Cohort Wage Effects and Job Mobility

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September 1, 2009

Abstract

Economic conditions at the time of labour market entry can induce wage differentials between entry cohorts. While there exists much empirical evidence on the existence and persistence of these cohort effects, little is known about their interaction with employees' mobility behaviour. Using German administrative data, this paper analyzes the determinants of job mobility, emphasizing the role of cohort wage differentials. The analysis suggests that cohort effects play an important role in explaining job transitions. Labour market entrants affected by unfavourable conditions and earning less than the average starting wage are more likely to change jobs, directly as well as indirectly. Moreover, our empirical analysis shows that labour market transitions tend to reduce the cohort effects in earnings, implying that job mobility operates as an adjustment mechanism that reduces the initial wage differences between entry cohorts.

JEL codes: *E24, J31, J62, J64*
Keywords: *mobility, job-to-job, wages, cohort effects.*

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†We are grateful to Thomas Zwick, as well as to participants at the 12th IZA Summer School, the ESPE 2009 Annual Congress, the Scottish Economic Society 2009 Annual Meeting, the COST workshop (Wages and Firms), the DFG-SPP March 2009 workshop, the Leibniz Seminar of BENA, and seminars at RWI for helpful comments. We also thank the staff of the Forschungsdatenzentrum of the IAB for their help with the data. Part of this research was carried out while Peggy David was visiting Humboldt-Universität Berlin, which she thanks for their hospitality. Financial support from the German Science Foundation (DFG) through SPP1169 "Flexibility in Heterogeneous Labor Markets" and from the Leibniz Association is gratefully acknowledged. All correspondence to Ronald Bachmann and Peggy David, RWI, Hohenzollernstr. 1-3, 45128 Essen, Germany, Email: bachmann@rwi-essen.de and david@rwi-essen.de.

1 Introduction

A number of empirical studies suggests that economic conditions prevailing at the time workers enter the labour market significantly affect their earnings (e.g. Bloom and Freeman, 1986, and Shin, 1994). Whether these wage effects are long-run in nature has been a widely studied question, yielding ambiguous results (e.g. Baker, Gibbs, and Holmstrom, 1994, Oreopoulos, Heisz, and von Wachter, 2006, Harris and Holmstrom, 1982, Welch, 1979). The standard competitive model, for example, implies that the labour market operates as a spot market, where wages are solely determined by labour demand and labour supply and thus are equal to the individual's marginal productivity. In such a model, labour market shocks at the beginning of a worker's career are temporary and do not lead to long-lasting wage effects. However, alternative economic theories suggest that differences in initial labour market conditions - arising, for example, from variations in the cohort size or fluctuations in the business cycle - can induce persistent wage differentials between entry cohorts (e.g. Harris and Holmstrom, 1982).

While theoretical and empirical studies confirm the existence as well as the persistence of entry cohort effects in wages, research on how these cohort wage differentials are related to workers' job mobility remains scarce. Looking at a sample of Canadian college graduates, Oreopoulos, Heisz, and von Wachter (2006) provide one of the few studies analyzing the impact of job-starting conditions on worker's early career. They document that the unemployment rate at job entry, diminishing the worker's starting wage, significantly raises the probability of job separation. Furthermore, they provide descriptive evidence that this increased job mobility in turn positively affects wages, and therefore is able to partly reverse the earnings losses experienced through less favourable career starting conditions.

In a related vein, we study the relationship between cohort effects and early job mobility addressing two questions: Do cohort-induced wage differentials significantly affect individual's mobility behaviour? And can job mobility act in such a way as to reduce these initial wage gaps? In order to answer these two questions, we use a large administrative data set representing 2% of German employees during the time period 1975-2004. We proceed in three steps, which always contain a descriptive and an econometric analysis. First, we quantify the impact of cohort effects on the wages of labour market entrants. We then examine the determinants of individual job mobility, emphasizing the role of cohort wage differences in this context. In order to do so, the probability of certain labour market transitions is modelled as a function of worker and establishment characteristics, as well as entry cohort wage effects. Finally, we analyse to what extent worker mobility contributes to a reduction of the initial wage gaps between different entry cohorts.

We contribute to the existing literature in several ways: First, we use a large, and representative, sample of labour market entrants in West Germany. Second, we are able to distinguish various destination states as well as voluntary and involuntary job mobility. Third, using an instrumental variable approach, we are able to identify a causal effect of job mobility on wage growth.

Our results suggest that cohort-induced wage differentials play an important role in explaining job transitions. Entry cohorts affected by unfavourable conditions and earning less than the average cohort starting wage show an increased mobility compared to cohorts with average or higher-than-average earnings. Moreover, our empirical analysis shows that the wage discrepancies that occur across cohorts at the time of labour market entry decrease with experience. Direct and indirect labour market transitions further reduce the cohort effects in earnings, implying that job mobility indeed operates as an adjustment process that reduces the initial wage differences between entry cohorts.

The remainder of this paper is organized as follows. The next section contains a review of the literature on cohort effects and early job mobility. Section 3 presents a description of the data set, particularly addressing the identification of job transitions. In Section 4 the methodology used in this paper is discussed. Descriptive statistics and estimation results are presented in Section 5. Section 6 delivers a number of sensitivity analyses. Section 7 concludes.

2 Cohort Effects and Job Mobility in the Literature

The analysis conducted in this paper builds on two strands of the literature: (i) the cohort effects literature, studying the impact of initial labour market shocks on earnings, and (ii) the job mobility literature, analyzing the determinants and wage effects of individual job transitions. In this section, we provide a brief survey of the existing theoretical and empirical studies for both strands. Although the subsequent empirical analysis does not address the causes of cohort induced wage differentials, for the sake of completeness our overview also covers studies providing various explanations for differences in wages between entry cohorts.

2.1 Cohort Wage Effects

The economic literature points to several theories that explain why initial labour market conditions might lead to wage differentials between entry cohorts. On the supply side, one factor consists in variations in the size of entry cohorts. Studies examining the impact of the demographic cycle on earnings find that an important increase in labour supply - emanating, for example, from the entry of baby boomers into the job market - adversely affects entry wages (Freeman, 1979, and Welch, 1979). The analysis whether these wage disadvantages experienced by large cohorts at a young age remain throughout workers' careers has created contention among researchers (Berger, 1989, Bloom, Freeman, and Korenman, 1987, and Murphy, Plant, and Welch, 1988). In particular, Bloom, Freeman, and Korenman (1987) track the progress of different U.S. cohorts from 1969 to 1984. Their results suggest that large cohorts are able to at least partly catch up in earnings within a decade after labour market entry. Welch (1979) finds similar results for the period from 1967 to 1975 and confirms that wage disadvantages do not persist as the cohort ages. However, Berger (1989) using almost identical data but less restrictive identification assumptions does not find any convergence in wages across cohorts.

Wage differentials between entry cohorts may also be the result of labour demand shocks, such as technological progress and business cycle fluctuations. For example, there is evidence that individuals hired during economic recessions experience lower entry wages than individuals hired in economic upturns (e.g. Bils, 1985, and Solon, Barsky, and Parker, 1994), with several studies finding that this cohort effect is persistent (e.g. Oreopoulos, Heisz, and von Wachter, 2006, Oyer, 2006, and von Wachter and Bender, 2007). Several theories on wage determination can be put forward to explain this long-term impact of poor initial economic conditions. Models of implicit contracts, developed for example by Azariadis (1975) as well as Harris and Holmstrom (1982) and empirically tested by Beaudry and DiNardo (1991), and Baker, Gibbs, and Holmstrom (1994), suggest that business cycle conditions at the time of labour market entry may affect individuals' long-term wages, because of missing or insufficient wage adjustments. Another type of model focuses on cyclical variations in hiring and promotion standards, which might lead to differences in workers' productivity and hence to differences in current and future earnings (Okun, 1973, and Reder, 1955). A prevalent explanation for persistent cohort effects is based on the neoclassical human capital model, stating that the initial economic situation affects workers' opportunity to accumulate skills and thus has a sustained impact on individual labour market performance (Gibbons and Waldman, 2004).

2.2 Early Job Mobility

Workers' careers - and in particular young workers' careers - can be characterized by a collective search process: Workers search for firms that value their skills most highly, while firms search for the most productive workers (cf. Jovanovic, 1979). Given their characteristics, labour market entrants may not be able to immediately find an employer that offers them the most productive jobs, which implies that job transitions are an integral part of early working lives (Topel and Ward, 1992). Thus early job mobility plays an important role in improving the quality of job matches and thus for the evolution of workers' wages. This fact especially holds true in times of unfavourable economic conditions, when suitable jobs are particularly hard to find. However, job transitions as a mechanism to adjust workers' early wages to average market wages are not taken into account by the theories of cohort wage effects mentioned above.

Empirical studies examining the determinants of job transitions early in the career suggest that the wage level is crucial for individual mobility. Topel and Ward (1992), for example, analyze the mobility patterns of young men and find a lower job stability for lower-paid jobs. This corresponds to results reported by Oreopoulos, Heisz, and von Wachter (2006), who show that economic downturns, diminishing workers starting wage, significantly raise the rate of job change. Common explanations for these findings are based on job search (Burdett, 1979) and job matching approaches (Jovanovic, 1979), which predict a long-lasting catch-up process if wages have temporarily declined. Thus, workers in employment relationships where they do not experience sustained productivity increases tend to search for better jobs that offer higher wages as well as higher match qualities. This implies that employer-to-employer transitions that occur for voluntary reasons are able to increase young workers' wages. There exists empirical evidence that confirms the beneficial wage effects of voluntary job mobility which takes place during the early stages of peoples' working lives. Antel (1986), and Bartel and Borjas (1978), for example, find mobility-induced wage premiums that range between 8% and 20%. Similarly, the analysis by Oreopoulos, Heisz, and von Wachter (2006) indicates that wage disadvantages, experienced by workers graduating in a recession, are partly reversed through job changes. This implies that individuals affected by poor initial labour market conditions might use the opportunity to advance in their careers through job changes, avoiding persistent earnings disadvantages and yielding a convergence between cohort and average market wages.

Likewise, firms may eventually lay off workers who experience relatively high wages because of favourable starting conditions. This kind of separation might lead to a loss of initial wage advantages and therefore contribute to a reduction of cohort effects. A prevalent explanation for wage losses of displaced workers is based on the human capital theory (Becker, 1975). It suggests that investments in job specific skills create a higher earnings potential, making job mobility less profitable. In a related vein, the segmentation hypothesis as well as the model of seniority wages (Lazear, 1981) predict that workers who change employers loose their firm-specific human capital and thus experience a wage cut. In line with the model of imperfect information, Gibbons and Katz (1991) argue that at the time of hiring, employers are insufficiently informed about workers' productivity. Since firms have an incentive to lay off less able workers, displacements may serve as a negative signal to other employers. This adverse selection of job movers implies that involuntary employer-to-employer transitions may entail negative wage effects. Consistent with that, empirical studies examining the costs of job displacements, suggest that lay-offs adversely affect workers' earnings. Kletzer and Fairlie (2003) and von Wachter and Bender (2006) point to the fact that job displacements in workers' early careers lead to sizeable and persistent wage losses. Similarly, von Wachter and Bender (2007) show that initial wage advantages, obtained from favourable labour market conditions, are reduced when workers lose their job.

With the exception of, for example, Perez and Sanz (2005) and Perticara (2004), only a few empirical papers analyze voluntary and involuntary job changes simultaneously and thus allow for both beneficial as well as adverse mobility. The one most closely resembling our analysis is the study by Perticara (2004). Using US panel data, she finds that workers earning less than the customary wage rate are more likely to initiate a job change, which leads to a post-separation wage gain. On the contrary, workers earning more than the average wage have a higher probability of being laid off and often experience wage losses after separation. However, like many other studies that analyze the returns to job mobility, Perticara (2004) fails to explicitly control for the potential endogeneity of job transitions, which may lead to biased and inconsistent OLS estimates.

3 Data

In the following analysis we employ a data set provided by the Institute for Employment Research (IAB), the IAB Employment Sample (IABS). The basis of this data set is the *Employment Statistics Register*, an administrative panel data set of the employment history of all individuals in Germany who worked in an employment covered by social security between 1975 and 2004. For 1995, this data source contains the employee history of nearly 79.4% of all employed persons in Western Germany, and 86.2% of all employed persons in Eastern Germany.[1] The IABS is a 2% representative sample of the Employment Statistics Register for the time period 1975-2004, supplemented with information on all unemployment spells of the workers covered.

A key information included in the data set are gross daily wages subject to social security contributions, which we deflate using consumer prices. Further worker characteristics included are the employees' year of birth, sex, education, and nationality.[2] To meet the problem of inconsistent and missing information on the individual's education, we use an education variable corrected following an imputation procedure provided by Fitzenberger, Osikumino, and Völter (2006).[3]

We restrict our sample to West-German individuals who started their career between 1980 and 1999, such that we are able to follow their career paths for the first five years on the labour market. Because the record on unemployment benefit recipients are unreliably measured before 1980, we restrict our analysis to persons that enter the labour market in 1980 or later. For a better comparison of wages, we exclude part-time workers, homeworkers, trainees, and individuals with parallel employment spells. We leave unconsidered the

[1] The employee history is based on the integrated notification procedure for health insurance, the statutory pension scheme, and unemployment insurance. At the beginning and at the end of any employment spell, employers have to notify the social security agencies. This information is exact to the day. For spells spanning more than one calendar year, an annual report for each employee registered within the social insurance system is compulsory, and provides an update on, for example, the qualification and the current occupation of the employee.

[2] A detailed description of the Employment Statistics Register and the notification procedure is given by Bender, Haas, and Klose (2000). Note that civil servants and self-employed workers are not included in the data.

[3] Particularly, we use the imputation procedure 2B, where education reports are extrapolated if a person's education sequence is consistent. Only for individuals having inconsistencies in their education reports, indicating reporting errors, extrapolation is restricted to degrees that are reported at least three times.

starting wages close to the contribution ceiling.[4] Finally, we drop individuals with missing values for the variables used in the empirical analysis. Using these selection criteria our final sample comprises 195,384 labour market entrants with a total of about 1.3 million spells.

The IABS data are representative regarding employment covered by the social security system but not regarding unemployment because only those unemployed who are entitled to transfer payments are covered. In the data, we can derive three labour market states at each moment in time: employment (E) covered by social security, unemployment (U), if the worker is receiving transfer payments, and non-participation (N). Since the latter state cannot be directly observed, we define non-participants as individuals leaving the sample. Transitions to non-participation, however, can also mean transitions to the civil service, to self-employment, retirement or marginal employment, because these destinations are not covered by social security legislation and are therefore not covered by the data set.

Regarding these labour market states, measurement errors might exist. Because of the way the data are collected, both firms' reports of a new employee and individuals' notifications of moving into or out of unemployment are not exactly consistent with the actual change of labour market state. This potential measurement error can be corrected in the following way: If the time lag between two employment or unemployment notifications does not exceed 30 days, it is defined as a direct transition between the two states recorded, and as an intervening spell of non-participation if the time interval between two records is larger than 30 days.

Since the IABS data set contains daily information on the employment and unemployment history of every individual in the sample, it is possible to calculate separation flows taking into account every change of the labour market state that occurs within a certain time period. Using the three mentioned states E, U and N, as well as the establishment identification number provided in the data set, we are able to identify three different separation flows: transitions (i) from employment to nonparticipation (EN), (ii) from employment to unemployment (EU) and (iii) from employment to another employment (EE). It should be noted here that our definition of a job is based on the establishment level and not on the

[4] Other studies based on administrative individual data are usually subject to the problem that the wage information in the IABS is censored at the social contribution ceiling. Because we only consider individuals entering the labour market for the first time, these data problems barely affect our analysis: Less than 0.4% of the workers' starting wages are top coded. Within the first five experience years about 10% of the workers reach wages affected by the contribution ceiling.

firm level. A transition from one establishment to another one within the same firm will therefore be identified as an employer-to-employer flow.

Throughout the following empirical analysis, we focus on EE flows. In this context, recent research has pointed out that a distinction between voluntary and involuntary job changes proves to be important (Perticara, 2004). Since the IABS data do not designate any reason for a job separation, we are not able to directly differentiate between voluntary and involuntary moves. As an alternative, we follow previous studies (e.g. Perez and Sanz, 2005) and compare direct employer-to-employer transitions and those with an intervening unemployment spell of less than 1 month (EE_d) to employer-to-employer transitions with an intervening unemployment spell that is larger than 1 month (EE_{id}).[5] Corresponding to the notion in the job mobility literature, the first type of separation is with a high probability initiated by the worker and can usually be seen as a voluntary move. The latter one, however, results in all likelihood from a lay-off and can be considered as an involuntary move.[6] Transition rates are calculated by using aggregate employment as the denominator.

4 Econometric Framework

In the first part of the empirical analysis, we estimate the probability of experiencing a certain job separation by using a hazard rate model. Since the IABS data set contains daily information, a continuous-time framework is used. The hazard rate is assumed to take the following proportional hazard from:

$$\lambda(t, X(t)) = \lambda_0(t) exp(X(t)\beta), \tag{1}$$

where λ_0 is the baseline hazard function, which measures the effect of the elapsed employment duration on the separation rate of a certain reference group. X refers to a set of

[5] Our data set only records unemployment spells if the worker receives unemployment benefits. We are thus not able to identify the true length of unemployment. Following Fitzenberger and Wilke (2006), we therefore use the nonemployment period as a proxy for the true unemployment period, which is defined as all nonemployment spells after an employment spell including at least one period with receipt of transfer benefits.

[6] Using this definition, job separations induced by the employer might be considered as voluntary moves. This is possible, for example, if the employer notifies the worker in advance that he will be laid off, giving him the opportunity to search on-the-job. We therefore view our measure of voluntary transitions as an upper bound.

time-varying covariates, and β is a vector of coefficients to be estimated. To take unobserved heterogeneity into account, we follow Lancaster (1990) and include a term α, which is assumed to have a multiplicative effect on the individual hazard and thus leads to the mixed proportional hazard model

$$\lambda(t, X(t), \alpha) = \alpha\lambda_0(t)exp(X(t)\beta). \tag{2}$$

We parameterize the hazard function as a piecewise-constant exponential model, i.e. we assume a baseline hazard rate which is constant within given time intervals, but is allowed to vary between them. The basic duration is partitioned into k prespecified sub-segments with cutpoints $0 = t_0 < t_1 < \ldots < t_k$. The baseline hazard then becomes

$$\lambda_0(t) = \begin{cases} \lambda_1, & t \in (0, \tau_1], \\ \lambda_2, & t \in (\tau_1, \tau_2], \\ \ldots & \\ \lambda_k, & t \in (\tau_{k-1}, \infty], \end{cases}$$

where the k parameters $\lambda_1, \ldots, \lambda_k$ represent the separation probability for a certain reference group in one particular time interval. Thus, in contrast to the Cox proportional hazard model, explicit estimates of the baseline hazard function can be obtained, enabling us to directly assess the effect of duration dependence. In the subsequent analysis we distinguish between seven sub-segments: 0-6 months, 7-12 months, 13-18 months, 19-24 months, 2-3 years, and more than 3 years of employment duration.

Using the piecewise constant exponential model, we estimate two different competing risk models. In order to get a general idea of young workers' mobility behaviour, we first distinguish between three possible separation destination states: individuals may transit from one employer to another (EE), from employment to unemployment (EU), and from employment to non-participation (EN). In a second step, we focus on employer-to-employer transitions and estimate the competing hazards of changing employers directly (EE^D flows: EE flows with an intervening nonemployment spell $<$ 1 month) and changing employers indirectly (EE^{ID} flows: EE flows with an intervening nonemployment spell $\geq$ 1 month). In the case of continuous time models with multiple destinations, the log-likelihood can be divided into the sum of multiple sub-contributions. Given this separability property, it is possible to estimate a competing-risk model by estimating a single-risk model for each destination. On the basis of the IABS data set, we are now able to explain the probability of certain transitions by a set of individual and establishment characteristics. As covariates we use gender, skill level, and employment duration at the individual level, as well as establishment size and industry at the establishment level. In order to account for differences in economic conditions at the time of separation, we also include yearly dummies.

The explanatory variable of main interest is the cohort effect in wages at the beginning of the worker's career. In order to calculate these initial wage differentials we estimate, in a first step, the following wage regression by using a simple OLS approach:

$$\ln w_{i0} = \alpha_1 + \alpha_2 X_{i0} + \alpha_3 Z_{e0} + \sum_{j=2}^{J} \beta_j C_j + \epsilon_{i0}, \tag{3}$$

where $\ln w_{i0}$ refers to the real hourly log wage of individual i at the time of entering the labour market (t=0), X_{i0} is a vector of individual characteristics, Z_{e0} is a vector of establishment characteristics, C_j denotes a set of $j-1$ cohort dummies indicating the year of entry, and the vectors α_1, α_2, α_3, and β_j are coefficients to be estimated. In a second step, the coefficients of the cohort dummy variables β_j, obtained from equation (3) by using an arbitrarily chosen reference cohort, are transformed into deviations from the grand mean of starting wages following **?**). These starting wage deviations enter the hazard rate equation with one variable comprising values smaller than zero and a second one comprising values larger than zero. This allows positive and negative deviations to have different effects on the transition probabilities.[7] Table A.1 provides definitions as well as summary statistics of all the worker and establishment characteristics used in the empirical analysis.

In the second part of the empirical analysis, we aim at investigating the contribution of individuals' mobility behaviour to the adjustment of cohort and market wages. That is, we want to examine whether job mobility can significantly decrease the initial wage differential between job starting cohorts within the first five years of their labour market career. To do so, we concentrate on individuals who stayed in their first job, individuals who directly transit from one employer to another one, and individuals who indirectly change employers. We then compare how the initial cohort effect changes for these three groups. This is done by estimating the following model:

$$\ln w_{i5} = \gamma_1 + \gamma_2 X_{it} + \gamma_3 Z_{et} + \sum_{j=2}^{J} \delta_j C_j + \sum_{j=2}^{J} \theta_{1j} C_j EE_{it}^{D} + \sum_{j=2}^{J} \theta_{2j} C_j EE_{it}^{ID} + \epsilon_{it}. \tag{4}$$

[7]Since a predicted variable is included as a regressor, standard errors are corrected following Murphy and Topel (1985).

In contrast to model (3), we now examine the workers' wages five years after their labour market entry. Moreover, equation (4) extends the previous one as it additionally includes two dummy variables EE_{it}^{D} and EE_{it}^{ID}, indicating whether only direct or only indirect employer changes take place.[8] In order to gauge the wage effect of mobility for different entry cohorts, we interact these two indicator variables with the cohort dummies C_j.

Previous studies point out that the failure to control for the simultaneous determination of wage and mobility may result in biased and inconsistent estimators (Abowd, Kramarz, and Roux, 2006, Altonji and Shakotko, 1987, and von Wachter and Bender, 2006). This could also arise in our analysis of the impact of job mobility on the variation of earnings. We address the possible endogeneity of changing employers by using an instrumental variable approach. We instrument for the probability of voluntary mobility with the regional share of workers who are older than 40 years. We argue that workers at this stage of their labour market career *ceteris paribus* change jobs less often than younger workers. This implies that in regions with a high share of older workers less job openings are available which negatively affects the likelihood of a voluntary job change. Following Goeggel and Zwick (2009), who analyze the job and wage mobility behaviour of German apprentices, we use a mass layoff indicator as an instrument for the probability of involuntary job mobility. Mass layoffs are defined as a 30 percent reduction of the employment in an establishment. It is assumed that workers are more likely to leave the job involuntarily if an establishment's labour force was reduced significantly in the year of separation.

The regional age-share as well as the mass-layoff indicator are highly correlated with the workers' likelihood to change employers directly and indirectly, respectively, making them strong instruments. Moreover, it seems plausible to argue that both instruments are uncorrelated to unobservable individual characteristics affecting wages.[9]

[8]By this definition, workers are allowed to switch employers several times within the first five years. Restricting the sample to workers who changed jobs only once leads to very similar, but slightly reduced estimation results. This implies that the first move is crucial for the workers' wage development.

[9]It has to be mentioned that the instrumental variable regression yields a local average treatment effect on the treated, i.e. the estimated wage effect of job changing only applies for those individuals affected by the instrument (Imbens and Angrist, 1994).

5 Descriptive Evidence and Estimation Results

Before we turn to the impact of initial cohort effects on workers' mobility behaviour and the role of job mobility in reversing wage differentials, we first show the pattern of cohort wage effects. Figure B.1 plots the development of average log real daily wages for workers starting their career between 1980 and 1999. It also shows the average wages at the time of labour market entry as well as five years later. It is apparent that the average starting wage varies significantly over time, suggesting wage differentials across entry cohorts. These initial cohort effects in wages are relatively persistent, decreasing slowly over time. As mentioned in Section 2, cohort wage effects at the time of labour market entry might be the result of labour demand shocks. The relation between this type of shock and average entry wages is shown in Figure B.2, which compares detrended average starting wages with variations in the business cycle. It is evident that wages at the time of labour market entry follow the GDP growth rate. Table A.2 shows that the correlation between entry wages and GDP growth is increasing up to the 2-year lag and is falling afterwards.

So far, our analysis has been purely descriptive and at an aggregate level. However, the observed variations in starting wages may not solely be driven by differences in labour market entry conditions, but also by variations in cohort composition. In Table A.3 we report summary statistics of cohort characteristics. One can see - beside the downward trend in cohort size - that the entry cohorts under consideration only slightly differ in observable characteristics (share of females, share of skill groups and cohort size). This issue is examined explicitly in Table A.4, which presents the initial cohort effects obtained by estimating several specifications of wage equation (3). With the exception of workers entering the labour market in 1990, all cohorts earn starting wages that significantly differ from the average. For example, workers starting their working career in 1980 earn about 22% less than the average, while entrants in 1999 have starting wages 20% above the average.[10] Taking into account observable person (skill level, gender) and establishment characteristics (industry, region, establishment size) reduces the estimated cohort wage effects slightly by about 1 to 3 percentage points. Therefore, composition effects only explain a small part of the initial wage differential across cohorts.

[10] The probability to enter the labour market follows a strongly procyclical pattern. Therefore we argue that the estimated wage losses constitute a lower bound of costs due to unfavourable starting conditions, as the costs of an increased unemployment probability would add.

5.1 The Impact of Cohort Wage Effects on Job Mobility

To illustrate the job mobility behaviour of individuals affected by diverse starting conditions, we begin by presenting different separation flows that occur within the first five years after labour market entry. Table A.5 displays several separation transitions by labour market experience and deviation from the mean starting wage. It becomes apparent that all transition rates are decreasing with the individual's labour market experience. Furthermore, we see that, in general, workers of the lower quintiles tend to be more mobile at the beginning of their career. More precisely, one year after labour market entry, workers with starting wages below the average show employer-to-employer transition rates ranging from 19.2% to 19.4%. By contrast, the first year EE transition rates of workers whose entry wage lies above the sample mean only reach about 15%. The transitions from employment to non-participation (EN) show a very similar pattern, while for employment-to-unemployment transitions (EU) slightly different properties can be observed. Workers with starting wages near the average and those with positive deviations seem to have the lowest transition rates, varying from 11.0% to 11.3% one year after labour market entry. On the other hand, workers with negative deviations from the mean starting wage show increased inflows to unemployment (about 15.5%).

Since we are mainly interested in employer-to-employer transitions, Table A.5 additionally displays the transition rates for direct (EE^{D}) and indirect employment changes (EE^{ID}). Here, we can again observe that direct EE flows are higher for workers from the lower quintiles. Moreover, EE flows with an intervening nonemployment spell increase with negative wage deviations and are least likely to occur for workers with starting wages near or above the sample mean. From these two descriptive tables we can infer that - compared to entrants earning average wages - workers with positive wage deviations show equivalent EE^{ID} but lower EE^{D} transition rates. Individuals with negative initial wage deviations, however, tend to be more mobile.

In order to analyse the determinants of different separation transitions, we estimate the hazard rate models described in the previous section. First, we estimate a basic specification including positive and negative cohort effects as main explanatory variables. To examine how the impact of initial wage differentials varies with employment duration, the interaction of cohort effects with workers' employment duration is featured in the second specification. For both specifications, Table A.6 displays the estimated hazard rates of the three separation flows EE, EN, and EU, obtained from a piecewise constant exponential hazard model that takes into account person-specific unobserved heterogeneity. The results are generally in line with the literature on job mobility. For example, the hazard rates of all three transitions indicate that job separations are negatively correlated with the skill level, suggesting that transitions are a more common phenomenon among less educated individuals. Moreover, the transition probabilities decline with establishment size, which implies that larger establishments offer more stable employment relationships. The estimation results also show that employment duration negatively affects the hazard of separating. This negative duration dependence can be attributed to the accumulation of firm-specific human capital, which makes it less profitable to dissolve the worker-firm match. The result that women face a significantly lower risk of job separation than men, irrespective of the destination state, is however not in line with other studies on labour market flows, which usually find women to be more mobile than men, and to be more likely to transit from employment to unemployment or nonparticipation. This result is mainly due to maternity and child care. For our sample, consisting of job starters these factors do not play an important role, which explains the discrepancy between our results and those usually found in the literature.

As for the impact of the cohort wage effect on transition probabilities, the estimation results largely confirm the results from the descriptive analysis. The probability that an EE flow occurs is increasing with negative cohort effects, implying that workers with a cohort wage below the average are more likely to move from one employment to another one. In contrast to this, EE transitions are decreasing with positive cohort effects. While for transitions from employment to nonparticipation one can see very similar results, a different pattern occurs for the outflows to unemployment. Here the estimation results indicate that positive cohort wage effects do not have a statistically significant impact on the transition probability. Negative cohort effects, however, significantly increase the probability of moving. Our overall finding is that workers entering the labour market during poor economic conditions tend to be more mobile, which is in line with the evidence presented by Oreopoulos, Heisz, and von Wachter (2006).

The coefficients obtained from estimating the hazards of changing jobs directly and indirectly are shown in Table A.7. For both types of transitions, we find very similar features with respect to the workers' gender and skill level, the size of the establishments, as well as job duration. The same is true for the cohort wage effects. Positive cohort effects significantly reduce the likelihood of direct employer-to-employer transitions: A one percentage point increase of the positive cohort effect lowers the transition probability by 0.24%. Arguably, labour market entrants affected by advantageous economic conditions earn wages above the average and do not have an incentive to search for better paid jobs.

An inspection of the estimation results for the interaction terms in the second specification reveals that at the time of labour market entry, EE transitions are not affected by positive wage deviations. However, the effect of positive wage deviations rises with increasing employment duration. The probability of changing employers through a nonemployment spell, however, is not significantly affected by positive cohort effects.

Negative cohort effects, on the other hand, are positively correlated with employer-to-employer transitions, irrespective of whether they occur directly or indirectly. If the negative wage differential is raised by one percentage point, the hazards increase by 0.29% or 0.37%, respectively. This implies that workers with cohort wages 20% below the average, as for example the cohort entering in 1980, have a 5.8% (7.4%) higher probability to switch employers directly (indirectly). The estimation hazard rates of the interaction terms in the second specification indicate that this effect is even larger at the beginning of a worker's career and then gradually declines with employment duration. An increased probability of direct employer-to-employer transitions could be the result of on-the-job search. Workers entering the labour market during unfavourable starting conditions, and earning less than the average cohort wage, might feel underpaid. While they are employed, these workers search for better jobs, and are likely to switch jobs without an intervening nonemployment spell. With respect to the increased likelihood of indirect employer changes, one could argue that negative cohort effects lower the individual's motivation and thus increase shirking. Since these workers face a higher risk to be laid off, employer-to-employer transitions are likely to occur through a period of nonemployment.

5.2 Adjustment of Cohort Wage Effects

We now want to examine the role of job mobility in the reduction of cohort wage effects over time. In order to do so, we compare wages and wage growth between stayers and movers five years after labour market entry. Stayers are defined as workers who stay in their first job. Movers are classified into three groups: workers who change jobs within the first five years of their labour market career through (i) direct, (ii) indirect, or through (iii) direct as well as indirect moves.[11] Table A.8 reports how workers are distributed over these different categories.

A first impression of the wage development is given in Table A.9, which presents the wage growth for the five quintiles of the entry wage distribution by deviation from the mean starting wage. One can see that in general, wage growth is much higher for workers whose entry wages show negative deviations, irrespective of whether they stay at their first employer, move directly or indirectly. For example, the wages of stayers in the first quintile grow by more than 160%, while the wages of those in the upper quintiles only grow by 11%. Furthermore, wage growth is steeper when workers directly change employers (206.3% and 16%, respectively). For workers in the lower quintiles, even indirect job changes lead to higher wage growth than staying with the same employer (188.6%).[12] This implies that unfavourable labour market conditions at the time of labour market entry result in inappropriate job matches, such that any kind of job change seems to be beneficial to workers. In contrast to this, wages grow less when workers of the upper quintiles switch employers. Direct employer-to-employer transitions result in a 16% wage growth, while indirect transitions even lead to a wage cut (-7.5%). These patterns of wage growth indicate that cohorts with wage disadvantages at the beginning of their career are able to at least partly catch up in earnings and close the initial wage gaps. An inspection of the corresponding wages in Table A.10 reveals that at the time of entering the labour market, log real daily wages range from 3.09 in the first quintile to 4.28 in the fifth quintile, implying that there exist clear differences in wages when workers start their career. Looking at stayers, we see that after gaining five years of labour market experience, these wage differentials have decreased. Since workers of the lower quintiles experience a higher wage growth than those of the upper quintiles, a convergence between wages occurs. It is apparent that this wage convergence is even more pronounced when workers are mobile. Given the results from the descriptive analysis, we can argue that job mobility leads to a stronger reduction of initial

[11] By this definition, workers are allowed to switch employers several times within the first five years. Restricting the sample to workers who only once changed jobs leads to very similar, but slightly reduced effects.

[12] The lower wage growth of indirect moves outweighs the higher wage growth of direct moves. Due to this, workers who change jobs directly and indirectly within their first five career years show an intermediate wage growth rate.

wage differentials than staying with the same employer.

To empirically test this statement, we estimate Equation (4) as described in the previous section.[13] The estimation results obtained from a simple OLS model are shown in Table A.11. The coefficients in the first row of this table refer to the effect staying, direct job mobility, as well as indirect job mobility have on workers' wages. The interaction terms indicate how these main effects are modified when we distinguish between the different entry cohorts. Overall, one can see that wages of stayers are 5.3% above the average. This positive effect is even higher for workers who started their career after 1988, but lower for those who entered the labour market before. The latter workers, suffering from initial wage disadvantages, can benefit from changing employers without an intervening nonemployment spell. While the main effect of direct job mobility lies at about 8%, it is even higher for these older cohorts. The older cohorts also benefit strongly from changing employers indirectly. However, this does not completey outweigh the average wage decrease of 10.3% for all involuntary job movers. One can observe opposite results for cohorts who start their career later and initially earn wages above the average: Compared to the main effect, direct job changes cause a lower wage, while indirect job mobility results in a more pronounced reduction. There are some exceptions to this (e.g. for cohorts 1991 and 1997), but mostly with insignificant effects. Figure B.3 illustrates the estimation results. It displays the estimated cohort effects five years after labour market entry for stayers and movers compared to the initial cohort effects at the time of labour market entry. It is apparent that five years later, the wage differentials across cohorts have decreased for both, movers and stayers. This reduction is much stronger when workers change their employers, implying that job mobility leads to a stronger convergence between cohort wages. For example, the 22.7% wage disadvantage of cohorts who started their career in 1982 and stay in their first job is reduced to 16.4%, while direct and indirect movers experience a reduction of negative wage differentials to 5.3% and 7.7%, respectively.

As discussed in Section 4, due to the endogeneity of the mobility decision, these estimation results may not reflect the causal impact of mobility on workers' wages. Therefore, in Table A.12 we additionally present the estimation results from an IV approach, where the regional share of workers older than 40 years and a mass-layoff indicator are used as

[13] Here we only consider workers who change jobs only directly or indirectly within the first five years. Those who show both types of job mobility (about 18.000 workers) are not included.

instruments for voluntary and involuntary job mobility. In the first step of the estimation, we regress the mobility variables on the instruments and a set of covariates. It reveals that both, the age share as well as the mass-layoff indicator strongly determine the workers' probability to change jobs. As the F-test of joint significance of the instruments shows, the problem of weak instruments does not apply here.[14]

We use the predicted values resulting from the first stage regression to conduct the IV estimation. The results in Table A.12 show that the OLS and IV coefficients show a very similar pattern, although the sign of the estimated IV coefficients changes earlier for stayers as well as for direct movers and later for indirect movers.[15] The major difference between OLS and IV results regards cohorts who entered the labour market between 1997 and 1999. For both, direct and indirect job mobility, OLS yields wage effects that negatively deviate from the main effect. However, using the regional age-ratio as an instrument for direct mobility, results in positive coefficients, implying that the wages of these cohorts grow particlarly strong. We put this down to the fact that younger cohorts initially experience wage advantages and therefore require a much stronger wage increase in order to consider voluntary job changes. This stands in contrast to older cohorts who are disadvantaged by poor starting conditions and therefore willing to switch jobs for smaller wage raises. Moreover, for younger cohorts changing jobs indirectly the IV estimators are more negative than the analogous OLS coefficients. Workers tend to perform poorly after an indirect job move, which is in line with the fact that separations due to mass layoffs are involuntary. Anticipating the massive labour force reduction, better workers engage in on-the-job search and thus are able to find appropriate new jobs. The remaining worse workers are not able to find adequate job offers, such that their reservation threshold and therefore their future wages are reduced. The estimated wage cut is particularly strong for the more recent cohorts, who benefitted from good economic starting conditions at the time of labour market entry. They will lose their initial wage advantages when they have to leave their job involuntarily. O

[14]The test statistic strongly exceeds the critical value of 10 recommended by Stock and Watson (2003).

[15]As described before, the IV approach yields a local average treatment effect on the treated. Therefore, the IV coefficients only apply for those affected by the instrument and in fact cannot be compared to the OLS coefficients.

6 Sensitivity Analysis

In order to test the robustness of our results, we first address the endogenous nature of the labour market entry decision.[16] It might be the case that in times of unfavourable economic conditions, individuals decide not to enter the labour market and postpone their career start by lengthening their educational training. We therefore separately analyze workers who start working after finishing their apprenticeship. We argue that these workers are not able to respond to fluctuations in economic conditions and are thus unlikely to defer the starting point of their labour market career. The regression results for these workers (not displayed) show that workers who start their career after an apprenticeship and who are affected by positive wage deviations experience almost the same transition probabilities as the entire sample. With respect to negative deviations, however, apprentices are more likely to separate from their employer. Being less able to postpone their labour market entry and to avoid poor match qualities, unfavourable economic conditions have a stronger effect on their transition probabilities.

In a second robustness test, we carry out both parts of the empirical analysis for different sub-populations separately, since estimation results may vary across different groups. Distinguishing between gender and skill groups leads to very similar estimation results as for the whole sample. Finally, we conduct a robustness check with respect to the definition of the instrumental variables. The regional share of workers older than 40 years, which serves as an instrument for voluntary job mobility, is redefined as the regional share of workers older than 45 years. This variation leads to similar results, although positive cohort effects seem to be somewhat larger.

7 Conclusion

In this paper, we aim at investigating the relationship between cohort effects in wages and workers' mobility behaviour early in their career. Throughout the analysis, we use a large administrative panel data set which contains detailed information on workers in the German labour market and covers the time period 1975-2004. As a first step, we model the probability of experiencing job transitions, where we focus on direct and indirect

[16]The results from the robustness tests are not displayed in this paper, but are available from the authors upon request.

employer-to-employer transitions. One of the explanatory variables, which is included in the regressions and is of particular interest, is the cohort wage effect caused by variations in economic conditions at the time of labour market entry. These wage differentials varying significantly across cohorts are found to be an important determinant of job mobility. For all types of transitions we can show that workers affected by poor economic starting conditions are more likely to separate from their job. For example, workers who earn wages 20% below the mean starting wage face a 5.8% higher risk to directly switch employers than workers who start their career ten years later. This reflects the fact that young workers try to enhance their career by searching for jobs that offer them higher wages and better job match qualities.

In the second step of the analysis, we estimate the change in the cohort wage effect that can be attributed to job mobility. To tackle the endogeneity problem which emerges from the fact that mobility is likely to be correlated with unobserved individual and job characteristics affecting earnings, we apply an instrumental-variable approach. As an instrument for the probability of direct job mobility, we use the regional share of workers older than 40 years. The probability of indirect job mobility, in return, is instrumented by a mass layoff indicator. This proceeding allows us to identify and exactly quantify a causal effect of job mobility on wage growth. We find that wage differentials across cohorts decrease with labour market experience. Moreover, the estimation results show that cohorts with wage advantages can benefit from direct job changes, but are adversely affected by employer transitions with an intervening unemployment spell. For workers with initial wage disadvantages, however, job mobility in general increases wages. From this we can conclude that cohort effects in wages are further reduced by job mobility, which implies that job mobility operates as an adjustment process that reverses the initial wage discrepancies.

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Appendix A Tables

Table A.1: Definition of characteristics

Variable	Mean	Std. Dev.	Definition
EE flows	0.1098	0.3127	Transitions from one employer to another one.
EU flows	0.0645	0.2456	Transitions from employment to unemployment.
EN flows	0.1206	0.3257	Transitions from employment to nonparticipation.
Separations	0.2950	0.4560	EE + EU + EN.
EE^{D} flows	0.1164	0.3207	Direct EE flows and EE flows with an intervening nonemployment spell $<$ 1 month.
EE^{ID} flows	0.0465	0.2106	EE flows with an intervening nonemployment spell $\geq$ 1 month.
Age	23.146	3.2361	Age of individual.
Low-skilled	0.1417	0.3488	Dummy=1 if individual holds a lower secondary school diploma without a professional degree.
Medium-skilled	0.7881	0.4086	Dummy=1 if individual has a lower secondary school diploma and professional degree; or a high school diploma and without a professional degree; or a school diploma as well as a professional degree.
High-skilled	0.0690	0.2534	Dummy=1 if individual holds a university degree or university of applied sciences degree.
Industry dummies	0.0221	0.1468	Agriculture, Mining and Energy
	0.3087	0.4619	Production
	0.0917	0.2885	Construction
	0.2517	0.4340	Trade, Transport
	0.2759	0.4469	Services
	0.0368	0.1882	State.
Establishment size dummies	0.3257	0.4686	1-19 employees
	0.2314	0.4218	20-99 employees
	0.2871	0.4524	100-999 employees
	0.1157	0.3626	more than 1000 employees
Entry Wage	48.236	20.992	Real daily wage at the time of labour market entry.
Wage	56.946	27.798	Real daily wage.
Positive cohort effect	13.473	7.7953	Positive deviation from grand mean starting wage in %.
Negative cohort effect	18.437	4.4881	Negative deviation from grand mean starting wage in %.

Source: Authors' calculations, based on IABS 1975-2004.

Table A.2: Correlation between cohort entry wages and GDP growth

	GDP_t	GDP_{t-1}	GDP_{t-2}	GDP_{t-3}	GDP_{t-4}
cohort entry wage	0.0169	0.2208	0.4456	0.3956	0.3427

Source: Authors' calculations, based on IABS 1975-2004.

Table A.3: Cohort characteristics at labour market entry

Year of	Characteristics					
entry	Age	Female	Low-skill	Med.-skill	High-skill	Cohort size
1980	19.81 (2.69)	0.46 (0.49)	0.31 (0.46)	0.64 (0.48)	0.05 (0.22)	14050 (0)
1981	19.79 (2.42)	0.44 (0.49)	0.28 (0.45)	0.67 (0.47)	0.05 (0.21)	12974 (0)
1982	19.97 (2.47)	0.45 (0.49)	0.24 (0.43)	0.71 (0.45)	0.05 (0.21)	11406 (0)
1983	20.11 (2.53)	0.46 (0.49)	0.23 (0.42)	0.71 (0.45)	0.06 (0.22)	10891 (0)
1984	20.18 (2.47)	0.46 (0.49)	0.24 (0.43)	0.71 (0.45)	0.05 (0.22)	11083 (0)
1985	20.42 (2.57)	0.46 (0.49)	0.23 (0.42)	0.71 (0.45)	0.06 (0.23)	11129 (0)
1986	20.58 (2.54)	0.46 (0.49)	0.21 (0.40)	0.74 (0.44)	0.05 (0.23)	12185 (0)
1987	20.77 (2.59)	0.47 (0.49)	0.20 (0.39)	0.74 (0.44)	0.06 (0.23)	12238 (0)
1988	20.99 (2.64)	0.46 (0.49)	0.20 (0.40)	0.74 (0.44)	0.06 (0.24)	11942 (0)
1989	21.03 (2.66)	0.46 (0.49)	0.20 (0.39)	0.73 (0.44)	0.07 (0.25)	12728 (0)
1990	21.14 (2.68)	0.45 (0.49)	0.19 (0.39)	0.74 (0.44)	0.07 (0.25)	12858 (0)
1991	21.36 (2.81)	0.47 (0.49)	0.19 (0.39)	0.73 (0.44)	0.08 (0.25)	11589 (0)
1992	21.69 (2.86)	0.47 (0.49)	0.17 (0.37)	0.76 (0.44)	0.07 (0.28)	11965 (0)
1993	21.58 (2.83)	0.47 (0.49)	0.17 (0.36)	0.76 (0.42)	0.07 (0.26)	9363 (0)
1994	21.62 (2.91)	0.46 (0.49)	0.16 (0.37)	0.74 (0.43)	0.08 (0.28)	8561 (0)
1995	21.69 (2.97)	0.44 (0.49)	0.18 (0.38)	0.73 (0.45)	0.09 (0.29)	8336 (0)
1996	21.68 (2.95)	0.46 (0.49)	0.17 (0.36)	0.74 (0.44)	0.09 (0.29)	7404 (0)
1997	21.75 (2.94)	0.45 (0.49)	0.18 (0.38)	0.73 (0.44)	0.09 (0.28)	7770 (0)
1998	21.82 (3.05)	0.46 (0.49)	0.19 (0.39)	0.71 (0.46)	0.10 (0.30)	8004 (0)
1999	21.61 (2.86)	0.45 (0.49)	0.21 (0.40)	0.71 (0.45)	0.08 (0.27)	7833 (0)

Source: Authors' calculations, based on IABS 1975-2004.

Table A.4: Cohort wage effects at labour market entry

Year of entry	(1) Coeff.	(S. D.)	(2) Coeff.	(S. D.)	(3) Coeff.	(S. D.)
1980	-0.2281***	(0.0045)	-0.1986***	(0.0041)	-0.1941***	(0.0039)
1981	-0.2377***	(0.0046)	-0.2128***	(0.0042)	-0.2097***	(0.0040)
1982	-0.2529***	(0.0048)	-0.2343***	(0.0044)	-0.2277***	(0.0041)
1983	-0.2433***	(0.0049)	-0.2283***	(0.0045)	-0.2228***	(0.0042)
1984	-0.2339***	(0.0048)	-0.2157***	(0.0044)	-0.2163***	(0.0041)
1985	-0.2064***	(0.0049)	-0.1924***	(0.0045)	-0.1948***	(0.0042)
1986	-0.1465***	(0.0047)	-0.1366***	(0.0043)	-0.1376***	(0.0040)
1987	-0.1102***	(0.0047)	-0.1011***	(0.0043)	-0.1012***	(0.0040)
1988	-0.0798***	(0.0049)	-0.0772***	(0.0045)	-0.0801***	(0.0042)
1989	-0.0492***	(0.0048)	-0.0509***	(0.0044)	-0.0541***	(0.0041)
1990	0.0037	(0.0048)	0.0013	(0.0044)	-0.0017	(0.0041)
1991	0.0726***	(0.0051)	0.0656***	(0.0047)	0.0629***	(0.0044)
1992	0.2166***	(0.0051)	0.2090***	(0.0047)	0.2074***	(0.0044)
1993	0.2212***	(0.0056)	0.2113***	(0.0051)	0.2173***	(0.0048)
1994	0.2219***	(0.0058)	0.2051***	(0.0053)	0.2068***	(0.0050)
1995	0.2421***	(0.0059)	0.2173***	(0.0054)	0.2102***	(0.0051)
1996	0.2240***	(0.0062)	0.2006***	(0.0057)	0.2033***	(0.0053)
1997	0.1923***	(0.0062)	0.1724***	(0.0056)	0.1742***	(0.0053)
1998	0.1935***	(0.0061)	0.1715***	(0.0056)	0.1691***	(0.0053)
1999	0.2001***	(0.0061)	0.1937***	(0.0056)	0.1890***	(0.0052)
Demographics			X		X	
Firm controls					X	

Source: Authors' calculations, based on IABS 1975-2004.

Note: Dependent variable is the log real daily wage. Cohort effects are calculated as deviations from the grand mean starting wage. Demographic characteristics include gender and skill level. Firm controls include dummy variables for establishment size, industry, and region. The three specifications differ by the inclusion of demographic and/or firm controls only.

Table A.5: Mobility statistics by deviation from mean entry wage

Deviation from entry wage	Experience year	Worker flow rates				
		EE	EN	EU	EE^D	EE^{ID}
1st quintile	1st year	0.194	0.217	0.155	0.213	0.111
	3rd year	0.180	0.182	0.149	0.196	0.103
	5th year	0.177	0.152	0.126	0.182	0.081
2nd quintile	1st year	0.192	0.190	0.121	0.196	0.089
	3rd year	0.182	0.171	0.111	0.186	0.083
	5th year	0.154	0.144	0.098	0.162	0.068
3rd quintile	1st year	0.184	0.189	0.112	0.185	0.078
	3rd year	0.177	0.161	0.103	0.176	0.074
	5th year	0.152	0.142	0.097	0.157	0.062
4th quintile	1st year	0.170	0.167	0.110	0.178	0.080
	3rd year	0.160	0.144	0.100	0.166	0.072
	5th year	0.143	0.132	0.095	0.150	0.060
5th quintile	1st year	0.150	0.145	0.113	0.158	0.081
	3rd year	0.142	0.127	0.097	0.151	0.073
	5th year	0.132	0.113	0.090	0.141	0.056

Source: Authors' calculations, based on IABS 1975-2004.

Note: The flow definitions are in Table A.1. The 1st quintile represents the bottom 20% of the wage distribution, the 5th quintile represents the top 20%.

Table A.6: Estimated hazard rates - EE, EU, EN

	EE		EU		EN	
	(1)	(2)	(1)	(2)	(1)	(2)
Female	0.8567***	0.8575***	0.8609***	0.8597***	0.7658***	0.7670***
	(0.006)	(0.006)	(0.009)	(0.009)	(0.003)	(0.003)
Mediumskill	0.7909***	0.7939***	0.5615***	0.5634***	0.5250***	0.5236***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.003)	(0.003)
Highskill	0.7002***	0.7005***	0.2680***	0.2668***	0.4516***	0.4544***
	(0.011)	(0.011)	(0.007)	(0.007)	(0.004)	(0.004)
20-99 empl.	0.9739***	0.9737***	0.7669***	0.7681***	0.7869***	0.7960***
	(0.007)	(0.007)	(0.008)	(0.008)	(0.005)	(0.005)
100-999 empl.	0.7634***	0.7628***	0.5392***	0.5412***	0.6124***	0.6897***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.005)	(0.005)
$\geq$ 1000 empl.	0.6901***	0.6892***	0.5329***	0.5371***	0.5879***	0.5967***
	(0.007)	(0.007)	(0.009)	(0.009)	(0.007)	(0.007)
pos. cohort effect	0.9986***	1.0000	0.9997*	1.0001	0.9961***	0.9906***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
neg. cohort effect	1.0017***	1.0043***	1.0049***	1.0063***	1.0046***	1.0010***
	(0.000)	(0.000)	(0.000)	(0.0003)	(0.000)	(0.000)
pos. cohort effect		0.9999***		0.9999*		1.0002***
*empl. dur.		(0.000)		(0.0003)		(0.000)
neg. cohort effect		0.9998***		0.9996***		1.0002***
*empl. dur.		(0.000)		(0.0003)		(0.000)
	Baseline Hazard					
	Reference Category: 0-6 months employment					
7-12 months	0.8759***	0.9017***	0.9985	1.0455***	0.7118***	0.6875***
	(0.007)	(0.007)	(0.009)	(0.010)	(0.004)	(0.004)
13-18 months	0.8932***	0.9451***	0.7408***	0.8122***	0.6901***	0.6430***
	(0.008)	(0.009)	(0.009)	(0.011)	(0.005)	(0.005)
19-24 months	0.7084***	0.7696***	0.5353***	0.6137***	0.6261***	0.5629***
	(0.008)	(0.009)	(0.009)	(0.011)	(0.005)	(0.005)
25-36 months	0.7414***	0.8351***	0.4818***	0.5885***	0.5247***	0.5833***
	(0.007)	(0.010)	(0.007)	(0.011)	(0.005)	(0.004)
37-60 months	0.6479***	0.7828***	0.3576***	0.4961***	0.4703***	0.4570***
	(0.007)	(0.012)	(0.006)	(0.012)	(0.007)	(0.004)
No. of observations	141,902		77,640		292,461	

Source: Authors' calculations, based on IABS 1975-2004.

Note: Standard errors in parentheses. Significant levels: *: 10%, **: 5%, ***: 1%. Each regression includes region, year, and month dummies. EE: employer-to-employer flows, EU: employment- to-unemployment flows, EN: employment-to-nonparticipation flows (see Table A.1).

Table A.7: Estimated hazard rates - EE^D, EE^{ID}

	EE^D		EE^{ID}	
	(1)	(2)	(1)	(2)
Female	0.8527***	0.8529***	0.7128***	0.7123***
	(0.005)	(0.005)	(0.013)	(0.013)
Mediumskill	0.7717***	0.7758***	0.5136***	0.5158***
	(0.006)	(0.006)	(0.010)	(0.010)
Highskill	0.6189***	0.6196***	0.1863***	0.1859***
	(0.008)	(0.008)	(0.010)	(0.010)
20-99 empl.	0.9429***	0.9430***	0.8746***	0.8749***
	(0.006)	(0.006)	(0.017)	(0.017)
100-999 empl.	0.7475***	0.7475***	0.6610***	0.6616***
	(0.005)	(0.005)	(0.014)	(0.014)
$\geq$ 1000 empl.	0.6781***	0.6786***	0.7271***	0.7300***
	(0.006)	(0.006)	(0.020)	(0.021)
pos. cohort effect	0.9976***	1.0003	0.9997*	1.0000
	(0.000)	(0.000)	(0.000)	(0.001)
neg. cohort effect	1.0029***	1.0063***	1.0037***	1.0068***
	(0.000)	(0.000)	(0.000)	(0.001)
pos. cohort effect		0.9998***		0.9997*
*empl. dur.		(0.000)		(0.000)
neg. cohort effect		0.9998***		0.9998***
*empl. dur.		(0.000)		(0.000)
	Baseline Hazard			
	Reference Category: 0-6 months employment			
7-12 months	0.7351***	0.7651***	0.5180***	0.5576***
	(0.005)	(0.005)	(0.010)	(0.011)
13-18 months	0.6545***	0.7077***	0.3566***	0.4120***
	(0.005)	(0.006)	(0.009)	(0.011)
19-24 months	0.5063***	0.5682***	0.2437***	0.3014***
	(0.005)	(0.006)	(0.008)	(0.011)
25-36 months	0.4897***	0.5789***	0.2024***	0.2750***
	(0.004)	(0.006)	(0.006)	(0.010)
37-60 months	0.3377***	0.4449***	0.0912***	0.1501***
	(0.003)	(0.006)	(0.003)	(0.008)
No. of observations	153,061		38,485	

Source: Authors' calculations, based on IABS 1975-2004.

Note: Standard errors in parentheses. Significant levels: *: 10%, **: 5%, ***: 1%. Each regression includes region, year, and month dummies. EE^D: direct employer-to-employer flows, EE^{ID}: indirect employer-to-employer flows (see Table A.1).

Table A.8: Distribution of stayers and movers

	Stay	EE^D (>1)	EE^D (=1)	EE^{ID} (>1)	EE^{ID}(=1)	EE^D+EE^{ID}
1980	0.381	0.155	0.203	0.030	0.080	0.150
1981	0.374	0.153	0.191	0.037	0.088	0.157
1982	0.393	0.150	0.186	0.037	0.083	0.152
1983	0.382	0.159	0.192	0.034	0.084	0.148
1984	0.412	0.134	0.212	0.027	0.072	0.143
1985	0.419	0.140	0.217	0.021	0.068	0.134
1986	0.421	0.145	0.221	0.019	0.056	0.139
1987	0.440	0.139	0.220	0.015	0.057	0.128
1988	0.446	0.141	0.230	0.014	0.050	0.118
1989	0.430	0.154	0.237	0.014	0.046	0.119
1990	0.441	0.139	0.237	0.015	0.045	0.122
1991	0.472	0.120	0.225	0.019	0.054	0.109
1992	0.467	0.113	0.224	0.020	0.059	0.118
1993	0.472	0.091	0.225	0.022	0.063	0.128
1994	0.466	0.103	0.227	0.023	0.060	0.120
1995	0.449	0.105	0.229	0.023	0.068	0.127
1996	0.447	0.117	0.223	0.024	0.060	0.128
1997	0.432	0.118	0.244	0.017	0.061	0.128
1998	0.440	0.123	0.231	0.018	0.059	0.129
1999	0.430	0.114	0.251	0.018	0.069	0.118

Source: Authors' calculations, based on IABS 1975-2004.

Note: See notes to Table A.5.

Table A.9: Wage growth by quintile

Deviation from	Average Wage growth			
entry wage	Stay	EE^D	EE^{ID}	EE^D+EE^{ID}
1st quintile	1.692 (3.096)	2.063 (2.932)	1.886 (3.063)	1.926 (2.009)
2nd quintile	0.686 (0.717)	0.843 (0.771)	0.589 (0.820)	0.641 (0.792)
3rd quintile	0.380 (0.571)	0.475 (0.581)	0.263 (0.581)	0.320 (0.571)
4th quintile	0.217 (0.473)	0.301 (0.484)	0.077 (0.500)	0.138 (0.479)
5th quintile	0.112 (0.389)	0.160 (0.403)	-0.075 (0.415)	-0.017(0.420)

Source: Authors' calculations, based on IABS 1975-2004.

Note: See notes to Table A.5.

Table A.10: Wage by quintile

Deviation from	Average Wage Changes				
entry wage	At entry	Stay	EE^D	EE^{ID}	EE^D+EE^{ID}
1st quintile	3.090 (0.441)	3.982 (0.786)	4.117 (0.611)	3.816 (1.020)	3.898 (0.888)
2nd quintile	3.532 (0.254)	4.053 (0.687)	4.125 (0.584)	3.764 (1.052)	3.847 (0.957)
3rd quintile	3.791 (0.209)	4.113 (0.666)	4.168 (0.617)	3.795 (1.056)	3.900 (0.932)
4th quintile	4.001 (0.185)	4.195 (0.668)	4.251 (0.612)	3.816 (1.140)	3.918 (0.988)
5th quintile	4.281 (0.201)	4.379 (0.638)	4.408 (0.600)	3.963 (1.080)	4.026 (1.001)

Source: Authors' calculations, based on IABS 1975-2004.

Note: See notes to Table A.5.

Table A.11: OLS-Estimation of cohort wage effects five years after labour market entry

	Coeff.	(S. D.)		Coeff.	(S. D.)		Coeff.	(S. D.)
Stay	0.053***	(0.006)	EE^{D}	0.083***	(0.007)	EE^{ID}	-0.135***	(0.010)
1980	-0.184***	(0.002)	1980*EE^{D}	0.044***	(0.004)	1980*EE^{ID}	0.053***	(0.007)
1981	-0.174***	(0.002)	1981*EE^{D}	0.041***	(0.004)	1981*EE^{ID}	0.051***	(0.007)
1982	-0.164***	(0.002)	1982*EE^{D}	0.053***	(0.005)	1982*EE^{ID}	0.077***	(0.007)
1983	-0.153***	(0.002)	1983*EE^{D}	0.082***	(0.005)	1983*EE^{ID}	0.101***	(0.007)
1984	-0.126***	(0.002)	1984*EE^{D}	0.048***	(0.004)	1984*EE^{ID}	0.092***	(0.008)
1985	-0.093***	(0.002)	1985*EE^{D}	0.029***	(0.004)	1985*EE^{ID}	0.049***	(0.008)
1986	-0.058***	(0.002)	1986*EE^{D}	0.012***	(0.004)	1986*EE^{ID}	0.041***	(0.009)
1987	-0.040***	(0.002)	1987*EE^{D}	0.024***	(0.004)	1987*EE^{ID}	0.040***	(0.009)
1988	-0.007***	(0.002)	1988*EE^{D}	0.013***	(0.004)	1988*EE^{ID}	0.053***	(0.010)
1989	0.016***	(0.002)	1989*EE^{D}	-0.002	(0.004)	1989*EE^{ID}	0.027**	(0.011)
1990	0.046***	(0.002)	1990*EE^{D}	-0.025***	(0.004)	1990*EE^{ID}	-0.085***	(0.008)
1991	0.075***	(0.002)	1991*EE^{D}	-0.049***	(0.004)	1991*EE^{ID}	0.002	(0.010)
1992	0.108***	(0.002)	1992*EE^{D}	-0.063***	(0.004)	1992*EE^{ID}	-0.062***	(0.010)
1993	0.098***	(0.003)	1993*EE^{D}	-0.041***	(0.005)	1993*EE^{ID}	-0.051***	(0.010)
1994	0.113***	(0.003)	1994*EE^{D}	-0.044***	(0.005)	1994*EE^{ID}	-0.093***	(0.011)
1995	0.123***	(0.003)	1995*EE^{D}	-0.045***	(0.005)	1995*EE^{ID}	-0.096***	(0.010)
1996	0.103***	(0.003)	1996*EE^{D}	-0.008	(0.005)	1996*EE^{ID}	-0.073***	(0.011)
1997	0.099***	(0.003)	1997*EE^{D}	0.007	(0.005)	1997*EE^{ID}	-0.062***	(0.012)
1998	0.120***	(0.003)	1998*EE^{D}	-0.020***	(0.005)	1998*EE^{ID}	-0.037***	(0.011)
1999	0.099***	(0.003)	1999*EE^{D}	-0.057***	(0.005)	1999*EE^{ID}	-0.024**	(0.011)

Source: Authors' calculations, based on IABS 1975-2004.

Note: Dependent variable is the log real daily wage. Regression also includes gender, skill level, establishment size, industry and region. Cohort effects are calculated as deviations from the grand mean wage.

Table A.12: IV-Estimation of cohort wage effects five years after labour market entry

	Coeff.	(S. D.)		Coeff.	(S. D.)		Coeff.	(S. D.)
Stay	0.039***	(0.004)	EE^D	0.405***	(0.022)	EE^{ID}	-0.444***	(0.024)
1980	-0.151***	(0.006)	$1980{*}EE^D$	0.023**	(0.011)	$1980{*}EE^{ID}$	0.086***	(0.012)
1981	-0.114***	(0.005)	$1981{*}EE^D$	0.059***	(0.011)	$1981{*}EE^{ID}$	0.044***	(0.012)
1982	-0.069***	(0.006)	$1982{*}EE^D$	0.058***	(0.011)	$1982{*}EE^{ID}$	0.059***	(0.012)
1983	-0.057***	(0.006)	$1983{*}EE^D$	0.040***	(0.012)	$1983{*}EE^{ID}$	0.077***	(0.013)
1984	-0.039***	(0.006)	$1984{*}EE^D$	0.037***	(0.012)	$1984{*}EE^{ID}$	0.029**	(0.013)
1985	-0.015**	(0.007)	$1985{*}EE^D$	0.006	(0.013)	$1985{*}EE^{ID}$	0.065***	(0.014)
1986	0.020***	(0.007)	$1986{*}EE^D$	-0.049***	(0.014)	$1986{*}EE^{ID}$	0.055***	(0.014)
1987	0.038***	(0.007)	$1987{*}EE^D$	-0.036**	(0.014)	$1987{*}EE^{ID}$	0.037***	(0.015)
1988	0.037***	(0.007)	$1988{*}EE^D$	-0.027	(0.015)	$1988{*}EE^{ID}$	0.026	(0.015)
1989	0.036***	(0.006)	$1989{*}EE^D$	-0.050***	(0.015)	$1989{*}EE^{ID}$	0.050	(0.015)
1990	0.035***	(0.006)	$1990{*}EE^D$	-0.025**	(0.011)	$1990{*}EE^{ID}$	0.025*	(0.012)
1991	0.030***	(0.005)	$1991{*}EE^D$	-0.033*	(0.016)	$1991{*}EE^{ID}$	0.032	(0.016)
1992	0.021***	(0.005)	$1992{*}EE^D$	-0.022	(0.017)	$1992{*}EE^{ID}$	-0.015	(0.017)
1993	0.015***	(0.005)	$1993{*}EE^D$	-0.014	(0.016)	$1993{*}EE^{ID}$	-0.004	(0.019)
1994	0.026***	(0.006)	$1994{*}EE^D$	-0.023	(0.018)	$1994{*}EE^{ID}$	-0.099***	(0.019)
1995	0.019***	(0.006)	$1995{*}EE^D$	-0.030	(0.018)	$1995{*}EE^{ID}$	-0.107***	(0.021)
1996	0.019***	(0.007)	$1996{*}EE^D$	-0.015	(0.015)	$1996{*}EE^{ID}$	-0.024	(0.021)
1997	0.064***	(0.008)	$1997{*}EE^D$	0.047**	(0.019)	$1997{*}EE^{ID}$	-0.148***	(0.024)
1998	0.061***	(0.009)	$1998{*}EE^D$	0.052***	(0.020)	$1998{*}EE^{ID}$	-0.187***	(0.024)
1999	0.024	(0.025)	$1999{*}EE^D$	0.004	(0.012)	$1999{*}EE^{ID}$	-0.104***	(0.021)

Source: Authors' calculations, based on IABS 1975-2004.

Note: See notes to Table A.11.

Appendix B Figures

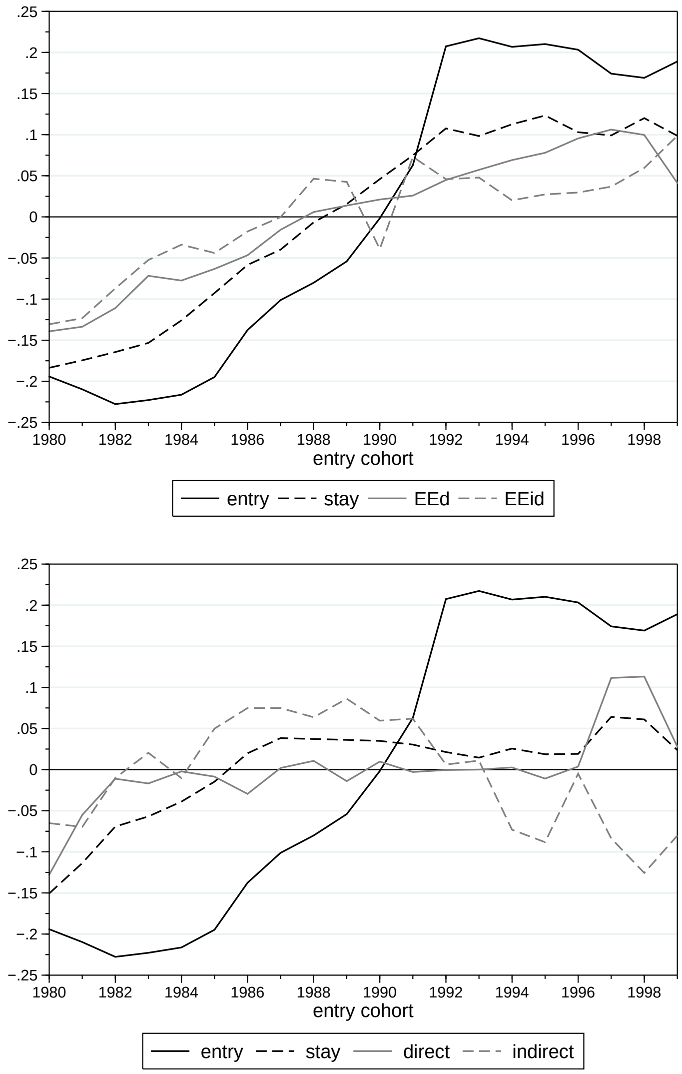
Figure B.1: Wages by year of labour market entry

Source: Authors' calculations, based on IABS 1975-2004.
Note: The grey broken lines show the evolution of wages for cohorts entering the labour market between 1980 and 1999. The black solid and the black broken lines show cohort wages at labour market entry and five years after labour market entry, respectively.

Figure B.2: Starting wages and GDP growth

Source: Authors' calculations, based on IABS 1975-2004.
Note: Wages are detrended by using a Hodrick-Prescott (HP) filter. Following Ravn and Uhlig (2002) we apply a HP smoothing parameter value of 6.25 for our yearly data.

Figure B.3: Estimated cohort effects five years after labour market entry



Source: Authors' calculations, based on IABS 1975-2004.

Note: See notes to Table A.11.