
Number of observations	640			
Individuals	544	545	551	550

Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Consideration of the robustness across the different samples reveals that some coefficients are non-robust because the effects in Table 4.7 have the opposite sign compared to the ones presented in Table 4.6. An example is the worsening in strain, where the estimated coefficient in Table 4.7 reveals a negative effect, whereas workers are, on average, compensated for an increase in this disamenity by higher wages in Table 4.6. Analogous argumentation holds for less flexible work schedules. In brief, trimming the sample has considerable consequences for some of the coefficients when compared to application of the full sample. Some effects, however, remain robust when trimming the sample. Tables 4.6 and 4.7 reveal a compensating wage differential for a worsening in perceived job security. In other words, workers get a wage markup in exchange for lower security against a job loss. This effect is robust, however, it is insignificant in most of the specifications. I also found that better application of individual skills in the current job compared to the previous one increases wages significantly. Workers changing to jobs with better promotion opportunities in the current job compared to the last one are suggested to earn, on average, (insignificantly) higher wages, whereas workers who change to jobs with fewer career prospects, in turn, suffer sizable wage cuts. A possible interpretation is that mobile individuals who accept fewer future career prospects in the new job are *double losers* who suffer not only lower wages but also fewer career prospects in the new job compared to the last one.

Next, I checked the robustness of the results presented in Table 4.6 by consideration of all subjective comparisons between the current and the previous jobs as well as by consideration of the six factor scores. Tables 4.8 and 4.9 present the corresponding results. Table 4.8 shows that the (real) relative wage change is significantly affected by improvements in commuting and better fringe benefits. The coefficient of improved commuting is significantly negative as in Tables 4.6 and 4.7 and, thus, is highly robust across different specifications. Interpretation of this effect is in line with trade-off reasoning where individuals pay for better commuting by lower wages. The effects are comparable in size to the ones obtained in Table 4.6 and somewhat larger than for the ones presented in Table 4.7. An improvement in fringe benefits increases wages by about 4.8% to 4.9%. Specification (4) in Table 4.8 reveals a statistically significant compensating wage differential for less flexible work time in the current job compared to the last one because worse work schedules increase wages by about 96 Cents. The match indicator is statistically significant and positive as in Table 4.6 whereas specialization might provide an interpretation for this result.

Table 4.8
OLS estimation results for wage change (all subjective comparisons included)

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Strain improved	-0.0168	-0.0171	-0.489	-0.550
	(0.0247)	(0.0243)	(0.327)	(0.348)
Strain worsened	-0.00287	-0.00399	-0.241	-0.318
	(0.0366)	(0.0358)	(0.522)	(0.552)
Work time improved	-0.0311	-0.0294	0.115	0.137
	(0.0248)	(0.0244)	(0.371)	(0.389)
Work time worsened	0.0312	0.0318	0.846	0.960*
	(0.0471)	(0.0461)	(0.539)	(0.580)
Security against job loss improved	-0.000429	-0.000471	-0.240	-0.261
	(0.0265)	(0.0260)	(0.356)	(0.380)
Security against job loss worsened	0.0771	0.0770	0.837	0.899
	(0.0619)	(0.0609)	(0.665)	(0.711)
Use of skills improved	0.0434	0.0425	0.657*	0.722*
	(0.0271)	(0.0267)	(0.373)	(0.393)
Use of skills worsened	0.0376	0.0371	0.745*	0.791*
	(0.0334)	(0.0328)	(0.417)	(0.434)
Commuting improved	-0.0485*	-0.0473*	-0.805*	-0.816*
	(0.0278)	(0.0273)	(0.452)	(0.475)
Commuting worsened	0.00634	0.00632	-0.0632	-0.0504
	(0.0348)	(0.0342)	(0.427)	(0.462)
Chances for promotion improved	0.00590	0.00588	0.220	0.178
	(0.0264)	(0.0259)	(0.384)	(0.406)
Chances for promotion worsened	-0.0682	-0.0671	-0.913	-1.017
	(0.0552)	(0.0544)	(0.622)	(0.665)
Fringe benefits improved	0.0494*	0.0484*	0.383	0.428
	(0.0266)	(0.0262)	(0.400)	(0.423)
Fringe benefits worsened	-0.0147	-0.0158	-0.431	-0.435
	(0.0582)	(0.0568)	(0.562)	(0.608)
Job in general improved	-0.0122	-0.0124	0.0573	0.0381
	(0.0272)	(0.0268)	(0.365)	(0.384)
Job in general worsened	-0.0342	-0.0336	-0.511	-0.546
	(0.0750)	(0.0740)	(0.767)	(0.807)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes
R^2	0.0542	0.0542	0.0608	0.0587
Number of observations		800		

Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Before turning the focus to the effect of the factor scores, I calculated the wage change based on the logarithm of the hourly net wage in Euros in order to compare the results to the ones obtained in Villanueva (2007). It is important to note that the net wage is not directly bargained between employer and employee because it depends, among a variety of aspects, on marital status or confession. I exclusively concentrate on observations which are already included into the above analysis, and because of missing values on the net wages this sample contains fewer observations. It is also important to note that Villanueva (2007) examines a different time horizon (from 1984 to 2001) and that both studies consider a somewhat different set of control variables. The results, however, can be compared with each other because of the concentration on the effects of the same subjective comparison variables between jobs on the wage change of voluntary mobility in Germany. Specification (1) and (2) in Table 4.9 contradict Villanueva (2007) in subjective improvements as well as worsenings in strain because the signs are reversed in both studies. After trimming the sample at the 10% and 90% percentile (see specifications (3) and (4)), the effect of an improvement in strain is similar in both studies, whereas the effect of worsenings in strain remain contradictory. The highly significant and positive coefficient for less use of skills after mobility in Villanueva (2007) cannot be confirmed in this paper. Comparison of the effects presented in Tables 4.8 and 4.9 in this chapter reveals that most effects advert to similar directions of the effects. Note, again, that most of the estimated coefficients are insignificant. The most important inconsistency can can be found in the coefficients for an improvement in strain because Table 4.8 adverts to trade-off reasoning while Table 4.9 reveals a positive effect in specifications (1) and (2).

Table 4.9
OLS estimation results for wage change (alternative dependent variables)

Variables	(1) $log(w_{ijt}^{net})-$ $log(w_{i,j-1,t-1}^{net})$	(2) $log(w_{ijt}^{net,real})-$ $log(w_{i,j-1,t-1}^{net,real})$	(3) $log(w_{ijt}^{net})-$ $log(w_{i,j-1,t-1}^{net})$ (trimmed)	(4) $log(w_{ijt}^{net,real})-$ $log(w_{i,j-1,t-1}^{net,real})$ (trimmed)
Strain improved	0.00299	0.00237	-0.0235**	-0.0236**
	(0.0233)	(0.0233)	(0.0111)	(0.0110)
Strain worsened	-0.0163	-0.0169	-0.0115	-0.0133
	(0.0233)	(0.0233)	(0.0132)	(0.0133)
Work time improved	-0.00621	-0.00550	0.0152	0.0120
	(0.0197)	(0.0197)	(0.0111)	(0.0111)
Work time worsened	0.0636*	0.0640*	-0.00214	0.000990
	(0.0343)	(0.0343)	(0.0147)	(0.0151)
Security against job loss improved	-0.0320	-0.0319	-0.00885	-0.00718
	(0.0208)	(0.0207)	(0.0109)	(0.0108)
Security against job loss worsened	0.0153	0.0154	0.0276	0.0234
	(0.0347)	(0.0346)	(0.0213)	(0.0211)
Use of skills improved	0.0212	0.0211	0.0182	0.0170
	(0.0212)	(0.0212)	(0.0112)	(0.0112)
Use of skills worsened	0.0133	0.0138	-0.00904	-0.00931
	(0.0247)	(0.0247)	(0.0150)	(0.0151)
Commuting improved	-0.0554**	-0.0547**	-0.0197*	-0.0162
	(0.0217)	(0.0217)	(0.0116)	(0.0116)
Commuting worsened	-0.0251	-0.0248	-0.00790	-0.00603
	(0.0252)	(0.0252)	(0.0117)	(0.0117)
Chances for promotion improved	0.00492	0.00463	0.00766	0.00506
	(0.0198)	(0.0198)	(0.00989)	(0.00989)
Chances for promotion worsened	-0.0601	-0.0602	0.000610	-0.00606
	(0.0430)	(0.0429)	(0.0232)	(0.0228)
Fringe benefits improved	0.0373*	0.0369*	0.0220*	0.0212*
	(0.0215)	(0.0215)	(0.0115)	(0.0115)
Fringe benefits worsened	-0.0471	-0.0473	-0.0291*	-0.0340**
	(0.0294)	(0.0295)	(0.0155)	(0.0155)
Job in general improved	-0.00808	-0.00829	0.00351	0.00552
	(0.0198)	(0.0198)	(0.0101)	(0.0102)
Job in general worsened	-0.00383	-0.00402	-0.0205	-0.0177
	(0.0492)	(0.0491)	(0.0218)	(0.0218)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes
R^2	0.074	0.0733	0.0940	0.090
Number of observations	755	755	605	605

Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Table 4.10 presents the results wherein the factor scores instead of dummy variables for subjective worsenings and improvements in job characteristics are considered. An improvement in the match quality increases wages by about 1.5%, and this effect is statistically insignificant. Specifications (3) and (4) reveal that an improvement in the match quality contributes to a statistically significant increase in wages of about 30 Cents. Commuting exhibits a significant impact on wages throughout all specifications.

As discussed above, improvements in commuting lead to a reduction of the corresponding factor score. For this reason, better commuting decreases wages, whereas subjectively more commuting expenditures are compensated for by higher wages. Table 4.5 suggests that less job amenities are more likely to be accompanied by wage cuts. This pattern is generally confirmed in Table 4.10 where the corresponding coefficient is negative but, however, it tends to be of small economic importance.

Table 4.10

OLS estimation results for wage change (factors included)

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Factor score 1	0.0155	0.0149	0.175	0.187
(working conditions)	(0.0131)	(0.0129)	(0.157)	(0.165)
Factor score 2	0.00322	0.00330	0.00497	0.00203
(job amenities improved)	(0.0127)	(0.0124)	(0.144)	(0.154)
Factor score 3	0.0150	0.0146	0.296*	0.304
(match improved)	(0.0147)	(0.0144)	(0.178)	(0.192)
Factor score 4	0.0275*	0.0269*	0.358**	0.370**
(commuting)	(0.0141)	(0.0138)	(0.175)	(0.188)
Factor score 5	-2.35e-05	0.000165	0.0113	0.00915
(match worsened)	(0.0118)	(0.0116)	(0.138)	(0.145)
Factor score 6	-0.00717	-0.00702	-0.136	-0.136
(job amenities worsened)	(0.0143)	(0.0141)	(0.167)	(0.178)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes
R^2	0.0439	0.0439	0.0485	0.0456
Number of observations	800			

Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Further robustness checks are summarized in Table A4.4 in the appendix. The table presents OLS estimation results for different subsamples, taking into consideration the variables strain, work time, security against job loss, commuting, promotion chances, and the match indicator. The robustness checks divide the sample by homeowners and renters, by age (below and above median age), and by different phases of the business cycle (growth of unemployment smaller or larger than zero). The dummy variable for homeowners in Table 4.6 suggests that homeowners are significantly worse off when compared to renters with identical characteristics. Estimation by homeowner and renter reveals that owners pay for improvements in commuting by an average of more than 9% lower wages while renters only accept wage cuts of, on average, about 3% when changing jobs voluntarily. Although both coefficients are economically significant, only the effect for homeowners is statistically significant (at the 5% level). The results also suggest that renters are willing to pay for improvements in strain (partly significant) and for better work time regulations (insignificant) compared to homeowners, who sacrifice large wage gains for such improvements. When focusing on different phases of the business cycle, the main result is that mobility during downturns ($\Delta u > 0$) is characterized by significant wage gains if an improvement in the use of skills can be achieved. The relative wage gain of better application of skills amounts to an average of more than 11%, whereas the average absolute wage gain is 1.56 to 1.71 Euros for better use of individual skills. The coefficient for worsenings in promotion opportunities is estimated to be significantly negative during upswings ($\Delta u \leq 0$). A possible interpretation might be that mobility to worse future career prospects during booms signals low own career ambitions, where employers impose a penalty for this type of signal. A very interesting result can be obtained when the sample is divided into workers who are at least 34 years old and workers who are younger than the median age. In particular, young workers obtain an economically as well as statistically significant compensating wage differential for less job security against a job loss, whereas older workers, on average, even suffer wage cuts when being mobile voluntarily to less secure jobs. Investigation of the standard errors of less job security for workers above 34 years of age, however, reveals that this coefficient is imprecisely measured.

In Table A4.5, I compared the results obtained for different time horizons (1998-2004 and 1999-2003). An improvement in strain is paid for by wage cuts in specifications (1) and (2) by consideration of the time horizon from 1998 to 2004 while a positive coefficient is estimated for the period from 1999 to 2003. Similar results are obtained for an improvement in perceived security against job loss. Specifications (3) and (4) are also sensitive to different time horizons (see work time improved and use of skills improved). An investigation of Table A4.5 reveals evidence that poolability over time is problematic.

Finally, I conducted tobit (corner solution) estimation in order to explain the willingness to pay for improvements in job-specific amenities, given the probability that the individual changes to lower wages. As a consequence, interpretation of the marginal effects in Table 4.11 is based on the condition that workers changed jobs voluntarily to lower wages. The table reveals that workers significantly pay for improvements in strain. The corresponding marginal effects suggest that workers with wage cuts pay for better workload by an average of about 1.2% or 29 to 32 Cent, respectively. For worse strain, a statistically insignificant as well as economically small moderating effect on wage cuts is estimated. As above, the tobit model confirms positive effects for both indicators of the match quality. The effects of promotion opportunities are also robust to the results in the OLS regressions. Better promotion opportunities mitigate wage cuts while worse career prospects in the new job compared to the previous one increase wage cuts. It is also confirmed that workers seem to pay for less commuting expenses by earnings losses. One can find only few inconsistencies or non-robust coefficients when utilizing the full specification (see Table A4.6). The effects of a worsening in work time in specifications (1) and (2) as well as the coefficient for worse commuting expenses in specification (2) are non-robust compared to the results shown in Table 4.11. Note, however, that these effects are statistically insignificant and of small economic importance in both tables.

Table 4.11

Tobit regression results for wage cut, given that individuals change to lower wages

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Strain improved	-0.0119*	-0.0126**	-0.289*	-0.321**
	(0.00639)	(0.00566)	(0.153)	(0.143)
Strain worsened	0.00249	0.00400	0.0223	0.0589
	(0.00739)	(0.00681)	(0.207)	(0.204)
Work time improved	0.00379	0.00683	0.194	0.282*
	(0.00605)	(0.00546)	(0.157)	(0.155)
Work time worsened	-0.000116	-0.000629	0.0593	0.0518
	(0.00780)	(0.00676)	(0.195)	(0.178)
Security against job loss improved	-0.00662	-0.00611	-0.112	-0.102
	(0.00576)	(0.00524)	(0.137)	(0.131)
Security against job loss worsened	-0.00143	0.00532	-0.0282	0.154
	(0.00857)	(0.00773)	(0.212)	(0.203)
Use of skills improved	0.00799	0.00753	0.283*	0.284*
	(0.00571)	(0.00519)	(0.155)	(0.146)
Use of skills worsened	0.0125*	0.0115*	0.332**	0.332**
	(0.00691)	(0.00617)	(0.166)	(0.152)
Commuting improved	-0.0118*	-0.00982*	-0.305	-0.265
	(0.00635)	(0.00569)	(0.195)	(0.182)
Commuting worsened	0.000928	-3.97e-05	0.00465	-0.0211
	(0.00621)	(0.00557)	(0.157)	(0.148)
Chances for promotion improved	0.0114**	0.00982*	0.180	0.144
	(0.00559)	(0.00513)	(0.146)	(0.141)
Chances for promotion worsened	-0.0303***	-0.0315***	-0.706**	-0.762***
	(0.0116)	(0.0111)	(0.283)	(0.274)
Homeowner	-0.00973*	-0.00678	-0.324**	-0.260*
	(0.00545)	(0.00500)	(0.161)	(0.150)
Number of previous individual	-0.00110	-0.00209	-0.0142	-0.0460
mobility	(0.00282)	(0.00265)	(0.0714)	(0.0685)
Age	-0.00428*	-0.00140	-0.115**	-0.0416
	(0.00256)	(0.00227)	(0.0581)	(0.0521)
Age^2	3.83e-05	3.61e-06	0.00107	0.000183
	(3.43e-05)	(3.06e-05)	(0.000762)	(0.000703)
Education in years	0.00352***	0.00237**	0.0628**	0.0362
	(0.00112)	(0.000992)	(0.0274)	(0.0254)
Blue-collar to white-collar transition	-0.00271	-0.00134	0.0367	0.0896
	(0.0115)	(0.0109)	(0.246)	(0.236)
Male	-3.39e-05	0.00199	-0.0396	0.00864
	(0.00531)	(0.00484)	(0.132)	(0.126)
Partner	-0.000874	-0.000211	-0.00978	0.00527
	(0.00638)	(0.00568)	(0.153)	(0.143)
Firm more than 2,000 workers	-0.00564	-0.00545	-0.199	-0.201
	(0.00744)	(0.00672)	(0.237)	(0.228)
Previous firm more than 2,000 workers	-0.00333	-0.00537	-0.151	-0.228
	(0.00873)	(0.00790)	(0.259)	(0.251)
Regional mobility	-0.00281	0.000558	0.00295	0.105
	(0.0158)	(0.0151)	(0.362)	(0.357)
Growth in unemployment rate	-0.00326	-0.00134	-0.0776	-0.0324
	(0.00317)	(0.00284)	(0.0774)	(0.0717)
Pseudo R^2	0.129	0.130	0.0332	0.0247
Number of observations		800		
Uncensored observations	195	254	195	254
Censored observations	605	546	605	546

Marginal effects after tobit regression.
Robust standard errors clustered for 670 individuals in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.*

Table 4.12

Tobit regression results for wage cut, given that individuals change to lower wages (factor scores)

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$
Factor score 1	0.00352	0.00297	0.0575	0.0478
(working conditions)	(0.00242)	(0.00225)	(0.0569)	(0.0555)
Factor score 2	4.03e-05	0.000575	0.0272	0.0446
(job amenities improved)	(0.00249)	(0.00237)	(0.0592)	(0.0589)
Factor score 3	0.00512*	0.00425*	0.151**	0.132*
(match improved)	(0.00267)	(0.00238)	(0.0752)	(0.0690)
Factor score 4	0.00533**	0.00402*	0.121*	0.0905
(commuting)	(0.00251)	(0.00230)	(0.0716)	(0.0668)
Factor score 5	-0.000386	-0.000560	-0.0128	-0.0138
(match worsened)	(0.00236)	(0.00217)	(0.0540)	(0.0514)
Factor score 6	-0.00662***	-0.00551***	-0.129**	-0.102**
(job amenities worsened)	(0.00215)	(0.00206)	(0.0507)	(0.0509)
Additional control variables (as in Table 4.11)	yes	yes	yes	yes
Pseudo R^2	0.107	0.094	0.027	0.017
Number of observations	800			
Uncensored observations	195	254	195	254
Censored observations	605	546	605	546

Marginal effects after tobit regression.
Robust standard errors clustered for 670 individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Table 4.12 presents the results for the tobit model when considering the factor scores. Workers changing to lower wages, on average, pay for less commuting expenses by lower wages. The factor score, however, reveals an economically small effect because, on average, less than 1% or an average maximum of 12 Cent are paid for the improvement in commuting. This effect is considerably smaller compared to the one presented in Table 4.10 but reveals robustness of this particular coefficient. A further similarity to the results in the OLS regressions is the moderating effect on wage cuts if the match quality in the current job is better than in the previous job. A worsening in job amenities is significantly paid for by lower wages. This result is highly robust when compared to Table 4.10. Albeit highly statistically significant, the effect is of small economic significance. It might be argued that workers changing to worse job amenities are changing to some sort of low-pay sector with dead-end jobs, low job stability (or low job security), and low (or inexistent) fringe benefits. This result is also consistent with the ”segmented labor market” in Villanueva (2007), where wage penalties are attached to job-specific disamenities.

An analysis for different subgroups is displayed in Table A4.7. The tobit approach reveals a much smaller negative impact of commuting on the acceptance of wage cuts, given that the individual changes to lower wages. This general result is reflected by consideration of different samples for homeowners and renters. The coefficients for less commuting expenses in Table A4.7 are much smaller than the ones presented in Table A4.4. The statistical significance of this particular effect, furthermore, is not given any longer for the subgroup of homeowners, whereas the coefficient becomes statistically significant for renters. It is, however, confirmed that renters pay for improvements in strain, while better workload mitigates wage cuts of homeowners. Consideration of different phases of the business cycle leads to statistical insignificance for almost all of the coefficients. The finding of a negative effect of mobility to worse career prospects in the new job compared to the previous one is confirmed across different phases of the business cycle. Similar to Table A4.4, Table A4.7 reveals alleviation for worse security against a job loss for the group of workers who are younger than the median age compared to the group of older workers.

To sum up, trade-off reasoning, as hypothesized above, is a key feature of the acceptance of wage cuts. Especially improved commuting is found to be paid for by lower wages. There is also weak evidence for compensating wage differentials for worse commuting to and from the workplace. This implies that the expenses for changes in commuting and changes in wages are positively correlated, whereas workers with higher commuting expenses compared to the ones in the previous job are paid for this disamenity, while workers with lower commuting expenses accept wage cuts when changing jobs. There is also some evidence in favor for individual trade-off reasoning between wages and improved strain. I am, however, not able to find distinct support for the hypothesis that workers trade off improvements in work time arrangements, better security against job loss, and the acceptance of lower wages. There is weak (mostly statistically insignificant) evidence in favor of compensating wage differentials for less security against a job loss. In addition, the hypothesis that workers pay for better career prospects by wage cuts cannot be supported in this paper. In fact, the reverse is suggested because individuals are not compensated for worse promotion opportunities by higher wages. The findings on some of the job-specific amenities differ when considering subjective perceptions about worsenings in wages instead of using objective measures for wage cuts. This might be driven by cognitive dissonance reduction where workers adjust their perceptions about the job in a positive way to resolve cognitive dissonance introduced by mobility to lower wages.

4.5 Discussion

This chapter investigates the relationship between subjective improvements between two jobs and voluntary mobility to lower wages. This allows to assess the impact of trade-off reasoning in individual labor market decisions. The results suggest that job-specific (non-wage) amenities affect the job choice. More specifically, workers are shown to voluntarily accept wage cuts when improvements in commuting expenses or subjectively better strain can be achieved. Note that commuting expenses are predictable before changing employer. In fact, economic uncertainty is almost inexistent because the new firm's location and the frequency of commuting is known beforehand. Because workers are sure about the improvements in this particular job characteristic, they are even willing to accept wage cuts. The loss of utility through decreasing wages is, thus, compensated for by an increase in utility through improvements in job-specific amenities in the new job.

The results also have important implications for employers. Offering non-wage amenities can attract workers of competitors who pay higher wages. This implies that those employers who offer, for example, activities to decrease job-specific strain are suggested to attract employees of competitors despite lower wages. In addition, employers are able to attract workers by locating companies in regions where workers have low costs of commuting. This also includes the time component of commuting to the workplace. It is also suggested that the absence of compensating wage differentials can be explained by such non-wage amenities. Since Schneck (2010) showed that transitions to permanently lower wages are common, it might be hypothesized that workers trade off permanent lower wages with subjective improvements in certain job-specific characteristics. This study shows that commuting expenses are a potential candidate for the acceptance of downward mobility.

4.6 Appendix

Table A4.1
Frequencies of subjective comparisons between old and new job

How would you judge your present position compared to your last one? In what ways has it improved, stayed the same, or worsened			
Variable	improved	stayed the same	worsened
Length of commute to and from work	280	274	246
Work load (strain)	269	367	164
Work schedule regulations (work time)	349	324	127
Security against job loss	276	471	53
Chances for promotion	348	402	50
General Job type	461	306	33
Fringe benefits	280	428	92
Wages	521	203	76
Are you able to use your professional skills and abilities today more, about the same, or less than in your previous position?			
	more (improved)	about the same (stayed the same)	less (worsened)
Use of skills	328	370	102
Number of observations		800	

Table A4.2

Descriptive statistics of the control variables

	Mean	Standard Deviation
Subjective improvement in		
Work load (strain)	0.3363	0.4727
Work schedule regulations (work time)	0.4363	0.4962
Security against job loss	0.3450	0.4757
Use of skills	0.4100	0.4921
Commuting	0.3500	0.4773
Chances for promotion	0.4350	0.4961
Fringe benefits	0.3500	0.4773
Job type	0.5763	0.4945
Subjective worsening in		
Work load (strain)	0.2050	0.4040
Work schedule regulations (work time)	0.1588	0.3657
Security against job loss	0.0663	0.2489
Use of skills	0.1275	0.3337
Commuting	0.3075	0.4617
Chances for promotion	0.0625	0.2422
Fringe benefits	0.1150	0.3192
Job type	0.0413	0.1990
Dummy variable for homeowners	0.3063	0.4612
Number of individual quits	1.7500	0.8992
Age	35.0725	7.9551
Age^2	1293.2850	593.8163
Education (in years of schooling)	12.7719	2.5192
Dummy variable for blue-collar to white-collar	0.0575	0.2329
Dummy variable for males	0.6475	0.4780
Dummy variable for partner	0.2125	0.4093
Dummy variable for workforce$_{ijt}$>2,000	0.1713	0.3770
Dummy variable for workforce$_{i,j-1,t-1}$>2,000	0.1325	0.3392
Dummy variable for regional mobility	0.0475	0.2128
Growth in unemployment rate	-0.0571	0.7854
Number of observations	800	
Number of individuals	670	

Table A4.3

Spearman's correlation and Tetrachoric correlations for binary variables

	strain		work time		job security		use of skills		commuting		promotion possibilities		fringe benefits		job type		wage	
	↑	↓	↑	↓	↑	↓	↑	↓	↑	↓	↑	↓	↑	↓	↑	↓	↑	↓
strain ↑	1.000																	
strain ↓	-0.361*	1.000																
	-1.000*																	
work time ↑	0.334*	-0.197*	1.000															
	0.514*	-0.359*																
work time ↓	-0.150*	0.313*	-0.382*	1.000														
	-0.315*	0.537*	-1.000*															
job security ↑	0.140*	-0.082*	0.199*	-0.063	1.000													
	0.228*	-0.154*	0.316*	-0.127														
job security ↓	-0.019	0.089*	-0.052	0.104*	-0.193*	1.000												
	-0.049	0.216*	-0.129	0.255*	-1.000*													
use of	0.031	0.030	0.061	0.041	0.122*	-0.059	1.000											
skills ↑	0.050	0.054	0.097	0.078	0.197*	-0.148												
use of	0.117*	-0.008	0.117*	0.029	0.030	0.079*	-0.319*	1.000										
skills ↓	0.229*	-0.019	0.231*	0.067	0.061	0.207*	-1.000*											
commuting ↑	0.099*	-0.061	0.195*	-0.118*	0.019	-0.016	0.022	0.081*	1.000									
	0.162*	-0.114	0.308*	-0.241*	0.031	-0.041	0.036	0.161*										
commuting ↓	-0.062	0.118*	-0.045	0.133*	0.029	0.084*	0.072*	0.005	-0.489*	1.000								
	-0.106	0.210*	-0.075	0.248*	0.049	0.199*	0.119*	0.011	-1.000*									
promotion	0.091*	0.054	0.047	0.060	0.164*	-0.102*	0.268*	-0.063	-0.025	0.104*	1.000							
possibilities ↑	0.147*	0.096	0.074	0.114	0.262*	-0.260*	0.412*	-0.129	-0.041	0.170*								
promotion	-0.020	0.022	0.023	0.043	-0.046	0.201*	-0.068	0.134*	-0.005	0.041	-0.227*	1.000						
possibilities ↓	-0.051	0.060	0.056	0.118	-0.122	0.471*	-0.177	0.324*	-0.014	0.101	-1.000*							
fringe	0.177*	-0.100*	0.337*	-0.125*	0.416*	-0.122*	0.076*	0.042	0.022	0.011	0.154*	-0.005	1.000					
benefits ↑	0.285*	-0.188*	0.516*	-0.256*	0.618*	-0.346*	0.122*	0.084	0.036	0.018	0.246*	-0.014						
fringe	-0.124*	0.157*	-0.151*	0.240*	-0.122*	0.345*	-0.014	0.015	0.007	-0.003	-0.048	0.150*	-0.265*	1.000				
benefits ↓	-0.287*	0.316*	-0.326*	0.460*	-0.278*	0.650*	-0.029	0.038	0.014	-0.005	-0.010	0.359*	-1.000*					
job type ↑	0.134*	-0.003	0.117*	0.006	0.112*	-0.016	0.406*	-0.074*	0.030	0.045	0.252*	-0.061	0.078*	0.016	1.000			
	0.219*	-0.006	0.185*	0.011	0.182*	-0.038	0.617*	-0.147*	0.049	0.075	0.394*	-0.148	0.127*	0.033				
job type ↓	-0.015	0.082*	-0.018	0.185*	0.035	0.096*	-0.160*	0.279*	0.006	0.066	-0.068	0.232*	0.032	0.043	-0.242*	1.000		
	-0.043	0.226*	-0.051	0.451*	0.097	0.295*	-0.610*	0.607*	0.017	0.182	-0.201	0.544*	0.091	0.140	-1.000*			
wage ↑	-0.051	0.047	-0.002	0.045	0.079*	-0.058	0.071*	-0.035	-0.057	0.016	0.171*	-0.125*	0.130*	-0.123*	0.121*	-0.033	1.000	
	-0.084	0.087	-0.003	0.089	0.132*	-0.140	0.117*	-0.071	-0.094	0.027	0.277*	-0.296*	0.218*	-0.246*	0.194*	-0.093		
wage ↓	0.085*	-0.017	0.042	0.011	-0.047	0.085*	-0.088*	0.081*	0.030	-0.013	-0.190*	0.233*	-0.086*	0.191*	-0.067	0.061	-0.443*	1.000
	0.183*	-0.042	0.090	0.028	-0.108	0.231*	-0.200*	0.195*	0.067	-0.029	-0.450*	0.506*	-0.203*	0.405*	-0.145	0.195	-1.000*	
Number of observations:	800																	

Spearman's correlation and Tetrachoric correlations for binary variables (below).
↑ *describes improvements,* ↓ *refers to subjective worsenings.*
* *p<0.05.*

Table A4.4

OLS estimation results for wage change by groups

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$		$\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$		$w_{ijt} - w_{i,j-1,t-1}$		$w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$	
	owner	renter	owner	renter	owner	renter	owner	renter
Strain improved	0.0136	-0.0304	0.0137	-0.0306	0.0776	-0.659*	0.0934	-0.747*
	(0.0375)	(0.0320)	(0.0369)	(0.0315)	(0.679)	(0.389)	(0.696)	(0.422)
Strain worsened	0.0623	-0.0396	0.0608	-0.0406	0.943	-0.782	0.901	-0.876
	(0.0444)	(0.0482)	(0.0437)	(0.0471)	(1.186)	(0.531)	(1.238)	(0.566)
Work time improved	0.0275	-0.0486	0.0277	-0.0466	0.581	-0.139	0.632	-0.131
	(0.0345)	(0.0333)	(0.0340)	(0.0327)	(0.680)	(0.400)	(0.697)	(0.431)
Work time worsened	0.0151	0.0313	0.0155	0.0315	1.157	0.643	1.266	0.743
	(0.0528)	(0.0583)	(0.0522)	(0.0569)	(1.306)	(0.588)	(1.363)	(0.644)
Security against job loss	0.00786	0.0201	0.00747	0.0197	0.371	-0.299	0.363	-0.311
improved	(0.0357)	(0.0326)	(0.0353)	(0.0320)	(0.746)	(0.355)	(0.758)	(0.385)
Security against job loss	0.0254	0.0821	0.0250	0.0820	0.871	0.613	0.979	0.632
worsened	(0.0698)	(0.0665)	(0.0684)	(0.0655)	(1.001)	(0.804)	(1.032)	(0.854)
Use of skills improved	0.0573*	0.0348	0.0565*	0.0337	1.631**	0.307	1.751**	0.344
	(0.0306)	(0.0362)	(0.0301)	(0.0356)	(0.739)	(0.429)	(0.761)	(0.463)
Use of skills worsened	0.0994	0.00820	0.0994	0.00762	2.123**	0.0900	2.246**	0.0954
	(0.0724)	(0.0374)	(0.0717)	(0.0367)	(0.976)	(0.412)	(0.984)	(0.442)
Commuting improved	-0.0959**	-0.0319	-0.0933**	-0.0312	-1.399	-0.505	-1.412	-0.507
	(0.0416)	(0.0358)	(0.0411)	(0.0351)	(1.078)	(0.431)	(1.119)	(0.459)
Commuting worsened	-0.0248	0.0198	-0.0235	0.0192	-0.180	0.0687	-0.193	0.105
	(0.0418)	(0.0487)	(0.0413)	(0.0478)	(0.827)	(0.521)	(0.848)	(0.579)
Chances for promotion	0.0127	0.00687	0.0116	0.00728	0.138	0.333	0.103	0.291
improved	(0.0378)	(0.0347)	(0.0372)	(0.0340)	(0.820)	(0.395)	(0.843)	(0.427)
Chances for promotion	-0.121	-0.0317	-0.121	-0.0301	-2.207*	-0.324	-2.347**	-0.388
worsened	(0.0758)	(0.0700)	(0.0747)	(0.0690)	(1.149)	(0.756)	(1.172)	(0.821)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes	yes	yes	yes	yes
Number of observations	245	555	245	555	245	555	245	555
R^2	0.120	0.046	0.120	0.046	0.088	0.063	0.088	0.058

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$		$\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$		$w_{ijt} - w_{i,j-1,t-1}$		$w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$	
	$\Delta u \leq 0$	$\Delta u > 0$	$\Delta u \leq 0$	$\Delta u > 0$	$\Delta u \leq 0$	$\Delta u > 0$	$\Delta u \leq 0$	$\Delta u > 0$
Strain improved	0.00291	-0.0322	0.00219	-0.0318	-0.151	-0.798	-0.191	-0.894
	(0.0317)	(0.0404)	(0.0312)	(0.0398)	(0.422)	(0.523)	(0.445)	(0.576)
Strain worsened	0.0159	-0.0407	0.0149	-0.0418	0.233	-0.995	0.195	-1.121
	(0.0357)	(0.0640)	(0.0351)	(0.0623)	(0.739)	(0.636)	(0.772)	(0.691)
Work time improved	-0.0101	-0.0215	-0.00904	-0.0193	0.628	-0.114	0.663	-0.0799
	(0.0331)	(0.0399)	(0.0325)	(0.0391)	(0.467)	(0.504)	(0.489)	(0.551)
Work time worsened	-0.0103	0.125	-0.00883	0.124	0.583	1.564	0.656	1.754
	(0.0435)	(0.0981)	(0.0427)	(0.0956)	(0.605)	(1.022)	(0.645)	(1.116)
Security against job loss	0.00539	0.0228	0.00523	0.0219	-0.0672	-0.343	-0.0820	-0.363
improved	(0.0301)	(0.0474)	(0.0296)	(0.0464)	(0.389)	(0.547)	(0.406)	(0.597)
Security against job loss	0.0327	0.162	0.0314	0.162	0.290	1.614	0.303	1.736
worsened	(0.0501)	(0.150)	(0.0491)	(0.148)	(0.774)	(1.315)	(0.813)	(1.429)
Use of skills improved	-0.00102	0.116**	-0.00119	0.114**	0.237	1.558***	0.244	1.706***
	(0.0297)	(0.0494)	(0.0292)	(0.0484)	(0.489)	(0.564)	(0.508)	(0.614)
Use of skills worsened	0.00718	0.0748	0.00706	0.0734	0.432	1.073	0.444	1.154
	(0.0414)	(0.0691)	(0.0408)	(0.0683)	(0.551)	(0.688)	(0.573)	(0.744)
Commuting improved	-0.0547	-0.0379	-0.0525	-0.0379	-0.747	-0.780	-0.749	-0.796
	(0.0335)	(0.0470)	(0.0329)	(0.0461)	(0.605)	(0.549)	(0.630)	(0.596)
Commuting worsened	0.0182	0.00322	0.0188	0.00286	0.612	-0.766	0.661	-0.779
	(0.0491)	(0.0480)	(0.0482)	(0.0471)	(0.572)	(0.646)	(0.623)	(0.694)
Chances for promotion	0.00530	0.0133	0.00450	0.0140	-0.0813	0.696	-0.131	0.668
improved	(0.0312)	(0.0439)	(0.0307)	(0.0430)	(0.548)	(0.537)	(0.569)	(0.579)
Chances for promotion	-0.123**	-0.0285	-0.122**	-0.0277	-1.914**	-0.0401	-2.099**	-0.0281
worsened	(0.0492)	(0.0988)	(0.0484)	(0.0976)	(0.836)	(0.916)	(0.878)	(0.990)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes	yes	yes	yes	yes
Number of observations	486	314	486	314	486	314	486	314
R^2	0.083	0.093	0.083	0.094	0.083	0.115	0.080	0.112

Continued on next page

Table A4.4 (continued)

OLS estimation results for wage change by groups

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$		$\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$		$w_{ijt} - w_{i,j-1,t-1}$		$w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$	
	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$
Strain improved	-0.0273	0.00688	-0.0275	0.00592	-0.946**	0.0327	-1.030**	-0.00458
	(0.0389)	(0.0350)	(0.0382)	(0.0345)	(0.415)	(0.529)	(0.452)	(0.555)
Strain worsened	-0.0767	0.0257	-0.0772*	0.0243	-1.806**	0.636	-1.929**	0.580
	(0.0469)	(0.0505)	(0.0462)	(0.0492)	(0.765)	(0.732)	(0.824)	(0.761)
Work time improved	-0.0532	-0.0125	-0.0523	-0.00999	-0.123	0.178	-0.120	0.224
	(0.0410)	(0.0334)	(0.0403)	(0.0327)	(0.400)	(0.550)	(0.438)	(0.571)
Work time worsened	0.0180	0.0435	0.0181	0.0442	0.803	0.735	0.870	0.894
	(0.0632)	(0.0711)	(0.0622)	(0.0691)	(0.654)	(0.873)	(0.718)	(0.933)
Security against job loss	0.0149	0.0276	0.0146	0.0269	-0.170	0.298	-0.211	0.333
improved	(0.0352)	(0.0370)	(0.0346)	(0.0362)	(0.362)	(0.564)	(0.395)	(0.588)
Security against job loss	0.198**	-0.0207	0.197**	-0.0205	2.523***	-0.290	2.696**	-0.317
worsened	(0.0991)	(0.0576)	(0.0976)	(0.0567)	(0.962)	(0.868)	(1.047)	(0.907)
Use of skills improved	0.0592	0.0208	0.0580	0.0202	0.750	0.740	0.793	0.807
	(0.0412)	(0.0347)	(0.0406)	(0.0339)	(0.479)	(0.596)	(0.527)	(0.618)
Use of skills worsened	0.0262	0.0492	0.0258	0.0490	0.483	1.057	0.515	1.130*
	(0.0590)	(0.0418)	(0.0583)	(0.0411)	(0.601)	(0.653)	(0.656)	(0.670)
Commuting improved	-0.0469	-0.0496	-0.0453	-0.0488	-0.428	-1.092	-0.410	-1.126
	(0.0415)	(0.0382)	(0.0408)	(0.0374)	(0.418)	(0.760)	(0.447)	(0.793)
Commuting worsened	0.000621	0.00384	0.000872	0.00360	-0.0650	-0.110	-0.0198	-0.125
	(0.0629)	(0.0344)	(0.0618)	(0.0337)	(0.623)	(0.565)	(0.697)	(0.586)
Chances for promotion	-0.00854	0.0174	-0.00820	0.0167	0.140	0.207	0.112	0.143
improved	(0.0409)	(0.0344)	(0.0403)	(0.0337)	(0.437)	(0.655)	(0.477)	(0.680)
Chances for promotion	-0.00375	-0.0872*	-0.00445	-0.0854*	-0.602	-1.051	-0.678	-1.150
worsened	(0.106)	(0.0480)	(0.105)	(0.0474)	(0.977)	(0.765)	(1.072)	(0.793)
Additional control variables (as in Table 4.6)	yes	yes	yes	yes	yes	yes	yes	yes
Number of observations	370	430	370	430	370	430	370	430
R^2	0.077	0.066	0.076	0.066	0.098	0.081	0.092	0.080

Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*
'Δu' refers to growth in unemployment rate.

Table A4.5

OLS estimation results for different time horizons

1998–2004	(1)	(2)	(3)	(4)
Variables	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$	$\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	$w_{ijt} - w_{i,j-1,t-1}$	$w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Strain improved	-0.00984	-0.0100	-0.199	-0.219
	(0.0263)	(0.0260)	(0.382)	(0.405)
Strain worsened	-0.00491	-0.00451	-0.212	-0.240
	(0.0340)	(0.0335)	(0.685)	(0.727)
Work time improved	0.00965	0.0103	0.139	0.173
	(0.0241)	(0.0237)	(0.398)	(0.419)
Work time worsened	0.0539	0.0532	0.977	1.048
	(0.0395)	(0.0390)	(0.653)	(0.690)
Security against job loss improved	-0.00241	-0.00217	-0.0781	-0.0669
	(0.0231)	(0.0229)	(0.358)	(0.377)
Security against job loss worsened	0.0604	0.0584	0.550	0.578
	(0.0483)	(0.0476)	(0.710)	(0.748)
Use of skills improved	0.0130	0.0126	0.405	0.425
	(0.0259)	(0.0256)	(0.409)	(0.431)
Use of skills worsened	0.0182	0.0183	0.375	0.407
	(0.0342)	(0.0339)	(0.398)	(0.422)
Commuting improved	-0.0704**	-0.0685**	-1.021*	-1.064*
	(0.0279)	(0.0275)	(0.557)	(0.584)
Commuting worsened	-0.0332	-0.0316	-0.538	-0.556
	(0.0293)	(0.0290)	(0.409)	(0.433)
Chances for promotion improved	0.0302	0.0289	0.483	0.487
	(0.0269)	(0.0265)	(0.449)	(0.472)
Chances for promotion worsened	-0.0859*	-0.0848*	-0.897	-0.997
	(0.0492)	(0.0487)	(0.745)	(0.793)
Number of observations			543	
R^2	0.091	0.090	0.086	0.084
1999–2003	(1)	(2)	(3)	(4)
Variables	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$	$\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$	$w_{ijt} - w_{i,j-1,t-1}$	$w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$
Strain improved	0.0103	0.00956	-0.133	-0.156
	(0.0299)	(0.0295)	(0.403)	(0.432)
Strain worsened	-0.00845	-0.00788	-0.537	-0.579
	(0.0391)	(0.0386)	(0.770)	(0.821)
Work time improved	0.00666	0.00756	-0.0320	-0.00731
	(0.0264)	(0.0261)	(0.357)	(0.379)
Work time worsened	0.0887*	0.0876*	1.461**	1.564**
	(0.0453)	(0.0446)	(0.730)	(0.774)
Security against job loss improved	0.00700	0.00721	-0.0396	-0.0198
	(0.0266)	(0.0263)	(0.355)	(0.377)
Security against job loss worsened	0.0857*	0.0832*	1.161*	1.216*
	(0.0510)	(0.0503)	(0.662)	(0.701)
Use of skills improved	-0.00750	-0.00792	-0.0371	-0.0519
	(0.0292)	(0.0289)	(0.389)	(0.414)
Use of skills worsened	-0.00412	-0.00360	0.00277	0.0136
	(0.0393)	(0.0389)	(0.439)	(0.466)
Commuting improved	-0.0663**	-0.0642**	-0.631	-0.659
	(0.0308)	(0.0304)	(0.443)	(0.471)
Commuting worsened	-0.0357	-0.0341	-0.505	-0.534
	(0.0338)	(0.0334)	(0.461)	(0.488)
Chances for promotion improved	0.0414	0.0399	0.941**	0.965**
	(0.0296)	(0.0292)	(0.383)	(0.407)
Chances for promotion worsened	-0.0808	-0.0797	-0.696	-0.790
	(0.0567)	(0.0560)	(0.830)	(0.885)
Number of observations	447	447	447	447
R^2	0.097	0.097	0.117	0.112

Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Table A4.6

Tobit regression results for wage cut, given that individuals change to lower wages (full specification)

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$	(2) $\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$	(3) $w_{ijt} - w_{i,j-1,t-1}$	(4) $w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$
Strain improved	-0.0128**	-0.0136**	-0.318**	-0.355**
	(0.00636)	(0.00563)	(0.153)	(0.142)
Strain worsened	0.00261	0.00426	0.0223	0.0624
	(0.00734)	(0.00676)	(0.206)	(0.203)
Work time improved	0.00154	0.00467	0.146	0.231
	(0.00621)	(0.00555)	(0.164)	(0.161)
Work time worsened	0.00109	0.00108	0.0715	0.0787
	(0.00783)	(0.00671)	(0.195)	(0.176)
Security against job loss improved	-0.00964	-0.00886	-0.180	-0.170
	(0.00631)	(0.00569)	(0.155)	(0.149)
Security against job loss worsened	-0.000167	0.00808	-0.0213	0.207
	(0.00922)	(0.00802)	(0.224)	(0.207)
Use of skills improved	0.00594	0.00618	0.210	0.227
	(0.00590)	(0.00540)	(0.150)	(0.144)
Use of skills worsened	0.0132*	0.0123**	0.333**	0.334**
	(0.00688)	(0.00613)	(0.166)	(0.151)
Commuting improved	-0.0117*	-0.00967*	-0.306	-0.264
	(0.00634)	(0.00568)	(0.196)	(0.182)
Commuting worsened	0.00126	9.43e-05	0.0112	-0.0210
	(0.00615)	(0.00552)	(0.156)	(0.147)
Chances for promotion improved	0.0105*	0.00908*	0.148	0.117
	(0.00573)	(0.00526)	(0.150)	(0.144)
Chances for promotion worsened	-0.0291**	-0.0297***	-0.696**	-0.738***
	(0.0119)	(0.0113)	(0.289)	(0.276)
Fringe benefits improved	0.00880	0.00849	0.178	0.186
	(0.00608)	(0.00559)	(0.155)	(0.152)
Fringe benefits worsened	-0.000998	-0.00540	0.0100	-0.116
	(0.00851)	(0.00762)	(0.202)	(0.190)
Job in general improved	0.00412	0.00215	0.184	0.133
	(0.00560)	(0.00516)	(0.150)	(0.141)
Job in general worsened	-0.00778	-0.00790	-0.0816	-0.0718
	(0.0134)	(0.0127)	(0.293)	(0.281)
Control variables (as in Table 4.11)	yes	yes	yes	yes
Pseudo R^2	0.135	0.138	0.035	0.026
Number of observations			800	
Uncensored observations	195	254	195	254
Censored observations	605	546	605	546

Marginal effects after tobit regression.
Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*

Table A4.7
Tobit regression results for wage cut, given that individuals change to lower wages by groups

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$ owner	(2) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$ renter	(3) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$ owner	(4) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$ renter	(5) $w_{ijt} - w_{i,j-1,t-1}$ owner	(6) $w_{ijt} - w_{i,j-1,t-1}$ renter	(7) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$ owner	(8) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$ renter
Strain improved	0.00968	-0.0191**	0.00950	-0.0196***	0.279	-0.375**	0.283	-0.415***
	(0.0108)	(0.00759)	(0.0100)	(0.00676)	(0.340)	(0.159)	(0.333)	(0.157)
Strain worsened	0.00520	0.00120	0.0105	0.000293	0.148	-0.00401	0.314	-0.0191
	(0.0133)	(0.00883)	(0.0127)	(0.00789)	(0.488)	(0.176)	(0.496)	(0.169)
Work time improved	-0.00239	0.00502	-0.000745	0.00841	0.147	0.145	0.205	0.228
	(0.0102)	(0.00730)	(0.00975)	(0.00645)	(0.322)	(0.149)	(0.322)	(0.148)
Work time worsened	0.00286	-0.00235	-0.00292	-2.98e-05	0.327	-0.0247	0.140	0.0236
	(0.0156)	(0.00887)	(0.0152)	(0.00749)	(0.473)	(0.178)	(0.465)	(0.164)
Security against job loss improved	0.000284	-0.01000	-0.00229	-0.00783	0.139	-0.189	0.0350	-0.157
	(0.0102)	(0.00686)	(0.00947)	(0.00620)	(0.338)	(0.132)	(0.312)	(0.129)
Security against job loss worsened	-0.00690	0.000745	-0.00123	0.00819	-0.210	-0.0185	0.00864	0.128
	(0.0130)	(0.0117)	(0.0122)	(0.0101)	(0.405)	(0.250)	(0.387)	(0.237)
Use of skills improved	0.0164*	0.00552	0.0152*	0.00524	0.674*	0.146	0.644*	0.152
	(0.00936)	(0.00705)	(0.00917)	(0.00632)	(0.347)	(0.139)	(0.344)	(0.135)
Use of skills worsened	0.0133	0.0108	0.0106	0.0110	0.436	0.227	0.373	0.255*
	(0.0130)	(0.00812)	(0.0113)	(0.00728)	(0.385)	(0.152)	(0.332)	(0.147)
Commuting improved	-0.00387	-0.0144**	-0.00968	-0.0100	-0.123	-0.293**	-0.343	-0.224*
	(0.0124)	(0.00727)	(0.0120)	(0.00627)	(0.457)	(0.145)	(0.488)	(0.133)
Commuting worsened	-0.000736	0.00349	-0.00350	0.00230	0.0236	0.0242	-0.0639	-0.00114
	(0.0109)	(0.00756)	(0.0107)	(0.00655)	(0.341)	(0.148)	(0.343)	(0.141)
Chances for promotion improved	0.00186	0.0154**	0.00589	0.0114*	-0.152	0.265**	0.0117	0.197*
	(0.0103)	(0.00645)	(0.00976)	(0.00585)	(0.372)	(0.123)	(0.345)	(0.118)
Chances for promotion worsened	-0.0453**	-0.0211	-0.0421**	-0.0236*	-1.197*	-0.410	-1.057*	-0.493*
	(0.0228)	(0.0133)	(0.0208)	(0.0133)	(0.636)	(0.258)	(0.555)	(0.281)
Additional control variables (as in Table 4.11)	yes	yes	yes	yes	yes	yes	yes	yes
Pseudo R^2	0.153	0.153	0.175	0.145	0.037	0.045	0.031	0.031
Number of observations	245	555	245	555	245	555	245	555
Uncensored observations	72	123	87	167	72	123	87	167
Censored observations	173	432	158	388	173	432	158	388

Variables	(1) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$ $\Delta u \leq 0$	(2) $\frac{w_{ijt}}{w_{i,j-1,t-1}}$ $\Delta u > 0$	(3) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$ $\Delta u \leq 0$	(4) $\frac{w^{real}_{ijt}}{w^{real}_{i,j-1,t-1}}$ $\Delta u > 0$	(5) $w_{ijt} - w_{i,j-1,t-1}$ $\Delta u \leq 0$	(6) $w_{ijt} - w_{i,j-1,t-1}$ $\Delta u > 0$	(7) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$ $\Delta u \leq 0$	(8) $w^{real}_{ijt} - w^{real}_{i,j-1,t-1}$ $\Delta u > 0$
Strain improved	-0.0102	-0.0143	-0.00936	-0.0158	-0.245	-0.338	-0.224	-0.409
	(0.00782)	(0.0106)	(0.00695)	(0.00967)	(0.196)	(0.242)	(0.180)	(0.251)
Strain worsened	0.00910	-0.00673	0.0127	-0.00815	0.196	-0.186	0.304	-0.239
	(0.0101)	(0.0103)	(0.00913)	(0.00989)	(0.320)	(0.216)	(0.308)	(0.230)
Work time improved	0.00741	0.00288	0.00901	0.00734	0.292	0.130	0.350*	0.247
	(0.00746)	(0.00957)	(0.00668)	(0.00894)	(0.216)	(0.209)	(0.208)	(0.224)
Work time worsened	0.00763	-0.00342	0.00412	-0.00162	0.251	-0.0116	0.156	0.0410
	(0.00921)	(0.0125)	(0.00779)	(0.0112)	(0.258)	(0.253)	(0.222)	(0.248)
Security against job loss improved	-0.00878	-0.00679	-0.00699	-0.00840	-0.141	-0.152	-0.0899	-0.201
	(0.00725)	(0.00940)	(0.00650)	(0.00897)	(0.184)	(0.196)	(0.170)	(0.211)
Security against job loss worsened	-0.00971	0.0143	0.000798	0.0156	-0.190	0.237	0.115	0.281
	(0.0107)	(0.0132)	(0.00955)	(0.0118)	(0.279)	(0.254)	(0.268)	(0.246)
Use of skills improved	0.00631	0.0119	0.00562	0.0117	0.305	0.271	0.285	0.299
	(0.00713)	(0.00845)	(0.00644)	(0.00801)	(0.216)	(0.178)	(0.200)	(0.186)
Use of skills worsened	0.00900	0.0181	0.00964	0.0163	0.313	0.363	0.345*	0.367
	(0.00880)	(0.0111)	(0.00755)	(0.0105)	(0.230)	(0.231)	(0.202)	(0.238)
Commuting improved	-0.0167**	-0.00420	-0.0155**	-4.60e-05	-0.452	-0.0858	-0.436	-0.00646
	(0.00830)	(0.00895)	(0.00755)	(0.00798)	(0.289)	(0.173)	(0.273)	(0.170)
Commuting worsened	0.00106	0.00413	-0.00105	0.00551	0.0558	-0.0162	-0.00122	0.00844
	(0.00787)	(0.00892)	(0.00730)	(0.00805)	(0.212)	(0.188)	(0.201)	(0.187)
Chances for promotion improved	0.00907	0.0116	0.00771	0.0101	0.0984	0.199	0.0687	0.174
	(0.00758)	(0.00792)	(0.00693)	(0.00745)	(0.223)	(0.160)	(0.210)	(0.164)
Chances for promotion worsened	-0.0385**	-0.0302*	-0.0385**	-0.0322*	-1.020**	-0.515	-1.037**	-0.590*
	(0.0163)	(0.0171)	(0.0155)	(0.0169)	(0.447)	(0.315)	(0.420)	(0.342)
Additional control variables (as in Table 4.11)	yes	yes	yes	yes	yes	yes	yes	yes
Pseudo R^2	0.191	0.130	0.185	0.154	0.048	0.034	0.035	0.030
Number of observations	486	314	486	314	486	314	486	314
Uncensored observations	116	79	153	101	116	79	153	101
Censored observations	370	235	333	213	370	235	333	213

Continued on next page

Table A4.7 (continued)

Tobit regression results for wage cut, given that individuals change to lower wages by groups

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\frac{w_{ijt}}{w_{i,j-1,t-1}}$		$\frac{w_{ijt}^{real}}{w_{i,j-1,t-1}^{real}}$		$w_{ijt} - w_{i,j-1,t-1}$		$w_{ijt}^{real} - w_{i,j-1,t-1}^{real}$	
	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$	$age \leq 34$	$age > 34$
Strain improved	-0.0157*	-0.00641	-0.0140*	-0.00886	-0.409*	-0.147	-0.399*	-0.214
	(0.00851)	(0.00899)	(0.00743)	(0.00816)	(0.221)	(0.213)	(0.204)	(0.198)
Strain worsened	-0.0108	0.00847	-0.0127	0.0135	-0.375	0.245	-0.435	0.386
	(0.0122)	(0.00918)	(0.0106)	(0.00877)	(0.358)	(0.267)	(0.349)	(0.281)
Work time improved	0.00115	0.00429	0.00231	0.00689	0.110	0.214	0.144	0.290
	(0.00793)	(0.00860)	(0.00711)	(0.00800)	(0.182)	(0.227)	(0.176)	(0.224)
Work time worsened	0.00111	-0.00135	0.00144	-0.000574	0.0859	0.0125	0.0924	0.0398
	(0.0111)	(0.0106)	(0.00873)	(0.0101)	(0.254)	(0.271)	(0.215)	(0.270)
Security against job loss improved	-0.00673	0.00135	-0.00882	0.00236	-0.150	0.149	-0.212	0.187
	(0.00737)	(0.00817)	(0.00667)	(0.00759)	(0.167)	(0.218)	(0.164)	(0.216)
Security against job loss worsened	0.0145	-0.0111	0.0267**	-0.00571	0.382	-0.252	0.695**	-0.0963
	(0.0130)	(0.0121)	(0.0103)	(0.0113)	(0.315)	(0.309)	(0.295)	(0.301)
Use of skills improved	0.0130	0.00334	0.0143**	0.00260	0.316	0.254	0.371*	0.242
	(0.00864)	(0.00725)	(0.00729)	(0.00709)	(0.206)	(0.212)	(0.195)	(0.210)
Use of skills worsened	0.00811	0.0168*	0.0158*	0.0117	0.197	0.497*	0.424**	0.386
	(0.0103)	(0.00944)	(0.00888)	(0.00861)	(0.222)	(0.262)	(0.215)	(0.239)
Commuting improved	-0.00943	-0.0142	-0.00525	-0.0137*	-0.175	-0.435	-0.0849	-0.434
	(0.00849)	(0.00879)	(0.00745)	(0.00819)	(0.195)	(0.312)	(0.190)	(0.302)
Commuting worsened	-0.00302	0.00560	-0.00655	0.00583	-0.0856	0.143	-0.185	0.157
	(0.00897)	(0.00832)	(0.00771)	(0.00783)	(0.208)	(0.218)	(0.195)	(0.213)
Chances for promotion improved	0.0100	0.0109	0.0110	0.00761	0.235	0.0788	0.267	-0.0173
	(0.00783)	(0.00758)	(0.00685)	(0.00742)	(0.187)	(0.213)	(0.170)	(0.229)
Chances for promotion worsened	-0.0232	-0.0218*	-0.0236	-0.0273**	-0.476	-0.578	-0.504	-0.755**
	(0.0180)	(0.0129)	(0.0182)	(0.0122)	(0.374)	(0.353)	(0.397)	(0.351)
Additional control variables (as in Table 4.11)	yes	yes	yes	yes	yes	yes	yes	yes
Pseudo R^2	0.245	0.145	0.232	0.178	0.074	0.035	0.049	0.034
Number of observations	370	430	370	430	370	430	370	430
Uncensored observations	74	121	108	146	74	121	108	146
Censored observations	296	309	262	284	296	309	262	284

Marginal effects after tobit regression.
Robust standard errors clustered for individuals in parentheses.
**** p<0.01, ** p<0.05, * p<0.1.*
'Δu': Growth in unemployment rate.

5 References

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