Job Stability Trends and Labor Market (Re-)Entry in West Germany 1984-1997

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Abstract

This paper investigates whether job stability in western Germany shows any signs of decline and compares the findings to evidence for the US and the UK. Cross sectional data and calendar information from the German Socioeconomic Panel 1984-1997 are combined allowing to check possible influences of oversampling long jobs in cross sectional data. Three different measures are looked at. All indicate that there is a decline in job stability, not fully explained by the business cycle: median elapsed tenure of male workers declined from around 10 years to 8.5, the probability to be in short jobs seems to increase relatively steadily for both males and females, and the hazard for job ending has become increasingly higher despite the fact that the economy experienced the post-unification boom and the current recession.

Cox proportional hazard models for different groups in the labor market show that men and women are equally affected. Part-time workers, although generally more likely to end their job, have suffered less. As 'outsiders' are more likely to have difficulties finding stable jobs in rough times separate analyses are carried out those who have entered the job directly from unemployment or non-participation and workers who enter the labor market having just finished their highest degree. These are compared to the 'insiders' who switch jobs directly. While 'insiders' are less likely to leave their new job, outsiders face increasing risks of job termination.

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1. Introduction

In Germany the casual public sentiment is that job security is on the decline and that jobs have become less stable since the 70s due to structural changes in labor markets. It is usually argued that globalization and technological progress have increased the need for flexibility leading to less secure jobs for workers. Obviously secure and stable jobs are desirable at least from the workers point of view, as long as no better employment opportunities arise. On the other hand too much stability can be harmful if necessary changes in the labor market are delayed. Job stability is discussed in western Europe as well as in the United States and while studies for the US and the UK present some limited evidence of increasing job instability (compare Burgess and Rees 1996, 1997, 1998; Farber 1995; Swinnerton and Wial 1995; Diebold et al. 1997) the rare empirical evidence for Germany is very controversial. Winkelmann and Zimmermann (1997) report decreasing numbers of job changes as evidence for even an increase in job stability. Bergemann and Schneider (1998), however, find some support for declining job stability in Germany. They present some descriptive statistics for jobs that started within a certain period, so their results cannot be directly compared to international evidence that looks at the tenure distribution for all jobs in the economy. Moreover, the group of job starters is very heterogeneous and different tendencies could exist for 'insiders' who only switch jobs and 'outsiders' who try to (re-)enter the labor market from unemployment, non-participation or training and education. This study adds to the literature by presenting a collection of cross sectional measures for overall job stability in Germany and looking at the different groups who start new jobs in the period 1984-1997 within duration models.

Previous studies are usually based only on cross sectional tenure information of those currently in work (with one exception being Booth et al. 1996). It is well known that this type of analysis undersamples short jobs. This is also a likely problem in the papers by Bergemann and Schneider (1998) and Bergemann (1999), who only used information on jobs existing at the time of the interview. Therefore, longitudinal information of the so-called "calendar" of the German Socioeconomic Panel (GSOEP) will be combined with cross sectional information, to see whether using only information of those currently in work influences the analysis. All currently available waves of the GSOEP between 1984 and 1997 are used.

The paper is set up as follows. Section 2 reviews the literature on job stability changes and shortly discusses the possible theoretical linkages between structural changes and job stability. In Section 3 some cross sectional evidence on elapsed tenure stratified by age, sex and industry is presented and compared to the most recent literature on job stability that focuses on the United States and the UK primarily. Multivariate analysis is introduced as well to show whether the probability to be in short or very long jobs has changed. The main hypothesis under investigation is whether average job tenure decreased since re-unification in West Germany. Section 4 takes the analysis further and looks at job stability for workers starting new jobs in a multivariate framework: men and women, part-time and full-time workers, job switchers and possible outsiders like labor market entrants and workers entering from unemployment or non-participation. Longitudinal data from the employment calendar in the GSOEP is combined with cross sectional information on important covariates to estimate Cox proportional hazard models, which directly take into account the right censoring of the tenure variable. Section 5 summarizes the findings and concludes.

2. A Decline in Job Stability - Evidence from the Literature

While the effects of structural change, "globalization" and technological progress on wage differentials have been of some concern in recent years[1], we hardly know anything about the consequences for job stability patterns. A decline in job stability, however, could certainly be part of the adaptation mechanism to structural changes and exogenous shocks, especially for sensitive groups in the labor market like unskilled workers and unemployed workers looking for new jobs (compare Farber 1995). Job stability can be measured in several different ways, but the most frequently used operationalizations are elapsed or completed tenure with one firm, that is job changes or promotions within firms are usually not looked at.[2]

We know from the literature that these job switches and therefore "unstable" jobs are not necessarily a bad thing. Workers who switch jobs voluntarily often improve their earnings or work conditions of their jobs (Jovanovic 1979).[3] We also know that (voluntary) quits are procyclical, while (involuntary) layoffs are countercyclical both influencing the tenure distribution. In boom periods more new jobs are created leading automatically to more jobs with short duration and, therefore, decreasing tenure usually even if layoffs are reduced. In recessions there will be more layoffs, less quits and average tenure is likely to increase as new hires are rare (Burgess and Rees 1996, Schettkatt 1996).

[1] See e.g. Katz and Murphy 1992; Levy and Murnane 1992; Krugman 1994; Leamer 1994 and 1996; Abraham and Houseman, 1995; Freeman and Katz 1995.
[2] As workers in large firms have more career possibilities within their firm it is usually observed that workers from large firms change their employer less often then workers from small employers. Therefore, firm size happens to be an important covariate to be included in the following multivariate studies.
[3] See Mertens (1998) for an overview of the theoretical and empirical literature on job mobility.

However, the distribution of job tenure in an economy seems to change very slowly so that business cycle influences seem to be of minor importance in comparison with structural conditions and institutional settings. Burgess and Rees (1996) report that although business cycle phenomena exist, changing the distribution of elapsed tenure, there is no long term trend. The same result is found when looking at the estimated distribution of completed tenure as indicated by retention rates. These are the probabilities that a worker with a certain age and tenure will retain her job for one, two, or more years (Burgess and Rees 1997).[4] Booth et al. (1996) analyze the tenure pattern using the BHPS retrospective data over a long time period. They show that since the 1950s there has been a small decline in job tenure in the UK.

Similar results are reported for the United States by Farber (1995) and Dieboldt et al. (1996, 1997). Only Swinnerton and Wial (1995, 1996) reported declines in job stability while Diebold et al. (1996,1997) show that tenure distributions are rather stable over time.[5] Finally, Farber (1995) reports that men are increasingly less likely to be observed in long jobs while women's probability of being in long jobs increased significantly. According to Farber's estimates much of the negative change for men is driven by a decreasing probability for male workers with low education to be in long jobs, i.e. those groups in the labor market that experienced the greatest decline in relative wage rates during the 1980s. Female workers in the lowest educational group experience the smallest increase in the probability of being in long jobs.

That there probably is less job stability in the U.S. can also be seen from studies on worker displacement.[6] These show that job loss increased since the 1970's affecting more and more also high tenure and white collar workers. Especially the facts reported by Farber (1997) cast some doubts on the notion that job stability did not change: job loss in the recession of 1981-1983 accounted to around 13% of the workforce. The three year rate of job loss decreased until the period 87-89 and then rose to the highest level since 1981: 15% of the workforce lost their job in a period of expansion between 1993 and 1995. As Kletzer (1998) writes, "These high rates of job loss are consistent with public perceptions of rising job insecurity" (see also DiPrete 1993). Even Dieboldt et al. (1996,1997) report for the U.S. that

[4] Hall (1982) was the first to estimate historical rates from CPS data by comparing the number of workers in an age-tenure category in one survey (e.g. those 25-29 years old with 0-5 years of tenure in 1968, say 5000) with the number in a later survey in correspondingly higher age and tenure categories (e.g. those 35-39 years old with 10-15 years of tenure in 1978, say 2000). The historical retention rate in this case equals 2000/5000 = 0.4, i.e. 40% of those between 25 and 29 in 1968 will hold jobs that will last for ten more years. Ureta (1992), Swinnerton and Wial (1995) and Diebold et al. (1996,1997) extended the analysis to later years and refined the estimation procedure.

[5] For recent overviews of the literature in the U.S. see Schmidt and Svorny (1998) or Jaeger and Stevens (1998).

[6] Another strand of literature not reviewed here is concerned with gross job and workers flows looking at job creation and destruction at the firm level (see Davis and Haltiwanger 1992; Leonard and Schettkatt 1991).

"Although the differences in the changes in retention rates by age group are not dramatic, it is nonetheless interesting that the decline is greatest for young workers, who also experienced the largest relative wage decrease" (p. 219). Gregg and Wadsworth (1995) have similarly argued for the UK that entry positions available to those currently not in employment might have become increasingly unstable and low paid. These studies suggest that looking at the most sensitive groups of the labor market in more detail may yield helpful insights into the question of whether job stability decreased. However, the next Section 3 will first look at general trends in job stability in western Germany to compare them to the international evidence. Section 4 then goes on to analyze more closely workers starting jobs within the time period under investigation like workers (re-)entering from unemployment or non-participation as well as workers entering from education and training.

3. General Trends in the Tenure Distribution

Between 1984 and 1989 the West German economy recovered from a recession in the early 80s leading to falling unemployment rates and slightly better job prospects for workers as can be seen from Figure 1. In 1990, the year of re-unification, however, there has been a pronounced boom bringing capacities in West Germany to their limits. This was primarily due to increasing demand for West German products in East Germany. This boom ended dramatically in 1993 followed by a recession that has not really come to an end, yet. Unemployment in West Germany has since then increased and growth rates are at relatively low levels. The following analysis will therefore compare the immediate post-reunification period 1990-92 and the later recession period 1993-97 with job stability from 1984-89.

As described in the preceding section one natural way to measure job stability is elapsed firm tenure or job duration. Using this simple measure some trends for western Germany will be described here. All calculations are based on the two original west German samples of the German Socioeconomic Panel (GSOEP) which includes information on elapsed firm tenure for all workers employed at the time of the interview.[7] Table 1 shows the well known fact that median tenure differs significantly by gender. While men's median tenure starts with 10 years in 1984 and ends with 8.5 years in 1997 women's median tenure is relatively constant at 6 years.

[7] As households with non-German heads are heavily oversampled in the GSOEP all frequencies and medians reported are weighted by the cross sectional sample weight made available in the GSOEP. For further information on the data set see the appendix.

Now, it is important to distinguish different age groups because obviously older workers are able to accrue longer tenure than young workers. Age is therefore used as a (non-ideal) proxy for labor market experience and median tenure is reported for the groups 16-25 years, 26-45 years and 46-65 years in Figure 2. The difference between men and women is evident only in the older age groups. Younger female workers between 16 and 25 have tenure comparable to men. What is more striking, however, is that median tenure decreases for the young and middle age groups between 25 and 45 but increases especially for male workers between 46 and 65. The female experience is strikingly different to results reported by Farber (1995), Marcotte (1995) and Burgess and Rees (1998) for the United States and the UK. It seems that while women in the US and the UK were able to accrue longer tenure over time, German women were only able to hold the level already acquired in the mid-1980s.

Another interesting detail of the tenure pattern has been pointed out by Gregg and Wadsworth (1995). They show for the UK that considerable variation in tenure by employment type can exist with increasing separation probabilities for part-timers. Therefore, Figure 3 shows average tenure by regular hours worked. Men in part-time jobs and marginal employment[8] obviously have lower median tenure than full-time workers, however only around 3% of male workers can be found in these categories (own calculations from the GSOEP, see also Hoffmann and Walwei 1998). Median tenure calculated for other than full-time male workers is very erratic and the observed decline in median tenure therefore seems to be due to decreasing median tenure in full-time employment.[9] As would have been expected women are more frequently found in part-time work (around 30%). It is interesting to see that this type of work is even associated with slightly higher median tenure than full-time work and there are no clear tendencies over time for all employment types.

Similarly, tenure by industry differs more strongly for men than for women as can be seen from Figure 5 where the most important sectors are depicted. In trade, services and manufacturing median tenure decreased between 1984 and 1997 by 1 to 2 years. In the state sector median tenure even increased. This is probably due to reduced hiring by the state.

Another possibility to analyze job stability is to look at the tenure distribution in a multivariate context originally proposed by Farber (1995) and extended by Burgess and Rees (1998). In order to measure changes in the distribution of job duration, dummy variables

[8] There are strict definitions for marginal employment in Germany. Working either below 15 hours or receiving monthly wages of only 620 DM (margin since 1998).

[9] In 1988 around 3000 male workers are observed in full-time employment and only around 60 in part-time or marginal employment.

indicating whether a worker is in a job with duration of one year or less (respectively morethan ten years or more than twenty years) are regressed on a set of covariates in a discrete choice model. Following Farber (1995) Logit estimations are presented, however using Probits does not change the pattern of the results. All employed individuals between 16 and 64 are selected for the logit of the probability of job duration one year or less. For the probability of a job duration of more than ten or twenty years the sample only includes workers who are at least 35 or 45 years old respectively. Moreover, self-employed, trainees and workers in agriculture (including forestry and fisheries), non-profit organizations and observations with missing values on at least one covariate were deleted from the sample.[10]

The covariates included here are age dummy variables for age categories (16-19, 20-24, 25-29, 30-34, and so on), dummy variables for educational attainment, changes in the national unemployment rate, firm size dummies, a part time indicator, white and blue collar workers dummies and industry dummies. All these variables are known to influence employment prospects and mobility patterns (for an overview see e.g. Mertens 1998). There are two different ways to test whether the probability to be in short jobs increased while the probability to be in long jobs decreased, both indicating a decline in job stability. Following Farber (1995) a linear time trend variable is included first. Results are presented in Table 2 columns I, IV and VI. A complete set of results can be found in Appendix Tables A1-A3. For both men and women the probability to be in short jobs has increased significantly, while the probability to be in jobs that last at least ten years has decreased. There is no significant change for jobs that last at least twenty years, i.e. for very long jobs corresponding to findings that elapsed tenure for the oldest group of workers did not decrease (see Figure 2).

As we are especially interested in the question of whether job stability declined in the post-reunification period a dummy variable approach is followed next. Column 2 shows that a dummy variable separating the post-reunification years 1990-1997 from the pre-unification period in our data set (1984-1989) is positive and significant for short jobs and negative significant for long jobs. However, the post-reunification period can be separated into a boom period (1990-1992) and a recession period (1993-1997) as already argued in the introduction (see Figure 1). In a boom period it should generally be the case that the probability to be in short jobs increases, as there are more new jobs leading to workers quitting more often and unemployed workers having better chances to find a job. In recessions, however, the opposite should be the case. If we find increasing probabilities to be in short jobs in a recession this could indeed be an indication of declining job stability and security for workers. Columns III,

[10] Including all sectors does not influence the central results of the analysis.

V and VII report the results of this exercise: in the latest period the probability to be in short jobs is higher than in the boom period following reunification. The probability to be in long jobs only decreased in the recession period. There is no theory to explain that such a development could be caused by voluntarily job terminations in a recession. It is more likely that involuntary separations are the major reason, but this will be discussed in more detail in the next section. Considering that this type of analysis does not account for the fact that there has been an increasing number of people looking for work, the result seems even more worrying.

The results presented so far suffice to give a first overview of what has been going on with respect to overall job stability in western Germany. It has been shown that especially younger and middle aged workers are affected by a decline in job stability. Moreover, the logit anaylsis has shown that jobs seem to end earlier. This question, however, can better be analyzed using duration analysis. Therefore, the next section goes one step further and focuses on the question of whether jobs that start after unification are likely to end earlier than jobs that started earlier. Having shown that older workers and jobs that have been going on for a very long time are not affected by a decline in job stability this is a reasonable procedure.

4. Do new jobs tend to end earlier in recent years?

In this section job stability will be studied collecting evidence on job termination directly. For Germany there exist few studies looking at similar problems. Winkelmann and Zimmermann (1997) estimate count data regression models from the German Socioeconomic Panel. They regress the number of job changes between 1974 and 1994 on a set of individual characteristics, however, they are interested in the general trend comparing the 70s with the 90s rather than pre- and post-unification. They find that the coefficient on a time dummy indicating the period between 1984 and 1994 is significantly negative, i.e. controlling for individual characteristics the number of job moves has decreased, which might be interpreted as an increase in job stability rather than a decrease. The analysis most related to this work is a study parallely developed by Bergemann (1999). While this study here focuses on the different groups in the labor market, hers focuses on the different reasons for job terminations. However, there are two drawbacks of her use of the GSOEP data. First, she does not use the spell data given in the GSOEP calendar that collects information on every month in the panel but rather extracts duration of jobs going on at the interview period only. Therefore, she cannot be sure whether the undersampling of short jobs biases her results seriously. The second problem arises from the same fact: looking at the different reasons for separations she has to match information on "Why did your previous job end" with job information from the last panel wave. Obviously there will be cases in which this is not the correct job as there might have been a short job in-between. This study overcomes these problems and shows that the general tendency of results is still the same.

The Cox Proportional Hazard Model

Cox proportional hazard models will be used to study the duration of jobs. This model has been frequently used in mobility analyses (see e.g. Mayer and Caroll 1990). Similarly to Booth et al. (1996) dummy variables indicating the time periods are included in the estimation controlling for a set of covariates known to influence mobility behavior. As already discussed in Section 3 three different time periods in Germany are distinguished: pre-unification (1984-1989) and post-unification separated into boom (1990-1992) and recession (1993-1997).

In the empirical analysis the dependent variable is the hazard rate $\lambda(t)$:

$$\lambda(t) = \lim_{\Delta \to 0} \frac{\Pr(t \leq T \leq t + \Delta \mid T \geq t)}{\Delta} = \frac{f(t)}{S(t)} \qquad (1)$$

where Pr(.) is the probability of leaving the job with one employer between t and $t+\Delta$ given that the job is still held at time t (see for this and the following e.g. Blossfeld et al. 1986, 1989 or Kiefer 1988). This probability is also defined by the density function $f(t)$ divided by the survival function $S(t)$. In the following the popular proportional hazard assumption will be employed to analyze the influence of important covariates x on hazard rate:

$$\lambda(t,x,\beta,\lambda_0) = \phi(x,\beta)\,\lambda_0(t) \qquad (2)$$

where β is the unknown parameter vector and λ_0 is the "baseline" hazard corresponding to $\phi(.) = 1$. If $\phi(.) = 1$ is measured at the mean of the regressors then λ_0 has an interpretation as the hazard function for the mean individual in the sample. Using the partial-likelihood approach originally suggested by Cox (1972) the coefficient vector β can be estimated without specifying the form of the baseline hazard function λ_0:

$$\lambda(t,x,\beta,\lambda_0) = \exp(x'\beta)\,\lambda_0(t) \tag{4}$$

One of the advantages of hazard rate models is that right censoring which limits the informational content of elapsed tenure in cross sectional analyses is explicitly dealt with (see Blossfeld et al. 1986 or Kiefer 1988).[11] In the remaining part of this section the job duration of all jobs beginning since 1984 will be analyzed. Special groups proposed to be likely 'victims' of a decrease in job stability like workers entering from unemployment, non participation or training and schooling will be investigated in turn.

Data

Like in the preceding section data from the GSOEP is used, but this time the calendar information, where all interviewed persons are asked about their major occupations on a monthly basis. The GSOEP is a yearly panel so this data base includes valuable information on job spells, especially also those spells that only last a few months and neither coincide with the interview month in the current year nor in the previous year. Additional information on job changes that are collected at each interview are used to split employment spells if necessary.[12] In the final sample only those spells were included that started since January 1984. Self-employed workers, trainees and civil servants were excluded as were the sectors agriculture (including forestry and fisheries) and private households. Full-time and part-time jobs enter the analysis. Table 3 gives information on the frequencies of all important covariates for the following analyses.

Unfortunately, there is no job information available for the jobs that do not coincide with interviews but it is possible to at least combine the jobs with detailed cross sectional information that coincide with the interview months. Two types of analysis are therefore possible: i) using all available spells and including dummy variables equaling one when the information is missing and ii) using only those spells with complete information on important covariates. Using calendar should in any case be superior to using only cross sectional

[11] Ties are handled using the Peto-Breslow approximation. The estimation used the software package STATA.
[12] Detailed information on the data selection process is given in the appendix.

information from the GSOEP. One major reason is that interview information on the reason for job change can be attributed to the latest observed employment spell even if that job did only last for a short time and did not coincide with the previous interview. The probability to combine job change information with the wrong job is therefore minimized here. Moreover, the calendar includes better information on the labor market status of workers prior to starting the new job, which enables to look at those workers who enter a job directly from unemployment.

The Average of all New Job Spells

Lets first look whether there is a general tendency that jobs starting between 1993-97 end earlier than jobs started in the 1980s or the immediate post-unification period. Table 4 presents the results for a variety of specifications of Cox proportional hazard models. The unemployment rate is included as a time varying variable. Specification I shows results for all 12,564 observed spells in the calendar. Personal characteristics like age at start of the job, sex, schooling and unemployment experience are available for all spells as is the full-time/part-time information. For other covariates missing indicator dummies have to be included. The results show that controlling for the aforementioned influences as well as firm size, industry, blue/white collar status, and the business cycle via the unemployment rate, there seems to be a tendency for increasing risks of job termination as indicated by the two time dummies (1990-1992) and (1993-1997). Reported are hazard ratios or α-coefficients from the Cox model, where $\alpha=\exp(\beta)$ and $(\alpha\text{-}1) \times 100$ equals the percentage effect of a covariate on the job separation risk. While the risk of job termination increased by around 9% in the immediate period following reunification in comparison with the 1980s, the risk increased by 18% in the period 1993-1997. Even controlling for unemployment in the year of job ending, jobs tend to end earlier in later years. This effect also remains true when censoring all observations at a maximum of 6 years duration, so this is not an artifact of the possible different time duration. Simple survivor functions in Figure 5 show a similar picture, although differences there are not very large. The multivariate analysis however is superior and indicates that workers with the same characteristics are more likely to leave their job now than in the mid-1980s. These findings support the logit models estimated in Section 3: job stability in Germany has been declining since the mid-1980s. Part-timers, though in general more likely to end a job, have been influenced less by this development and seem to experience a smaller increase in job terminations. This can be seen from an interaction variable between a part-time dummy and the time period dummies. Therefore, it is important to include this interaction in all specifications.

The following column II shows whether the same result can be obtained using only those observations with valid information on all included explanatory variables and the reason for job termination. While the direction and significance of the time dummy variables and its interactions are the same, the size of the coefficients increases. It therefore seems that using only those jobs with valid information on all variables exaggerates the size of the effect, though not the general trend in the data. However, one has to be extremely careful when interpreting the results. For the rest of this paper only those results using all spells will be reported.

So far the different reasons for job change have not been distinguished, but it is also interesting to know why jobs end, as quits and layoffs tend to have different consequences (for Germany see Burda and Mertens 1998, 1999). The increase in the risk of job termination between 1990-1992 could be explained with increasing separations given new job opportunities, the increase during the latter phase, however, is more likely to be caused by involuntary job endings as recessions and increasing unemployment are unlikely to go hand in hand with increasing voluntary quits. Bergemann (1999) follows Booth et al. (1996) and presents some evidence on this point by looking at the different reasons for job termination within competing risks models. She reports that especially men experience an increase in the risk of job termination by layoffs in the 1990s while the hazard of job termination by quits has not increased significantly. For women other reasons than quits or layoffs seem to be more important (everything from child care to early retirement). The specification presented in column III supports her findings. Including interactions between a layoff dummy and the time dummies it can be shown there is indeed a tendency for an increasing risk of job termination by layoffs. Inclusion of those interactions leaves only the second time dummy for 1993-1997 significant at a lower significance level than before. Again these effects are stronger when only including spells without missing values as can be seen from column IV.

Before turning to an analysis of different labor market groups the other controls in the Cox models will shortly be discussed in turn. The unemployment rate, which is the only variable that is allowed to vary in the analysis is always negative and significant, i.e. it acts as a kind of lagged indicator of the business cycle. When unemployment rates are already high, overall job terminations, voluntary and involuntary, are lower. Personal unemployment experience is only found to be significant at usual significance levels in the analysis where all spells are included. This is what could be expected, assuming that workers with unemployment experience differ in their unobserved characteristics from other workers who never experienced unemployment. Consistent with findings in the mobility literature workers who are older at entry into the labor market have a lower risk of job termination, as younger workers are usually found to be more mobile. Non-German citizens as well as females are found to have lower risk of job termination, however, only after controlling for the other worker and job characteristics. In general woman have higher separation hazards from jobs that last longer than around four years as can be seen from the graph of survivor functions in Figure 6. Females with the same characteristics as men, though, are found to be less mobile, at least in jobs that started since 1984.

Training dummies and occupational position dummies have to be interpreted together as both tell us something about the type of education. Significance varies strongly between the two specifications, but it seems that workers with technical training are less likely to leave a job than workers without training or workers holding a university degree. Much of the training and education effects are also captured in the worker status dummies: all status groups are less likely to end their job than unskilled blue collar workers. As usually found in the literature workers from small firms have higher risks of job termination than workers from medium sized or large firms. Finally, the dummies indicating missing variable status are always positive and significant as they catch much of the effect that short jobs are underrepresented in a cross section.

Workers Switching Jobs

Having looked at the average tendencies we now go on to look at special groups who start new jobs between 1984 and 1997. Workers switching directly from a full-time job into another job will serve as a reference for those workers who enter work from unemployment, non-participation or training. As the number of coefficients estimated is very high detailed results are to be found in the appendix only.[13] Table 5 reports the results on the time dummies and interactions for part time workers. Surprisingly the results turn around: workers switching into another full-time employment are less likely to leave their jobs in the post-unification period (though the coefficient is only marginally significant). On the other hand workers who switch from full-time into part-time employment are more likely to leave their job. Both

[13] See Appendix Table 4.

effects get even stronger when controlling for involuntary separations (see Appendix Table A4). These findings support results of Winkelmann and Zimmermann (1997) who reported a declining number of job switches over time in West Germany.

Summarizing we can already note that it might be misleading to look only at the average tendencies for all new jobs within a period, as workers starting new jobs are obviously very heterogeneous. The following sections will show which groups are the ones that are most seriously affected by an increasing risk in job termination.

Workers Entering From Unemployment or Non-Participation

As Table 3 shows, a considerable part of all workers who start new jobs come directly from unemployment (around 23%) or non-participation (25%). These groups should be of major concern as it is important especially for unemployed workers and workers seeking from non-participation to find their way back into the labor market. Table 5 consequently repeats the analysis for workers entering from unemployment and non-participation in rows III and IV. The unemployment experience dummy in the former case catches whether this has been their first unemployment spell (dummy=0) or at least their second unemployment spell (dummy=1). The other included variables are the same as in the previous analysis (see also Appendix Table A5).

For workers unemployed prior to the present job spell the recession period 1993-97 is of major importance. The estimated coefficient, however, hardly differs from the average reported in Table 4. Again the increase in job instability can be attributed primarily to layoffs as indicated by the interaction variable (and reported in Table 6). Workers entering from non-participation have obviously more problems as their risk of job termination is 25% in the period 1993-1997 in comparison with 18% on average. Considering that most workers entering from non-participation are females it will be interesting to see in a next step whether there is a difference between males and females.

Job Terminations by Females and Males

The analyses are now repeated for males and females and also reported in Table 5. Rows V and VI show that both males and females were obviously affected by a decline in job stability during the 1990s. However, it can be shown that the effect for women does not disappear once controlling for involuntary separations, indicating that other reasons are of greater importance (see Table 6 and Appendix Table A6; also compare Bergemann 1999). The following rows VII and VIII in Table 5 show that the decreasing risk of job termination for full-time working job switchers is driven by male workers experiences. Female full-time workers do not experience any significant change.

Separating again between workers coming from different non-employment status yields further interesting results. Male workers leaving unemployment seem to have increasing and higher difficulties in finding a stable job than others: their risk of job termination in the period 1993-97 is around 3% higher than on average (column IX). Workers who leave non-participation, however, are primarily female and have similarly problems in finding a stable job (see column XII).[14] In combination with the reported results on job switchers this clearly supports the idea that outsiders have less opportunities to get stable positions in the labor market in the 1990s. In the next and final section we therefore look at another group of outsiders, namely those who have just recently finished a degree and have to find an appropriate job now.

Young Workers Having Recently Finished a Degree

In Germany of all workers around 68% have some kind of practical training. Skills acquired during this training are usually thought to be transferable across firms but hardly across occupations. Winkelmann (1996) reports that 70% of all apprenticeship trained workers leave their training firm within 5 years but some occupational mobility is observed even directly following the apprenticeship (compare Allmendinger 1989 and Werwatz, 1998).

In the following we will see whether job duration in the first job after apprenticeship or other forms of vocational training decreased after re-unification in Germany. The definition of labor market entrants follows Winkelmann (1996) and includes three comparative groups: secondary school leavers, young workers who have just received a university degree and workers who have just left school and report full-time or part-time employment as primary occupation. Note that this group of workers *does not necessarily have to enter directly from education*, as there could have been short unemployment spells or non-participation spells in-between. Some graduates who did a degree in the evening while working are also included.

[14] Compare for complete set of results Appendix Table A7.

Only workers entering the labor market until the age of 30 are included. If there is more than a years' gap between an education and a job spell the individual is not included in the sample. Again Table 5 reports the results which show a similar picture: the risk of job ending increases following unification in western Germany, but labor market entrants do not seem to be affected above average (also compare Appendix Table A8) .

5. Conclusions

The analyses have found some support for decreasing job stability in western Germany, however not all groups in the labor market are equally affected. Using repeated cross sections from the German Socioeconomic Panel (GSOEP) it was shown that male workers have experienced a decrease in median tenure from around 10 years to 8.5 years between 1984 and 1997. While women in other countries like the US or the UK were able to accrue higher median tenure on average, there has not been such an increase in Germany between 1984 and 1997. On the contrary, women's probability to be in long jobs decreased just like men's and the probability to be in short jobs increased, which indicates that jobs tend to end earlier in recent years.

The latter questions, which is of major importance can, however, be better dealt with using duration analysis. Moreover, repeated cross sectional analysis, does not tell the whole story, as short jobs are undersampled and the effect of the increasing number of workers who are in and out of very short jobs is therefore neglected. Hence, the duration of jobs has been looked at more closely using job spell information from the GSOEP calendar combined with cross sectional information. Cox proportional hazard models for workers starting new jobs between 1984 and 1997 show that there is indeed a decline in job stability over the nineties that is even stronger in the recession period of 1993-1997 than in the post-reunification boom. The results are thus similar to those obtained using simple cross sectional information. However, disaggregating the job starters into distinct groups shows that not all of them are equally affected.

Considering insider-outsider theories (see e.g. Lindbeck and Snower 1988) it can be predicted that workers who are outside the system will be hurt most when there is a decrease in job stability. This has also been proposed by Gregg and Wadsworth (1995). Along this line of reasoning, separate analyses were performed for different groups of workers who started new jobs between 1984 and 1997. Likely insiders are the ones who switch directly between jobs while more disadvantaged groups should be the ones who have to find a way (back) into stable employment: workers entering from unemployment or non-participation as well as workers finishing a degree. The paper shows some support for this hypothesis. Full-time workers who switch jobs are less likely to leave their job in the 1990s than in the 1980s, but workers entering from unemployment and non-participation are more likely to terminate their jobs. Labor market entrants also experience an increase in the risk of job termination but not as strongly as males who were unemployed prior to the job or women who entered from non-participation.

Appendix

I. The German Socioeconomic Panel

The GSOEP is a panel of approximately 6000 German households and has been conducted annually since 1984. All adults in participating households are interviewed once a year, usually in spring. The questions cover economic and social conditions of all household members. The major topics of interest in this data set are population and demography, education, training and qualification, labor market and occupational dynamics, earnings and income, social security, housing, health, household production and preferences. The original survey started in western Germany and expanded to eastern Germany in 1990. For further detailed information on the data set see German Institute for Economic Research (1996).

Only those individuals living in western Germany were selected, including all available samples A and B. As foreigners are oversampled to allow individual analysis, the reported descriptive cross sectional statistics are weighted by the respective sample weights.

II. Extracting Spell Data from the GSOEP

The duration analysis is performed with the help of two different types of data sets from the GSOEP, a monthly spell data set with information from the calendar and a data set that uses the detailed information on the job at the interview time. How the two different types of information are matched is described in this appendix.

The spell data is derived from the calendar which includes information on the employment status each month. This is the only monthly information in the GSOEP. Other variables like wages, industry or occupation are only covered once a year in the survey. The monthly spell data set includes the following important variables: person identifier, begin of the spell, end of the spell and spelltype. The analysis cannot be build on these spell data alone, because important information on the characteristics of the job and job changes without intervening spells of other activities are missing. Therefore the following cross sectional information is merged to the spell data:

1. Several job characteristics: industry, occupation, firm size, employment status, 2. "Why did your last job end?", 3. "When did you last job end?", 4. "Did you finish training or school since the beginning of last year?", 5. "Which type of degree did you attain?". With the help of this information the spell data was transformed as follows: when a job ending was reported, the respective employment spell was splittet if it was not a within firm job change. One problem with this procedure, however, occurs when workers switch between part-time and full-time employment (or between full-time jobs with intervening breaks) and report the job ending slightly differently than in the calendar. In this case implausible short jobs could be introduced with the splitting procedure, so that it was checked whether a new part-time or full-time employment spell begins within three months. In this case the split was not performed. After splitting the spell data other information was added to the spells by checking the start and ending date of the spell and the time of the interview. That way the values at the start of the job were matched.

Finally, for the purposes of this investigation it was also necessary to get information on the origin and destination state, i.e. the major activities of the individuals post and prior to the job spell. As some activities can be parallel the following hierarchy was introduced to make most use of the data: 1. full-time employment, 2. school, 3. unemployment, 4. part-time, 5. non-participation.

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Tables

Table 1 - Median Tenure in Years

	Men		Women	
	Median tenure	Observations	Median tenure	Observations
1984	10.1	3650	6.1	2177
1985	9.7	3736	5.9	2290
1986	9.8	3546	6.4	2145
1987	9.9	3476	6.8	2108
1988	10	3291	6.6	2097
1989	9.9	3211	6.5	2096
1990	9.9	3176	6.4	2133
1991	9.9	3167	6.5	2167
1992	9.8	3070	6.7	2097
1993	9.6	3005	6.6	2109
1994	9.5	2867	6.8	2030
1995	9.3	2754	6.5	1941
1996	8.8	2715	6.8	1944
1997	8.5	2644	5.8	1910

Note: All observations are weighted by the GSOEP sample weight.
Source: Own calculations based on the GSOEP 1984-1997 (samples A and B)

Table 2 - Logit Analyis: Probability of Elapsed Tenure ≤ 1, ≥ 10 and ≥ 20 years

	All I	All II	All III	Men IV	V	Women VI	VII
Tenure ≤ 1 year [a]							
Time trend	0.051** (0.009)	.	.	0.049** (0.013)	.	0.050** (0.011)	.
1990-1997	.	0.262** (0.055)	.	.	.	.	.
1990-1992	.	.	0.171** (0.062)	.	0.079 (0.086)	.	0.262** (0.092)
1993-1997	.	.	0.397** (0.087)	.	0.462** (0.116)	.	0.330** (0.109)
unemployment growth	-0.723** (0.238)	-0.322 (0.236)	-.778** (0.274)	-1.002** (0.337)	-1.346** (0.383)	-0.336 (0.338)	-0.139 (0.391)
Tenure ≥ 10 year [b]							
Time trend	-0.022** (0.006)	.	.	-0.017* (0.007)	.	-0.030** (0.009)	.
1990-1997	.	-0.102* (0.041)	.	.	.	.	.
1990-1992	.	.	-0.044 (0.046)	.	0.016 (0.061)	.	-0.117 (0.074)
1993-1997	.	.	-0.182** (0.059)	.	-0.154* (0.075)	.	-0.230** (0.088)
unemployment growth	0.119 (0.176)	-0.061 (0.168)	0.214 (0.206)	0.160 (0.236)	0.360 (0.281)	0.080 (0.269)	0.096 (0.310)
Tenure ≥ 20 years [c]							
Time trend	0.003 (0.007)	.	.	0.011 (0.008)	.	-0.003 (0.014)	.
1990-1997	.	-0.021 (0.052)	.	.	.	.	.
1990-1992	.	.	-0.068 (0.061)	.	0.001 (0.074)	.	-0.160 (0.116)
1993-1997	.	.	0.046 (0.072)	.	0.130 (0.087)	.	-0.047 (0.130)
unemployment growth	-0.003 (0.231)	0.073 (0.218)	-0.154 (0.270)	0.201 (0.282)	0.038 (0.334)	-0.307 (0.409)	-0.430 (0.476)

Note: All models include controls for age, education, industry, firm size, worker status and changes in the national unemployment rate. See Appendix Tables 1-3 for complete set of results. Self-employed, trainees, agriculture (including forestry and fisheries), non-profit organizations and all observations with missing values are excluded. Numbers in parentheses are asymptotic standard errors. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (one-sided tests).

[a] The included age range is 16-64. The dependent variable is a dummy variable equaling 1 if job duration is less than or equal to one year.

[b] The included age range is 35-64. The dependent variable is a dummy variable equaling 1 if job duration is greater than or equal to ten years.

[c] The included age range is 45-64. The dependent variable is a dummy variable equaling 1 if job duration is greater than or equal to twenty years.

Source: Own calculations based on the GSOEP 1984-1997. Weighted using the GSOEP sample weights.

Table 3 - Frequencies in the Duration Analysis

Variable		All new jobs	Leaving unemployment	Leaving schooling or training
Observations	irrespective of missing values for covariates	12564	3020	1726
Failure		8911 (70.9%)	2183 (72.3%)	1174 (68.0%)
Observations	without missing values for covariates	4950	936	660
Failure		1813 (44.7%)	470 (50.2%)	308 (46.7%)
Time	1984-1989	1727 (42.5%)	473 (50.5%)	298 (45.2%)
	1990-1992	1111 (27.4%)	176 (18.8%)	160 (24.2%)
	1993-1997	1221 (30.1%)	287 (30.7%)	202 (30.6%)
Sex	Female	2072 (51%)	355 (37.9%)	298 (45.2%)
	Male	1981 (49%)	581 (62.1%)	362 (54.8%)
Age	< 20	245 (6%)	42 (4.5%)	136 (20.6%)
	20-25	1375 (33.9%)	315 (33.7%)	403 (60.1%)
	> 25 (max 55)	2439 (60.1%)	579 (61.9%)	121 (18.3%)
Education	Compulsory Education	1167 (28.8%)	362 (32.3%)	81 (12.3%)
	Training	2521 (62.1%)	562 (60.0%)	487 (73.8%)
	University	371 (9.1%)	72 (7.7%)	92 (13.9%)
Foreigner		1103 (27.2%)	325 (34.7%)	168 (25.5%)
Parttime		874 (21.5%)	89 (9.5%)	29 (4.4%)
Status	Less skilled blue collar	1203 (29.6%)	374 (40%)	191 (28.9%)
	Skilled blue collar	813 (20.0%)	197 (21%)	266 (40.3%)
	Less skilled white collar	1636 (40.3%)	291 (31.1%)	304 (46.1%)
	Skilled white collar	407 (10.0%)	74 (7.9%)	90 (13.6%)
Firm size	small	1188 (29.3%)	291 (31.1%)	186 (28.2%)
	medium	1218 (30.0%)	313 (33.4%)	165 (25.0%)
	large	1653 (40.7%)	332 (35.5%)	309 (46.8%)

Table 3 continued				
Variable		**All new jobs**	**Leaving unemployment**	**Leaving schooling or training**
Industry	Energy, water, mining	53 (1.3%)	11 (1.2%)	9 (1.4%)
	Manufacturing	1506 (37.1%)	372 (39.7%)	275 (41.7%)
	Construction	386 (9.5%)	135 (14.4%)	49 (7.4%)
	Trade	612 (15.1%)	137 (14.6%)	76 (11.5%)
	Transport	199 (4.9%)	42 (4.5%)	24 (3.6%)
	Credit and insurance	142 (3.5%)	17 (1.8%)	28 (4.2%)
	Services	949 (23.4%)	185 (19.8%)	170 (25.8%)
	State and social insurance	212 (5.2%)	37 (4.0%)	29 (4.4%)
Origin	Full-time employment	940 (23.2%)	-	10 (1.5%)
	Education/training/university	755 (18.6%)	-	487 (73.8%)
	Unemployment	936 (23.1%)	936 (100%)	105 (15.9%)
	Part-time /marginal employment	412 (10.2%)	-	9 (1.4%)
	Non-participation / other	1014 (25.0%)	-	49 (7.4%)
Destination	Full-time employment	643 (35.5%)	178 (36.6%)	161 (50.5%)
	Education/training/university	89 (4.9%)	11 (2.3%)	18 (5.6%)
	Unemployment	564 (31.1%)	229 (47.1%)	61 (19.1%)
	Part-time /marginal employment	99 (5.5%)	12 (2.5%)	6 (1.9%)
	Non-participation / other	418 (23.1%)	56 (11.5%)	73 (22.9%)
Reason	Involuntary	685 (16.9%)	247 (26.4%)	88 (13.3%)
	Voluntary	949 (23.4%)	206 (22%)	173 (26.2%)
Unemployment experience		1425 (35.1%)	641 (68.5%)	101 (15.3%)

Table 4 - The Risk of Job Termination for All New Jobs Between 1984-1997:
Hazard Ratios from Cox Proportional Hazard Models

	All Spells I		Without missings II		All Spells III		Without missings IV	
	α	t-statistic	α	t-statistic	α	t-statistic	α	t-statistic
1990-1992[a]	1.086*	(2.319)	1.175*	(2.553)	1.047	(1.196)	1.077	(0.985)
1993-1997[a]	1.181**	(4.451)	1.413**	(4.017)	1.096*	(2.228)	1.011	(0.101)
Parttime	1.350**	(9.422)	1.137	(1.364)	1.375**	(9.938)	1.176[+]	(1.741)
Parttime * (1990-1992)	0.981	(-0.354)	1.004	(0.030)	0.997	(-0.049	1.040	(0.310)
Parttime * (1993-1997)	0.829**	(-3.207)	0.656**	(-2.594)	0.866*	(-2.429)	0.780	(-1.517)
Layoff	.	.	.	.	1.319**	(6.616)	3.155**	(15.542)
Layoff * (1990-1992)	.	.	.	.	1.204**	(2.699)	1.261*	(2.019)
Layoff * (1993-1997)	.	.	.	.	1.337**	(3.893)	2.872**	(7.718)
Unemployment rate	0.875**	(-9.785)	0.805**	(-7.961)	0.886**	(-8.807)	0.830**	(-6.784)
Unemployment experience dummy	1.025	(1.055)	1.246**	(4.373)	0.992	(-0.361)	1.107*	(1.995)
Age	0.988**	(-10.408)	0.976**	(-8.427)	0.988**	(-10.977)	0.973**	(-9.635)
Foreign	0.818**	(-7.666)	0.765**	(-4.315)	0.816**	(-7.744)	0.743**	(-4.743)
Female	0.887**	(-4.885)	1.056	(0.898)	0.902**	(-4.189)	1.097	(1.553)
Training[b]	0.760**	(-9.945)	0.981	(-0.312)	0.762**	(-9.825)	1.006	(0.100)
University[b]	0.891[+]	(-1.885)	1.370**	(2.746)	0.903[+]	(-1.678)	1.384*	(2.816)
Skilled blue collar[c]	0.822**	(-4.225)	0.743**	(-3.887)	0.835**	(-3.901)	0.800**	(-2.951)
Unskilled white collar[c]	0.844**	(-4.373)	0.700**	(-5.144)	0.850**	(-4.184)	0.783**	(-3.500)
Skilled white collar[c]	0.641**	(-6.407)	0.497**	(-5.927)	0.642**	(-6.359)	0.529**	(-5.345)
Status missing[c]	3.914**	(28.091)	.	.	3.915**	(28.060)	.	.
Small firm < 20[d]	1.195**	(4.700)	1.222**	(3.294)	1.186**	(4.491)	1.152*	(2.329)
Large firm > 200[d]	0.802**	(-5.783)	0.729**	(-5.270)	0.803**	(-5.740)	0.738**	(-5.053)
Firm size missing[d]	1.610**	(7.615)	.	.	1.601**	(7.528)	.	.
Observations	12564		4059		12564		4059	
Failures	8911		1813		8911		1813	
Log Likelihood	-73867		-13715		-73778		-13372	
LR-Chi^2	7070**		378**		7245**		1064**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and [+] at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where $\alpha=\exp(\beta)$ and $(\alpha-1) \times 100$ = percentage effect of a covariate on the job separation risk.

Table 5 - The Risk of Job Termination for Different Groups 1984-1997:
Hazard Ratios from Cox Proportional Hazard Models

		1990-1992	**1993-1997**	**Part-time**	**Part-time *1990-1992**	**Part-time *1993-1997**
I	All new job spells	1.086* (2.319)	1.181** (4.451)	1.350** (9.422)	0.981 (-0.354)	0.829** (-3.207)
II	Workers switching from a full-time job	0.915 (-1.027)	0.814+ (-1.810)	1.365** (3.317)	1.315* (1.971)	1.491** (2.685)
III	Entering from unemployment	1.101 (1.303)	1.189** (2.750)	1.281** (2.673)	1.015 (0.099)	0.872 (-0.923)
IV	Entering from non-participation	1.104 (1.313)	1.250** (2.732)	1.116+ (1.779)	0.950 (-0.528)	0.740** (-2.716)
V	Males	1.005 (0.101)	1.104* (1.984)	1.698** (9.458)	1.055 (0.592)	0.840+ (-1.788)
VI	Females	1.224** (3.817)	1.289** (4.469)	1.261** (5.745)	0.877+ (-1.875)	0.768** (-3.401)
VII	Males switching from full-time job	0.900 (-0.945)	0.675* (-2.476)	2.335** (5.493)	1.168 (0.660)	1.369 (1.327)
VIII	Females switching from full-time job	0.865 (-1.053)	1.022 (0.128)	1.023 (0.181)	1.477* (2.039)	1.299 (1.267)
IX	Males entering from unemployment	1.098 (1.019)	1.209* (2.527)	2.495** (5.340)	0.535* (-2.186)	0.592+ (-1.848)
X	Females entering from unemployment	1.100 (0.752)	1.167 (1.316)	1.121 (1.001)	1.129 (0.600)	0.947 (-0.278)
XI	Males entering from non-participation	1.099 (0.808)	1.154 (1.137)	1.715** (2.965)	1.420 (1.249)	0.800 (-0.793)
XII	Females entering from non-participation	1.196+ (1.748)	1.310* (2.449)	1.065 (0.919)	0.856 (-1.304)	0.687** (-2.777)
XIII	Entering from Education or Training	1.073 (0.821)	1.206+ (1.952)	1.836** (5.165)	0.951 (-0.269)	0.762 (-1.456)

Note: Control group: Jobs starting between 1984-1989. One * indicates significance at the 5% significance level, two ** at the 1% significance level and + at the 10% significance level (two-sided tests). For the complete set of results see Appendix Tables A4-A8.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where $\alpha=\exp(\beta)$ and $(\alpha-1) \times 100$ = percentage effect of a covariate on the job separation risk.

Table 6 - The Risk of Job Termination for Different Groups 1984-1997:
Hazard Ratios from Cox Proportional Hazard Models

		1990-1992	1993-1997	Part-time	Part-time *1990-1992	Part-time *1993-1997	Displaced	Displaced *1990-1992	Displaced *1993-1997
I	All new job spells	1.047 (1.196)	1.096* (2.228)	1.375** (9.938)	0.997 (-0.049)	0.866* (-2.429)	1.319** (6.616)	1.204** (2.699)	1.337** (3.893)
II	Workers switching from a full-time job	0.829* (-2.022)	0.717** (-2.685)	1.380** (3.426)	1.417* (2.485)	1.624** (3.204)	1.526** (3.429)	1.802** (3.171)	1.373** (1.611)
III	Entering from unemployment	1.070 (0.809)	1.011 (0.151)	1.294** (2.784)	1.035 (0.221)	0.937 (-0.433)	1.272** (3.629)	1.115 (0.861)	1.567** (3.824)
IV	Entering from non-participation	1.025 (0.312)	1.185^{+} (1.951)	1.129* (1.961)	0.994 (-0.059)	0.771* (-2.313)	1.157 (1.621)	1.428* (2.546)	1.476* (2.115)
V	Males	0.966 (-0.662)	1.026 (0.458)	1.759** (10.010)	1.057 (0.613)	0.861 (-1.517)	1.337** (5.121)	1.181^{+} (1.778)	1.277* (2.450)
VI	Females	1.182** (3.012)	1.195** (2.932)	1.278** (6.072)	0.889^{+} (-1.674)	0.803** (-2.805)	1.279** (3.927)	1.280* (2.410)	1.473** (3.423)
VII	Entering from Education or Training	1.052 (0.563)	1.160 (1.466)	1.861** (5.272)	0.963 (-0.202)	0.748 (-1.552)	1.518** (3.412)	1.175 (0.803)	1.229 (0.973)

Note: Control group: Jobs starting between 1984-1989. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (two-sided tests). For the complete set of results see Appendix Tables A4-A8.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where α=exp(β) and (α-1) $\times$ 100 = percentage effect of a covariate on the job separation risk.

Appendix Tables

Appendix Table A1 - Logit Analyis: Probability of Elapsed Tenure ≤ 1 Year

	All I	All II	All III	Men IV	 V	Women VI	 VII
Constant	-6.937** (0.841)	-2.521** (0.121)	-2.544** (0.121)	-6.723** (1.090)	-2.464** (0.200)	-6.911** (0.992)	-2.562** (0.154)
Time trend	0.051** (0.009)	.	.	0.049** (0.013)	.	0.050** (0.011)	.
1990-1997	.	0.262** (0.055)	.	.	.	.	.
1990-1992	.	.	0.171** (0.062)	.	0.079 (0.086)	.	0.262** (0.092)
1993-1997	.	.	0.397** (0.087)	.	0.462** (0.116)	.	0.330** (0.109)
Δ unemployment rate	-0.723** (0.238)	-0.322 (0.236)	-.778** (0.274)	-1.002** (0.337)	-1.346** (0.383)	-0.336 (0.338)	-0.139 (0.391)
High School (*Abitur*)	0.490** (0.174)	0.510** (0.174)	-0.778** (0.274)	0.108 (0.318)	0.111 (0.320)	0.505^{+} (0.261)	0.519* (0.259)
Training	0.059 (0.087)	0.056 (0.087)	0.499** (0.174)	0.034 (0.151)	0.032 (0.152)	0.049 (0.099)	0.047 (0.099)
University	0.744** (0.123)	0.734** (0.123)	0.057 (0.087)	0.696** (0.213)	0.684** (0.215)	0.697** (0.169)	0.699** (0.170)
Parttime	0.267** (0.100)	0.279** (0.102)	0.735** (0.123)	1.163* (0.467)	1.185* (0.476)	0.211* (0.083)	0.219** (0.083)
Firm size < 20	0.250** (0.066)	0.252** (0.066)	0.253** (0.066)	0.247* (0.099)	0.256* (0.099)	0.304** (0.088)	0.304** (0.088)
Firm size > 200	-0.318** (0.068)	-0.318** (0.068)	-0.315** (0.068)	-0.536** (0.086)	-0.531** (0.087)	-0.144^{+} (0.086)	-0.143^{+} (0.086)
Skilled blue collar	-0.805** (0.084)	-0.793** (0.084)	-0.798** (0.084)	-0.822** (0.096)	-0.818** (0.096)	-0.789** (0.188)	-0.773** (0.189)
Unskilled white collar	-0.345** (0.074)	-0.355** (0.075)	-0.352** (0.075)	-0.193^{+} (0.116)	-0.195^{+} (0.117)	-0.353** (0.103)	-0.357** (0.104)
Skilled white collar	-0.646** (0.094)	-0.611** (0.091)	-0.625** (0.092)	-0.741** (0.139)	-0.728** (0.137)	-0.537** (0.116)	-0.508** (0.116)
Observations		49103		29481		19622	
Log-Likelihood	-13959	-13972	-13966	-7732	-7727	-6031	-6040
Wald-Chi^2	1601**	1605**	1616**	1053**	1076**	761**	750**

Note: The included age range is 16-64. The dependent variable is a dummy variable equaling 1 if job duration is less than or equal to one year. Self-employed, trainees, agriculture (including forestry and fisheries), non-profit organizations and all observations with missing values are excluded. All models include seven industry dummies for eight one-digit industries and eight age dummies for nine categories: 16-20, 21-25,26-30 and so forth. Numbers in parentheses are asymptotic standard errors. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (one-sided tests). Reference group: unskilled blue collar with compulsory education (*Hauptschule or Realschule*) in medium sized firm (20-200 employees).
Source: Own calculations based on the GSOEP 1984-1997. Weighted using the GSOEP sample weights.

Appendix Table A2 - Logit Analyis: Probability of Elapsed Tenure ≥ 10 Year

	All I	All II	All III	Men IV	 V	Women VI	 VII
Constant	1.590** (0.537)	-0.177** (0.066)	-0.165* (0.066)	1.172+ (0.661)	0.006 (0.095)	2.162** (0.799)	-0.354** (0.098)
Time trend	-0.022** (0.006)	.	.	-0.017* (0.007)	.	-0.030** (0.009)	.
1990-1997	.	-0.102* (0.041)	.	.	.	.	.
1990-1992	.	.	-0.044 (0.046)	.	0.016 (0.061)	.	-0.117 (0.074)
1993-1997	.	.	-0.182** (0.059)	.	-0.154* (0.075)	.	-0.230** (0.088)
Δ unemployment rate	0.119 (0.176)	-0.061 (0.168)	0.214 (0.206)	0.160 (0.236)	0.360 (0.281)	0.080 (0.269)	0.096 (0.310)
High School (*Abitur*)	-0.542* (0.255)	-0.517* (0.263)	0.214 (0.206)	-1.033** (0.310)	-1.014** (0.318)	0.235 (0.605)	0.263 (0.611)
Training	-0.177** (0.054)	-0.173** (0.054)	-0.509+ (0.262)	-0.251** (0.084)	-0.236** (0.086)	-0.160* (0.070)	-0.162* (0.070)
University	-0.643** (0.087)	-0.625** (0.088)	-0.172** (0.054)	-0.800** (0.119)	-0.770** (0.121)	-0.476** (0.172)	-0.475** (0.172)
Parttime	-0.544** (0.059)	-0.550** (0.060)	-0.624** (0.088)	-1.784** (0.594)	-1.779** (0.604)	-0.378** (0.057)	-0.378** (0.057)
Firm size < 20	-0.300** (0.055)	-0.300** (0.055)	-0.300** (0.055)	-0.351** (0.077)	-0.350** (0.078)	-0.339** (0.080)	-0.340** (0.080)
Firm size > 200	0.677** (0.045)	0.673** (0.046)	0.673** (0.046)	0.953** (0.058)	0.947** (0.058)	0.272** (0.069)	0.271** (0.069)
Skilled blue collar	0.656** (0.060)	0.654** (0.061)	0.654** (0.061)	0.538** (0.073)	0.529** (0.074)	0.652** (0.157)	0.643** (0.157)
Unskilled white collar	0.361** (0.059)	0.370** (0.060)	0.367** (0.060)	0.367** (0.091)	0.377** (0.092)	0.426** (0.082)	0.430** (0.082)
Skilled white collar	0.443** (0.062)	0.429** (0.062)	0.435** (0.062)	0.254** (0.083)	0.241** (0.084)	0.702** (0.094)	0.690** (0.094)
Observations		29443		18253		11190	
Log-Likelihood	-17499	-17455	-17458	-10361	-10293	-6905	-6909
Wald-Chi^2	1895**	1941**	1938**	1069**	1121**	781**	786**

Note: The included age range is 35-64. The dependent variable is a dummy variable equaling 1 if job duration is greater than or equal to ten years. Self-employed, trainees, agriculture (including forestry and fisheries), non-profit organizations and all observations with missing values are excluded. All models include seven industry dummies for eight one-digit industries and five age dummies for six catagories 35-39, 40-44, 45-49 and so forth). Numbers in parentheses are asymptotic standard errors. One * indicates significance at the 5% significance level, two ** at the 1% significance level and + at the 10% significance level (one-sided tests). Reference group: unskilled blue collar with compulsory education (*Hauptschule or Realschule*) in medium sized firm (20-200 employees).
Source: Own calculations based on the GSOEP 1984-1997. Weighted using the GSOEP sample weights.

Appendix Table A3 - Logit Analyis: Probability of Elapsed Tenure ≥ 20 Year

	All I	All II	All III	Men IV	V	Women VI	VII
Constant	-2.046**	-1.756**	-1.770**	-2.687**	-1.738**	-1.696	-1.884**
	(0.647)	(0.084)	(0.084)	(0.760)	(0.104)	(1.221)	(0.152)
Time trend	0.003	.	.	0.011	.	-0.003	.
	(0.007)			(0.008)		(0.014)	
1990-1997	.	-0.021	.	.	.	.	.
		(0.052)					
1990-1992	.	.	-0.068	.	0.001	.	-0.160
			(0.061)		(0.074)		(0.116)
1993-1997	.	.	0.046	.	0.130	.	-0.047
			(0.072)		(0.087)		(0.130)
Δ unemployment rate	-0.003	0.073	-0.154	0.201	0.038	-0.307	-0.430
	(0.231)	(0.218)	(0.270)	(0.282)	(0.334)	(0.409)	(0.476)
High School (*Abitur*)	0.353	0.350	0.341	0.461	0.456	0.127	0.117
	(0.367)	(0.364)	(0.367)	(0.447)	(0.448)	(0.692)	(0.693)
Training	0.056	0.055	0.055	0.028	0.027	-0.057	-0.056
	(0.063)	(0.063)	(0.063)	(0.083)	(0.083)	(0.102)	(0.102)
University	-0.263*	-0.270*	-0.269*	-0.304*	-0.307*	-0.410^{+}	-0.420^{+}
	(0.113)	(0.113)	(0.113)	(0.137)	(0.138)	(0.226)	(0.226)
Parttime	-0.746**	-0.742**	-0.745**	-1.624**	-1.624**	-0.461**	-0.462**
	(0.081)	(0.081)	(0.081)	(0.513)	(0.514)	(0.091)	(0.091)
Firm size < 20	0.222**	0.223**	0.223**	0.230*	0.230*	0.253^{+}	0.261^{+}
	(0.079)	(0.079)	(0.079)	(0.099)	(0.099)	(0.137)	(0.137)
Firm size > 200	0.814**	0.814**	0.815**	0.817**	0.818**	0.788**	0.791**
	(0.058)	(0.058)	(0.058)	(0.068)	(0.068)	(0.110)	(0.110)
Skilled blue collar	0.804**	0.805**	0.806**	0.745**	0.747**	0.844**	0.847**
	(0.073)	(0.073)	(0.073)	(0.082)	(0.082)	(0.230)	(0.232)
Unskilled white collar	0.441**	0.435**	0.437**	0.565**	0.565**	0.426**	0.413**
	(0.076)	(0.076)	(0.076)	(0.102)	(0.102)	(0.117)	(0.117)
Skilled white collar	0.929**	0.938**	0.934**	0.954**	0.956**	0.896**	0.913**
	(0.077)	(0.077)	(0.077)	(0.093)	(0.093)	(0.143)	(0.142)
Observations		17103		10894		6202	
Log-Likelihood	-9987	-9988	-9987	-6890	-6890	-3083	-3082
Wald-Chi^2	1895**	1941**	1938**	1069**	1121**	781**	786**

Note: The included age range is 45-64. The dependent variable is a dummy variable equaling 1 if job duration is greater than or equal to twenty years. Self-employed, trainees, agriculture (including forestry and fisheries), non-profit organizations and all observations with missing values are excluded. All models include seven industry dummies for eight one-digit industries and three age dummies for four categories: 45-49, 50-54, 55-59, 60-64. Numbers in parentheses are asymptotic standard errors. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (one-sided tests). Reference group: unskilled blue collar with compulsory education (*Hauptschule or Realschule*) in medium sized firm (20-200 employees).
Source: Own calculations based on the GSOEP 1984-1997. Weighted using the GSOEP sample weights.

Appendix Table A4 - The Risk of Job Termination for Workers Switching from Full-Time Employment: Hazard Ratios from Cox Proportional Hazard Models

	All Job Switchers				All Job Switchers			
	I		II		III Males		IV Females	
	α	t-statistic	α	t-statistic	α	t-statistic	α	t-statistic
1990-1992[a]	0.915	(-1.027)	0.829*	(-2.022)	0.900	(-0.945)	0.865	(-1.053)
1993-1997[a]	0.814[+]	(-1.810)	0.717**	(-2.685)	0.675*	(-2.476)	1.022	(0.128)
Parttime	1.365**	(3.317)	1.380**	(3.426)	2.335**	(5.493)	1.023	(0.181)
Parttime * (1990-1992)	1.315*	(1.971)	1.417*	(2.485)	1.168	(0.660)	1.477*	(2.039)
Parttime * (1993-1997)	1.491**	(2.685)	1.624**	(3.204)	1.369	(1.327)	1.299	(1.267)
Layoff	.	.	1.526**	(3.429)	.	.	.	.
Layoff * (1990-1992)	.	.	1.802**	(3.171)	.	.	.	.
Layoff * (1993-1997)	.	.	1.373**	(1.611)	.	.	.	.
Unemployment	0.962	(-1.103)	0.977	(-0.671)	0.984	(-0.301)	0.959	(-0.888)
Unemployment experience dummy	1.028	(0.456)	1.013	(0.221)	0.980	(-0.225)	1.119	(1.318)
Age	0.986**	(-4.266)	0.986**	(-4.200)	0.995	(-0.916)	0.985**	(-3.406)
Foreign	0.852*	(-2.284)	0.856*	(-2.201)	0.762**	(-2.721)	0.938	(-0.629)
Female	0.915	(-1.303)	0.940	(-0.908)	.	.	.	.
Training[b]	0.894	(-1.558)	0.923	(-1.119)	0.794*	(-2.220)	1.025	(0.239)
University[b]	0.910	(-0.660)	0.953	(-0.342)	0.587**	(-2.660)	1.283	(1.201)
Skilled blue collar[c]	0.803*	(-1.998)	0.822[+]	(-1.779)	0.751*	(-2.152)	0.942	(-0.260)
Unskilled white collar[c]	0.792*	(-2.435)	0.796*	(-2.379)	0.697*	(-2.176)	0.811*	(-1.670)
Skilled white collar[c]	0.735*	(-2.150)	0.723*	(-2.280)	0.555**	(-3.076)	0.963	(-0.170)
Status missing[c]	3.274**	(7.332)	3.424**	(7.653)	2.155**	(3.123)	4.064**	(6.355)
Small firm < 20[d]	1.196*	(2.104)	1.183*	(1.976)	1.364*	(2.443)	1.070	(0.582)
Large firm > 200[d]	0.734**	(-3.708)	0.750**	(-3.446)	0.763*	(-2.184)	0.670**	(-3.462)
Firm size missing[d]	1.724**	(3.079)	1.742**	(3.143)	2.086**	(2.679)	1.451	(1.560)
Observations	2086		2086		1040		1046	
Failures	1345		1345		627		718	
Log Likelihood	-8794		-8757		-3646		-4177	
LR-Chi^2	952		1026**		580**		453**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and [+] at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where $\alpha=\exp(\beta)$ and $(\alpha-1) \times 100$ = percentage effect of a covariate on the job separation risk.

Appendix Table A5 - The Risk of Job Termination for Workers Entering from Un- or Non participation: Hazard Ratios from Cox Proportional Hazard Models

	Unemployment				Non-participation			
	I		II		III		IV	
	α	t-statistic	α	t-statistic	α	t-statistic	α	t-statistic
1990-1992[a]	1.101	(1.303)	1.070	(0.809)	1.104	(1.313)	1.025	(0.312)
1993-1997[a]	1.189**	(2.750)	1.011	(0.151)	1.250**	(2.732)	1.185$^+$	(1.951)
Parttime	1.281**	(2.673)	1.294**	(2.784)	1.116$^+$	(1.779)	1.129*	(1.961)
Parttime * (1990-1992)	1.015	(0.099)	1.035	(0.221)	0.950	(-0.528)	0.994	(-0.059)
Parttime * (1993-1997)	0.872	(-0.923)	0.937	(-0.433)	0.740**	(-2.716)	0.771*	(-2.313)
Layoff	.	.	1.272**	(3.629)	.	.	1.157	(1.621)
Layoff * (1990-1992)	.	.	1.115	(0.861)	.	.	1.428*	(2.546)
Layoff * (1993-1997)	.	.	1.567**	(3.824)	.	.	1.476*	(2.115)
Unemployment rate	0.864**	(-4.857)	0.883**	(-4.140)	0.880**	(-5.009)	0.889**	(-4.590)
Unemployment experience dummy	1.090$^+$	(1.668)	1.082	(1.509)	1.065	(1.356)	1.045	(0.943)
Age	0.991**	(-4.275)	0.990**	(-4.672)	0.991**	(-4.084)	0.991**	(-4.179)
Foreign	0.921$^+$	(-1.682)	0.925	(-1.582)	0.795**	(-4.557)	0.796**	(-4.507)
Female	0.992	(-0.159)	1.005	(0.106)	0.958	(-0.747)	0.963	(-0.662)
Training[b]	0.742**	(-5.621)	0.743**	(-5.613)	0.750**	(-5.720)	0.749**	(-5.756)
University[b]	0.924	(-0.594)	0.950	(-0.383)	0.921	(-0.697)	0.936	(-0.553)
Skilled blue collar[c]	0.905	(-1.235)	0.924	(-0.979)	0.773**	(-2.641)	0.788*	(-2.448)
Unskilled white collar[c]	0.766**	(-3.175)	0.781**	(-2.944)	0.913	(-1.342)	0.923	(-1.178)
Skilled white collar[c]	0.733*	(-2.099)	0.727*	(-2.140)	0.677**	(-2.614)	0.683*	(-2.545)
Status missing[c]	3.572**	(13.121)	3.505**	(12.856)	4.134**	(17.178)	4.197**	(17.298)
Small firm < 20[d]	1.141$^+$	(1.802)	1.135$^+$	(1.725)	1.262**	(3.083)	1.253*	(2.989)
Large firm > 200[d]	0.794**	(-2.968)	0.791**	(-3.015)	0.874$^+$	(-1.787)	0.866$^+$	(-1.908)
Firm size missing[d]	1.318*	(2.248)	1.309*	(2.185)	2.047**	(6.088)	2.012**	(5.932)
Observations	3020		3020		3445		3445	
Failures	2183		2183		2466		2466	
Log Likelihood	-15059		-15025		-17298		-17282	
LR-Chi^2	1476**		1542**		1855**		1887**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^+$ at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.

Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where α=exp(β) and (α-1) × 100 = percentage effect of a covariate on the job separation risk.

Appendix Table A6 - The Risk of Job Termination for Men and Women: Hazard Ratios from Cox Proportional Hazard Models

	All Job Spells				All Job Spells			
	I Males		II Males		III Females		IV Females	
	α	t-statistic	α	t-statistic	α	t-statistic	α	t-statistic
1990-1992[a]	1.005	(0.101)	0.966	(-0.662)	1.224**	(3.817)	1.182**	(3.012)
1993-1997[a]	1.104*	(1.984)	1.026	(0.458)	1.289**	(4.469)	1.195**	(2.932)
Parttime	1.698**	(9.458)	1.759**	(10.010)	1.261**	(5.745)	1.278**	(6.072)
Parttime * (1990-1992)	1.055	(0.592)	1.057	(0.613)	0.877^{+}	(-1.875)	0.889^{+}	(-1.674)
Parttime * (1993-1997)	0.840^{+}	(-1.788)	0.861	(-1.517)	0.768**	(-3.401)	0.803**	(-2.805)
Layoff	.	.	1.337**	(5.121)	.	.	1.279**	(3.927)
Layoff * (1990-1992)	.	.	1.181^{+}	(1.778)	.	.	1.280*	(2.410)
Layoff * (1993-1997)	.	.	1.277*	(2.450)	.	.	1.473**	(3.423)
Unemployment	0.876**	(-6.400)	0.888**	(-5.720)	0.879**	(-7.053)	0.889**	(-6.386)
Unemployment experience dummy	0.977	(-0.670)	0.941^{+}	(-1.735)	1.064*	(1.970)	1.035	(1.074)
Age	0.993**	(-4.087)	0.992**	(-4.637)	0.987**	(-8.490)	0.987**	(-8.660)
Foreign	0.831**	(-4.906)	0.833**	(-4.837)	0.826**	(-5.206)	0.823**	(-5.322)
Female	.	.	0.784**	(-5.835)	.		0.780**	(-6.630)
Training[b]	0.774**	(-6.163)	0.806*	(-2.448)	0.782**	(-6.551)	1.026	(0.301)
University[b]	0.794**	(-2.621)	0.811**	(-3.676)	1.015	(0.180)	0.941	(-0.651)
Skilled blue collar[c]	0.802**	(-3.869)	0.768**	(-3.820)	0.937	(-0.699)	0.896^{*}	(-2.218
Unskilled white collar[c]	0.761**	(-3.934)	0.545**	(-6.362)	0.885*	(-2.492)	0.828^{+}	(-1.833)
Skilled white collar[c]	0.544**	(-6.379)	3.492**	(17.246)	0.830^{+}	(-1.813)	4.137**	(21.622)
Status missing[c]	3.491**	(17.228)	1.235**	(3.764)	4.126**	(21.644)	1.155**	(2.788)
Small firm < 20[d]	1.251**	(3.982)	0.747**	(-5.183)	1.160**	(2.864)	0.859**	(-2.896)
Large firm > 200[d]	0.748**	(-5.172)	1.619**	(5.163)	0.858**	(-2.921)	1.633**	(5.786)
Firm size missing[d]	1.644**	(5.321)			1.623**	(5.725)		
Observations	5903		5903		6661		6661	
Failures	4139		4139		4772		4772	
Log Likelihood	-31151		-31102		-36513		-36471	
LR-Chi^2	3667**		3761**		3480**		3565**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.

Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where $\alpha=\exp(\beta)$ and $(\alpha-1) \times 100$ = percentage effect of a covariate on the job separation risk.

Appendix Table A7 - The Risk of Job Termination for Men and Women Entering from Unemployment or Non-Participation: Hazard Ratios from Cox Proportional Hazard Models

	Entering from Unemployment				Entering from Non-participation			
	I Males		II Females		I Males		II Females	
	α	t-statistic	α	t-statistic	α	t-statistic	α	t-statistic
1990-1992[a]	1.098	(1.019)	1.100	(0.752)	1.099	(0.808)	1.196+	(1.748)
1993-1997[a]	1.209*	(2.527)	1.167	(1.316)	1.154	(1.137)	1.310*	(2.449)
Parttime	2.495**	(5.340)	1.121	(1.001)	1.715**	(2.965)	1.065	(0.919)
Parttime * (1990-1992)	0.535*	(-2.186)	1.129	(0.600)	1.420	(1.249)	0.856	(-1.304)
Parttime * (1993-1997)	0.592+	(-1.848)	0.947	(-0.278)	0.800	(-0.793)	0.687**	(-2.777)
Layoff	.	.	.	.	.	.	.	.
Layoff * (1990-1992)	.	.	.	.	.	.	.	.
Layoff * (1993-1997)	.	.	.	.	.	.	.	.
Unemployment	0.872**	(-3.526)	0.848**	(-3.451)	0.924	(-1.551)	0.874**	(-4.564)
Unemployment experience dummy	1.061	(0.857)	1.096	(1.141)	1.032	(0.318)	1.061	(1.103)
Age	0.993*	(-2.441)	0.986**	(-3.614)	0.997	(-0.902)	0.990**	(-4.012)
Foreign	0.921	(-1.353)	0.946	(-0.654)	0.808*	(-2.195)	0.751**	(-4.760)
Female								
Training[b]	0.765**	(-4.096)	0.708**	(-3.716)	0.674**	(-3.633)	0.781**	(-4.336)
University[b]	0.779	(-1.450)	0.982	(-0.087)	0.855	(-0.645)	1.050	(0.359)
Skilled blue collar[c]	0.900	(-1.156)	0.877	(-0.612)	0.736*	(-2.151)	0.865	(-0.905)
Unskilled white collar[c]	0.757*	(-2.185)	0.805+	(-1.756)	0.720+	(-1.838)	0.968	(-0.430)
Skilled white collar[c]	0.732+	(-1.730)	0.847	(-0.641)	0.469**	(-2.901)	0.905	(-0.523)
Status missing[c]	3.348**	(9.993)	3.766**	(7.958)	4.100**	(6.546)	4.104**	(15.700)
Small firm < 20[d]	1.098	(1.024)	1.255+	(1.818)	1.757**	(3.866)	1.121	(1.280)
Large firm > 200[d]	0.718**	(-3.314)	0.941	(-0.485)	0.822	(-1.427)	0.907	(-1.074)
Firm size missing[d]	1.319+	(1.804)	1.381	(1.534)	2.171**	(2.923)	1.972**	(5.075)
Observations	1903		1116		968		2477	
Failures	1384		799		636		1830	
Log Likelihood	-8912		-4699		-3631		-12228	
LR-Chi^2	951**		555**		612**		1278**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and + at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where α=exp(β) and (α-1) × 100 = percentage effect of a covariate on the job separation risk.

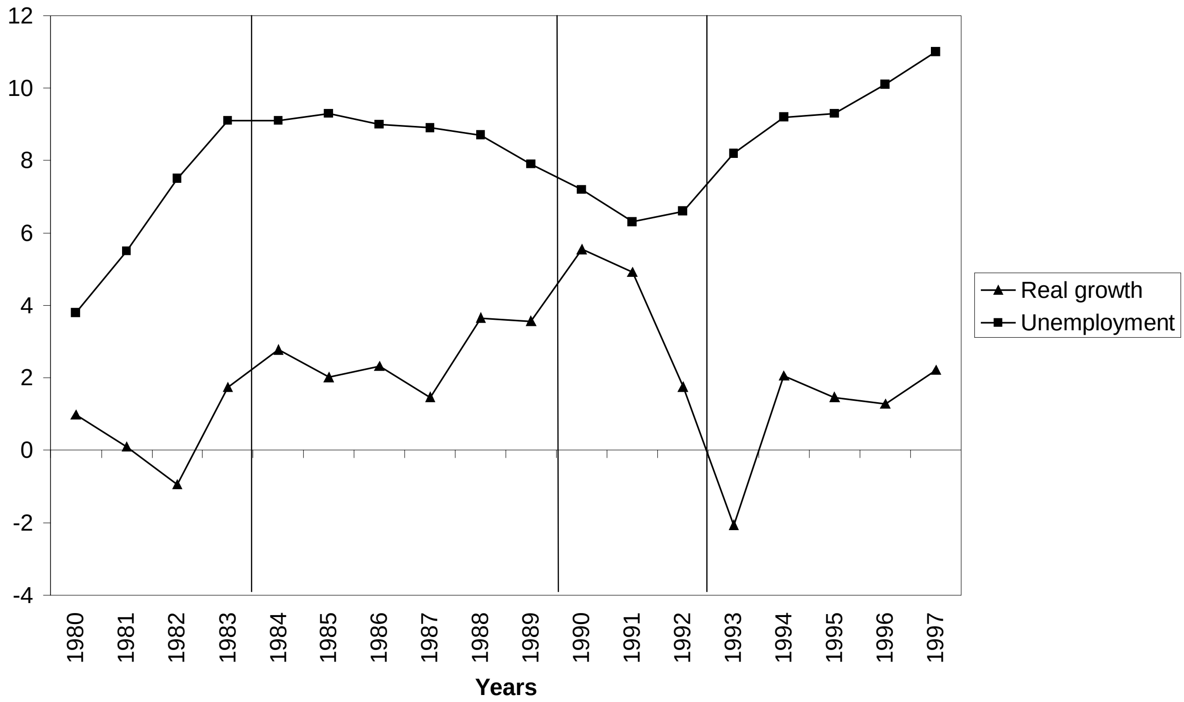
Appendix Table A8 - The Risk of Job Termination for Workers Having Recently Finished Education or Training: Hazard Ratios from Cox Proportional Hazard Models

	All Spells			
	I		II	
	α	t-statistic	α	t-statistic
1990-1992[a]	1.073	(0.821)	1.052	(0.563)
1993-1997[a]	1.206^{+}	(1.952)	1.160	(1.466)
Parttime	1.836**	(5.165)	1.861**	(5.272)
Parttime * (1990-1992)	0.951	(-0.269)	0.963	(-0.202)
Parttime * (1993-1997)	0.762	(-1.456)	0.748	(-1.552)
Layoff	.	.	1.518**	(3.412)
Layoff * (1990-1992)	.	.	1.175	(0.803)
Layoff * (1993-1997)	.	.	1.229	(0.973)
Unemployment rate	0.893**	(-3.029)	0.910*	(-2.504)
Unemployment experience dummy	1.100	(1.137)	1.067	(0.774)
Age	0.965**	(-3.240)	0.967**	(-3.002)
Foreign	0.616**	(-6.288)	0.619**	(-6.207)
Female	0.848*	(-2.477)	0.851*	(-2.410)
Training[b]	0.653**	(-4.941)	0.660**	(-4.811)
University[b]	0.829	(-1.201)	0.847	(-1.059)
Skilled blue collar[c]	0.683**	(-2.812)	0.698**	(-2.665)
Unskilled white collar[c]	0.674**	(-3.103)	0.690**	(-2.911)
Skilled white collar[c]	0.425**	(-4.580)	0.426**	(-4.572)
Status missing[c]	3.752**	(8.179)	3.867**	(8.537)
Small firm < 20[d]	1.361**	(2.895)	1.330**	(2.680)
Large firm > 200[d]	0.868	(-1.372)	0.860	(-1.458)
Firm size missing[d]	1.087	(0.430)	1.059	(0.300)
Observations	1726		1726	
Failures	1174		1174	
Log Likelihood	-7323		-7307	
LR-Chi^2	1886**		1119**	

Note:.[a]: Control group: Jobs starting between 1984-1989. [b]: Control group: compulsory education or Abitur only. [c]: Control group: unskilled blue collar worker. [d]: Control group: firm with 20-200 employees. One * indicates significance at the 5% significance level, two ** at the 1% significance level and $^{+}$ at the 10% significance level (two-sided tests). Controls for seven 1-digit industry groups are included. Unemployment is allowed to vary over time.
Source: Own calculations using the GSOEP. Reported are the hazard ratios or α-coefficients for the covariates, where $\alpha=\exp(\beta)$ and $(\alpha-1) \times 100$ = percentage effect of a covariate on the job separation risk.

Figures

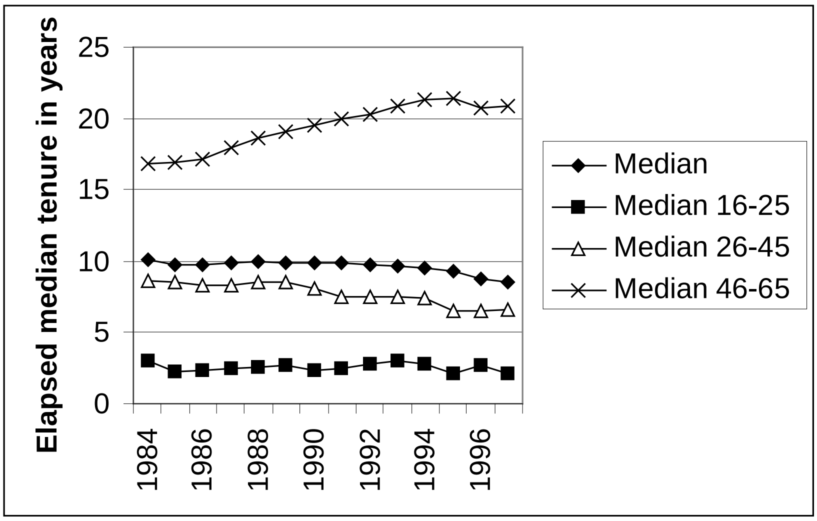
Figure 1 - Unemployment and the Business Cycle in West Germany



Source: Statistisches Jahrbuch für die Bundesrepublik Deutschland and Amtliche Nachrichten der Bundesanstalt für Arbeit.

Figure 2 - The Evolution of Elapsed Tenure by Age Groups

Panel A - Men



Panel B - Women

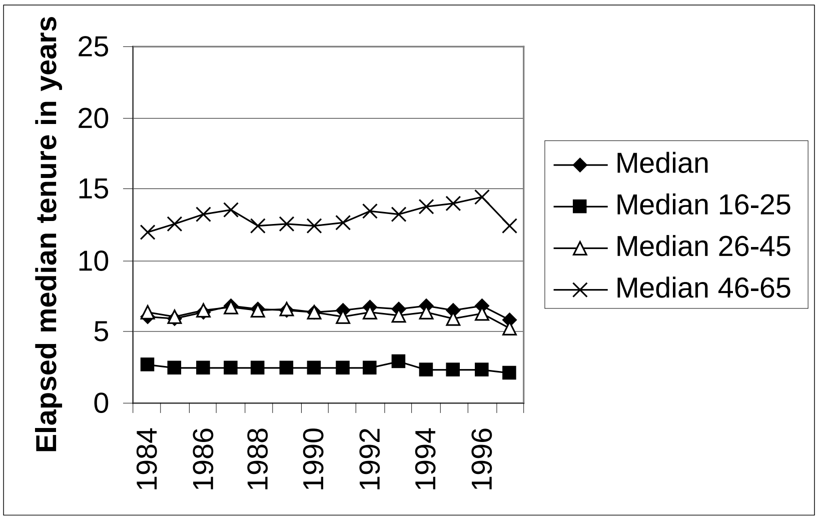
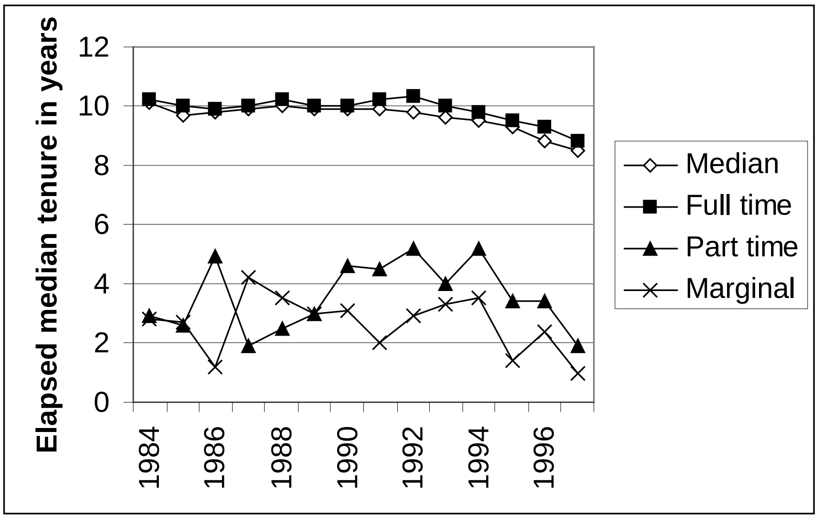
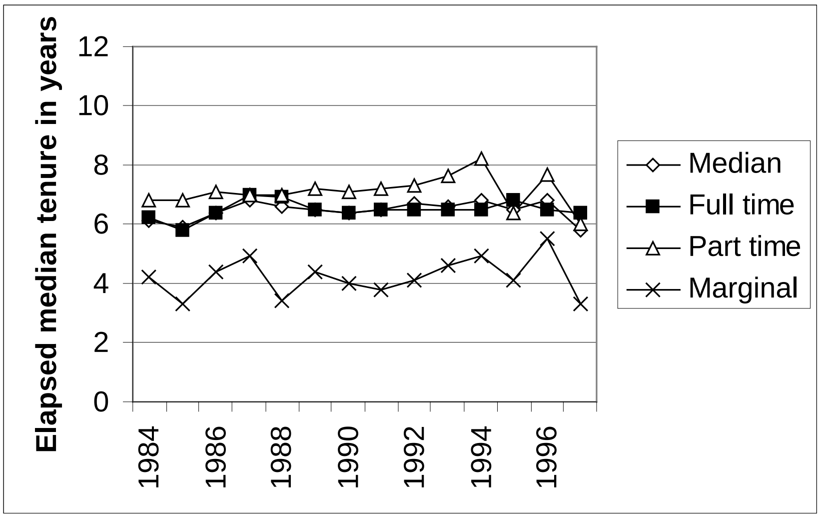


Figure 3 - The Evolution of Elapsed Tenure by Hours Worked

Panel A - Men



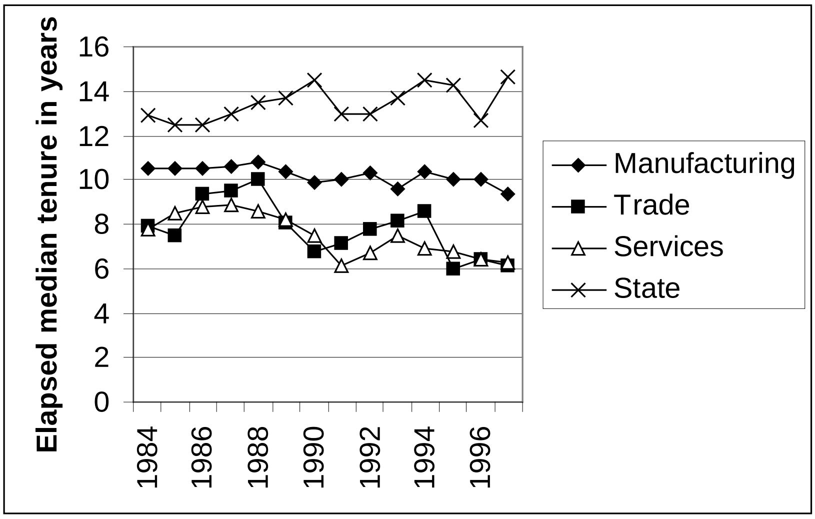
Panle B - Women



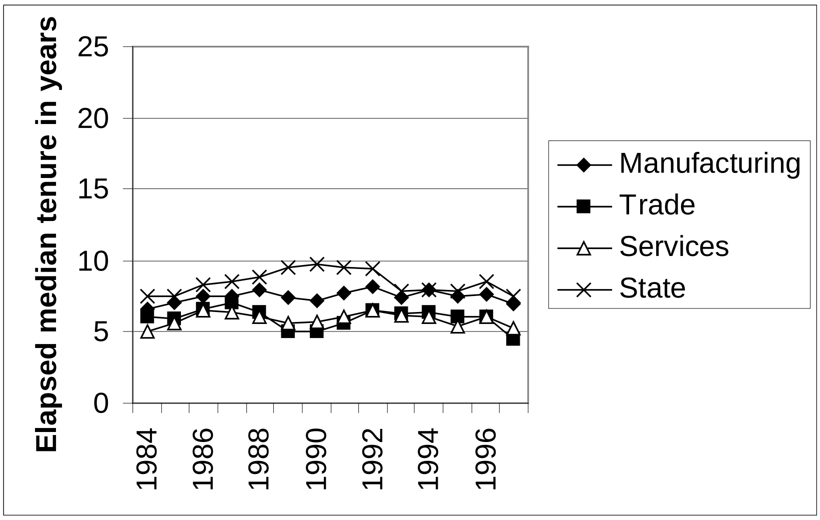
Source: Own calculations based on the GSOEP (samples A and B).

Figure 4 - The Evolution of Elapsed Tenure by Industry

Panel A - Men



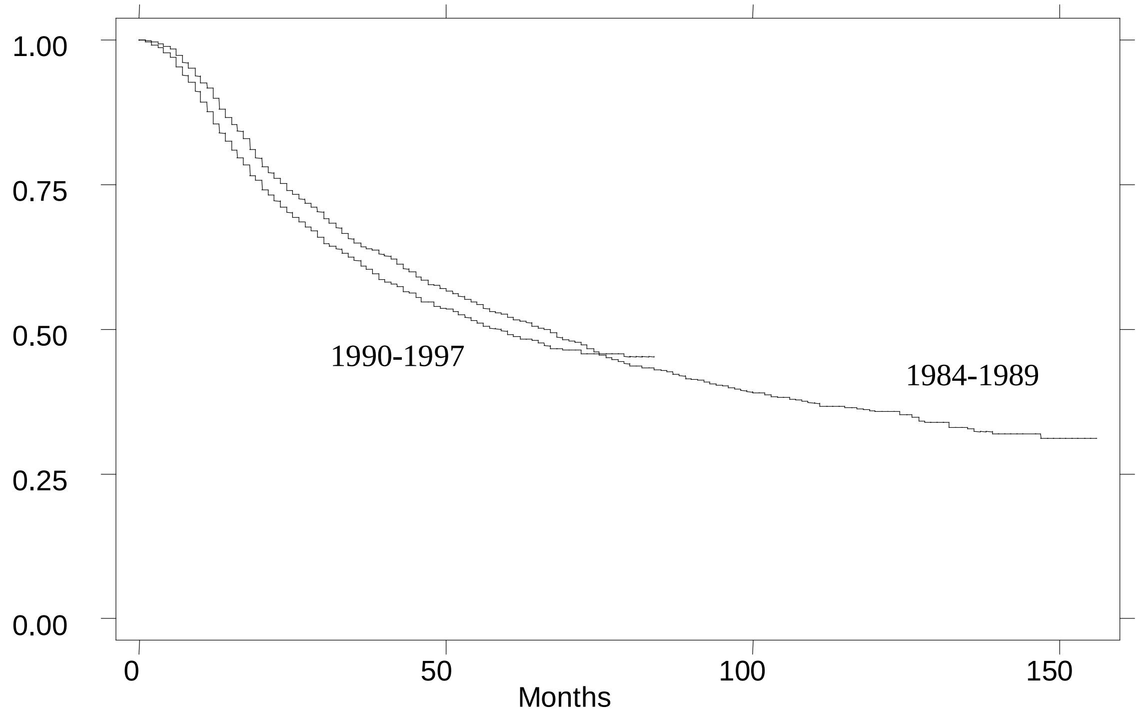
Panel B - Women



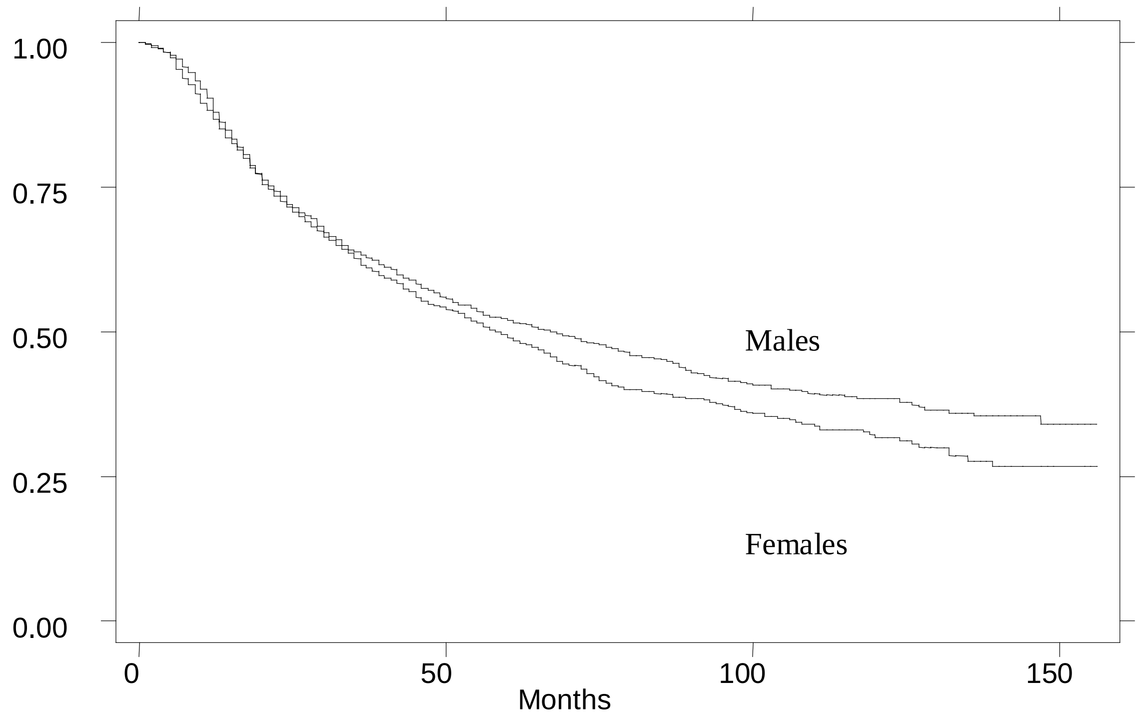
Source: Own calculations based on the GSOEP (samples A and B).

Figure 5 - Kaplan-Meier Survival Estimates of All New Job Spells

Panel A - Two time periods: 1984-1989 and 1990-1997



Panel B - Males and Females



Source: Own calculations based on the GSOEP (samples A and B).