CHANGES IN JOB STABILITY – EVIDENCE FROM LIFETIME JOB HISTORIES*

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We use individual-level panel data spanning over 42 years from the pension records to evaluate changes in job stability in Finland between 1963 and 2004. Compared with previous research on job stability we cover much longer period and for some cohorts observe the entire lifetime job histories. These data allow us to study job stability using standard duration models instead of simply examining changes in elapsed tenure. We find that hazard of job loss increased during the recession years in the early 1990s but has then returned to the level that prevailed in the 1970s. We also demonstrate that fluctuations in the hazard rate together with changes in labor market entry rates have complicated dynamic effects on the tenure distribution, and that analysing changes in job stability based on the elapsed duration of ongoing jobs may be quite misleading. (JEL: J63)

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1. Introduction

A common perception in both Finland and elsewhere in the western countries is that the labor markets are becoming more uncertain. According to this view the risk of job loss is increasing and "lifetime jobs", where the employees spend most of their careers, no longer exist for the younger generations. The Finnish discussion has focused on the effect of temporary jobs, but the question is associated with more general changes in the labour market. The share of manufacturing jobs is decreasing while the share of service sector is increasing which potentially affects the riskiness and hence duration of typical employment relationships. In this paper we document the long-term trends in job stability based on an unusual administrative data source where individual employees and their careers can be followed over four decades.

There is now a relatively large literature on the changes in job stability and on the incidence of lifetime jobs in the economy. A general conclusion from the existing studies is that changes in job stability were small up to the mid 1990s (e.g. Neumark 2000, Burgess and Rees 1996), but during the more recent years job stability has declined (e.g. Farber 2007, Gregg and Wadsworth 2002). While the majority of existing studies are based on US data, there are also several studies describing the trends in job durations in European countries. A cross-country study by OECD (2007) is still probably the most recent source of comparative data across countries. We use data from Finland that could be considered as a typical European country in terms of job stability. According to the Eurostat Labor Force survey, job stability, as measured by average job tenure, was 9.9 years in Finland in 2005. This is only slightly higher than the unweighted EU average of 9.7 years.

A common problem in existing studies of job stability is that genuine panel data, where job spells can be followed over their entire span, is generally lacking. This has prevented modelling job durations using standard methods of survival analysis. Instead, researchers have examined changes in the distribution of elapsed duration of ongoing jobs or used strategies based on inferring job durations from retention rates calculated from short panels or synthetic cohort data.

In this paper we use administrative data from the Finnish pension insurance companies. These data cover a time period starting from 1963, the year following the creation of the earnings related pension system in 1962, up to the recent pension reform in the end of 2004. We have access to a representative sample of fifteen cohorts of Finns who were employed in the private sector at some point during this 42-year interval. The oldest cohort was born in 1905 and the youngest cohort in 1975. The data include the starting and ending dates of all employment spells of the individuals in the sample. This allows us to create a sample of job spells that were ongoing on January 1st, 1963 or started sometime after that date and follow these spells until they end – even if the jobs last for several decades. For some cohorts this implies that we can observe all employment spells during the entire career and directly observe the completed durations of all employment spells. The key limitation of the data is that it only includes private sector employees covered by the Employees Pension Scheme (TEL). However, our analysis still covers most private sector jobs: for example, in 2004 about 90 percent of private sector employees paid TEL contributions.

We start by taking snapshots of data at regular intervals and by computing the distribution of the elapsed duration of ongoing jobs at various points in time. Since we are using administrative data, many problems related to recall errors and consistency of measurement over time can be avoided. Also problems due to non-response or panel attrition do not arise with these data. To verify that administrative data is consistent with typical survey data, we compare the results from administrative records with the data on tenure distribution in the Labor Force Survey. In Finland, this can be done from 1982 onwards.

Though the approach based on comparing the changes in elapsed duration is commonly used, we argue that it is not a particularly appealing way of analysing the *changes* in job stability. To demonstrate this, we simulate the effects of changes in the entry rate of new jobs and the hazard of job ending on this commonly used measure of job stability.

Since we have access to genuine panel data that for some cohorts cover their entire careers we can directly model changes in the hazard of job ending. We do this using a competing risks model and analyse separately exits to other jobs and exits to non-employment. We are primarily interested in how these hazards vary over time but also account for the effects of elapsed duration, age and gender.

2. Data

Our data are based on individual pension contribution records of workers covered by the Employee Pension Insurance Scheme. This data were originally collected for calculating pension accruals.

The Finnish pension system is a defined-benefit system where each employment spell contributes to the old-age pension with a fixed accrual rate multiplied by the duration of the employment spell. Current system was created in 1962 and reformed in 2005. Pension contributions are mandatory and employers are liable for arranging pension insurance for all their employees. The system is operated by private insurance companies. When the workers retire the full pension is paid by the insurance company that had insured the last employment spell. The Finnish Centre for Pensions supervises the system and acts as a clearinghouse that allocates pension liabilities to the companies that received the pension contributions and transfers the funds to the company that will pay the pension to the retiree.

To perform its task the Finnish Centre for Pensions has access to full pension contribution records of all pension insurance companies. Importantly for this study, these records include information on the starting and ending dates of all employment spells. The data set that is used for this study was created at the Finnish Centre for Pensions for developing indices that are used in converting the pension accruals to the price level prevailing at the time when the employees retire.

The sampling frame includes all individuals who have contributed to the pension system between 1963 and 2004. The sampling was done using a stratified design first selecting those born on the eighth day of each month from every fifth cohort born between 1905 and 1975. Within each cohort individuals were then picked at random until a desired sample size was reached. For the individuals that are selected into the sample, all employment spells between 1963 and 2004 are included in the data. Data contain cohort-specific sampling weights that are used throughout the paper.

In the Finnish pension system employment spells contributed to the pensions from age 23 onwards. This age limit was reduced to 18 in 2005. For this reason the data from the earlier years only contain employment spells from age 23 onwards. For the last three cohorts employment spells are recorded from age 18 onwards. However, data contain the starting date of the employment spell also for the spells that were ongoing on the 23^{rd}, and for the last cohorts, on the 18^{th} birthday.

The first observations in the data are from 1963. Also in this case the data contain starting dates of all employment spells that were ongoing on January 1^{st}, 1963 no matter when the job had started. Therefore, the sample is representative of all ongoing employment spells between 1963 and 2004 for the cohorts that are included in the sample. Naturally the follow-up period varies between cohorts. The oldest cohorts are only observed at the end of their careers and the youngest cohorts only at the beginning of their careers. The 1940 birth cohort is the only cohort that is observed from age 23 to age 64. The age range when each cohort is observed and the sample sizes available for each cohort are displayed in Table 1.

The main limitations of the data are due to changes in the pension coverage over time.

Table 1. The observation period for each cohort in the original data set

Birth cohort	Years when observed	Ages when observed	Sample size	
			Individuals	Job spells
1905	1963–1970	58–64	283	324
1910	1963–1975	53–64	295	422
1915	1963–1980	48–64	296	510
1920	1963–1985	43–64	331	658
1925	1963–1990	38–64	401	1,082
1930	1963–1995	33–64	442	1,593
1935	1963–2000	28–64	469	2,863
1940	1963–2004	23–64	456	2,878
1945	1968–2004	23–59	494	4,569
1950	1973–2004	23–54	514	4,955
1955	1978–2004	23–49	538	4,563
1960	1983–2004	23–44	533	5,890
1965	1983–2004	18–39	1,022	13,094
1970	1988–2004	18–33	1,020	13,009
1975	1993–2004	18–29	975	10,223

First, as already noted, the cohorts born before 1965 are only included in the data from their 23rd birthday. To keep the data consistent across cohorts, we have excluded all job spells ending before age 23 also for the younger cohorts. Second, the data excludes very short spells lasting for less than one month because these jobs were not insured under the Employees' Pension Scheme (TEL). Before 1965 this limit was six months and between 1965 and 1971 four months. We have no data from jobs insured under Temporary Employees' Pension Act (LEL) that covered the employees in construction, agriculture, forestry and harbor work. The main reason for this is that LEL-insurance is based on monthly gross earnings and has no need to record employment dates.

Finally, the data covers both public and private sector workers but the coverage of the public sector is incomplete before the 1980s. To minimize the effect of changes in insurance coverage, we exclude all public sector employment spells from most empirical calculations. In the empirical analysis we also exclude self-employed and farmers (both covered by their own pension schemes) and hence focus on the private sector employees covered by Employees Pension Scheme (TEL) for which we have information reported in a consistent way for the whole 42-year-period.

In addition to dates of job spells, the data contain only a limited amount of other information. Age and gender can be inferred from the id-codes. Reasons why job spells ended can be used to identify those who retire, but not to distinguish between dismissals and quits. As a partial solution, we can classify job endings as quits and layoffs based on whether the employee started a new job in the TEL sector within two weeks after the end of the previous job.

Other than removing the public sector employees, we have made only minimal adjustments to the original data. We removed short overlapping job spells in cases when a short spell begins and ends while a longer spell is ongoing. Job spells that are ongoing on December 31st, 2004 or ongoing on the day when the worker turns 65 are marked as censored. Jobs that were ongoing in January 1963 are also included in the data and coded according to their original starting date.

3. Comparison to other data sources

Most studies on job stability are based on consecutive cross-sectional surveys that have been widely available for research purposes. Several studies in the United States have used data from the Current Population Survey (CPS)1 (Farber 1997, 1998, 2007; Jaeger and Stevens 2000). It collects information on tenure with the current employer or in the current job in various mobility supplements conducted at irregular intervals since 1951.

Comparable cross-sectional survey data have also been used in several other countries to measure the changes in job stability. Heisz (1999, 2005) used monthly tenure data collected in the Canadian Labor Force Survey (CLFS) since 1976, Gregg and Wadsworth (1995, 2002) and Burgess and Rees (1996, 1997, 1998) used data from the British Labor Force Survey (BLFS) and data from the General Household Survey (GHS). Both data sets include annual information on the current tenure of the respondents since 1975 and 1974, respectively. Vejsiu (2001) used data from the Swedish Level of Living Surveys and the Swedish Labor Force Survey. These two data sources contain information on the current tenure in ongoing jobs in Sweden from 1968 onwards. The Swiss Labor Force Survey used by Sousa-Poza (2004) has gathered information on the current tenure of the respondents since 1991.

While independent cross-sectional surveys are designed to be representative for the target population in each cross-section, the fact that they lack information on eventual tenure after the interview date is unfortunate. CPS income supplements offer a slight improvement making it possible to follow respondents in two consecutive years (Stewart 2002). Similar two-year panel has been available in the aforementioned Swiss Labor Force Survey. Nevertheless, the main approach employed in studies using CPS data and comparable sources from other countries has been to analyze the changes in the current tenure distribution (Farber 1997, 1998, 2007; Jaeger and Stevens 2000) or to rely on strong assumptions that allow predicting future retention rates (Swinnerton and Wial 1995; Diepod, Neumark and Polsky 1997; Neumark, Polsky and Hansen 2000) or eventual competed duration of the job held by the respondents (Hall 1982; Ureta 1992).

Several studies have employed panel data to analyze the changes in job stability. In the United States, Marcotte (1999), Polsky (1999), Gottschalk and Moffit (2000) and Jaeger and Stevens (2000) all use the Panel Study of Income Dynamics (PSID). Since 1976, the PSID has included a question on the current tenure, and in principle the annual interviews allow one to track employment histories of the survey respondents.

Other panel data sets used in US studies on job stability include the Survey of Income and Program Participation (SIPP) and the National Longitudinal Surveys (NLS). The SIPP data used by Gottschalk and Moffit (2000) and Bansak and Raphael (2006) include information on the job histories of the individuals typically over a period lasting 32 months. The National Longitudinal Survey of Young Men and the National Longitudinal Survey of Youth, have been used for analyzing job stability by Bernhardt, Morris, Handcock and Scott (2000). These two data sets provide the researchers with a 16-year follow-up of the employment histories of young men starting from a cohort born in 1944, and include employer coding that captures job changes.

The German Socio-Economic Panel (GSOEP) was used by Winkelmann and Zimmermann (1998) and Bergemann and Mertens (2011). These data are available since 1984 and include information on both the current tenure of the individuals and their labor market transitions between consecutive interviews. In addition, when the GSOEP was initiated in 1984, the respondents were asked about the number of employers they had during the past 10 years. Even longer retrospective data is available in Britain. Booth, Francesconi and Garcia-Serrano (1999) use retrospective

[1] *See, for instance, Farber (1997, 1998, 2007), Jaeger and Stevens (2000) and Neumark, Polsky and Hansen (2000).*

employment history data gathered in 1993 as a part of the British Household Panel Survey (BHPS). This survey asked the respondents to list their employment history until September 1990 starting from the day they left full-time education.

Naturally, the risk in using retrospective data for analyzing changes in job stability is that the respondents may not recall short employment spells in distant past which could lead to a false impression of declining job stability. A potentially better alternative is to compare retrospective information in surveys conducted at different times. This approach was applied by Stevens (2005) who used data from the Retirement History Survey (RHS), the National Longitudinal Survey of Older Men (NLSOM) and the Health and Retirement Survey (HRS) to measure changes in the distribution of the longest job held by the respondents during their career.

Data from administrative registers has been used only recently for analyzing changes in job stability. Bratberg, Salvanes and Vaage (2010) use linked employer-employee data from the period 1986–2002 to investigate the changes in job stability in Norway. Mahringer (2005a, 2005b) uses administrative employment spell data based on the Austrian social security records from 1975 onwards. Main benefit of these administrative data is that following job spells over time is easier than in surveys. Our pension record data extends the follow-up times to a much longer period than the administrative data sets used in previous papers.

4. Descriptive analysis

4.1. Changes in average tenure over time

As a first attempt to describe changes in job stability we examine the changes in the average tenure over time. We pick an arbitrary date, October 15th, each year and report the average elapsed duration of the jobs that were ongoing on that day. The results are plotted in Figure 1. We report both simple averages and regression

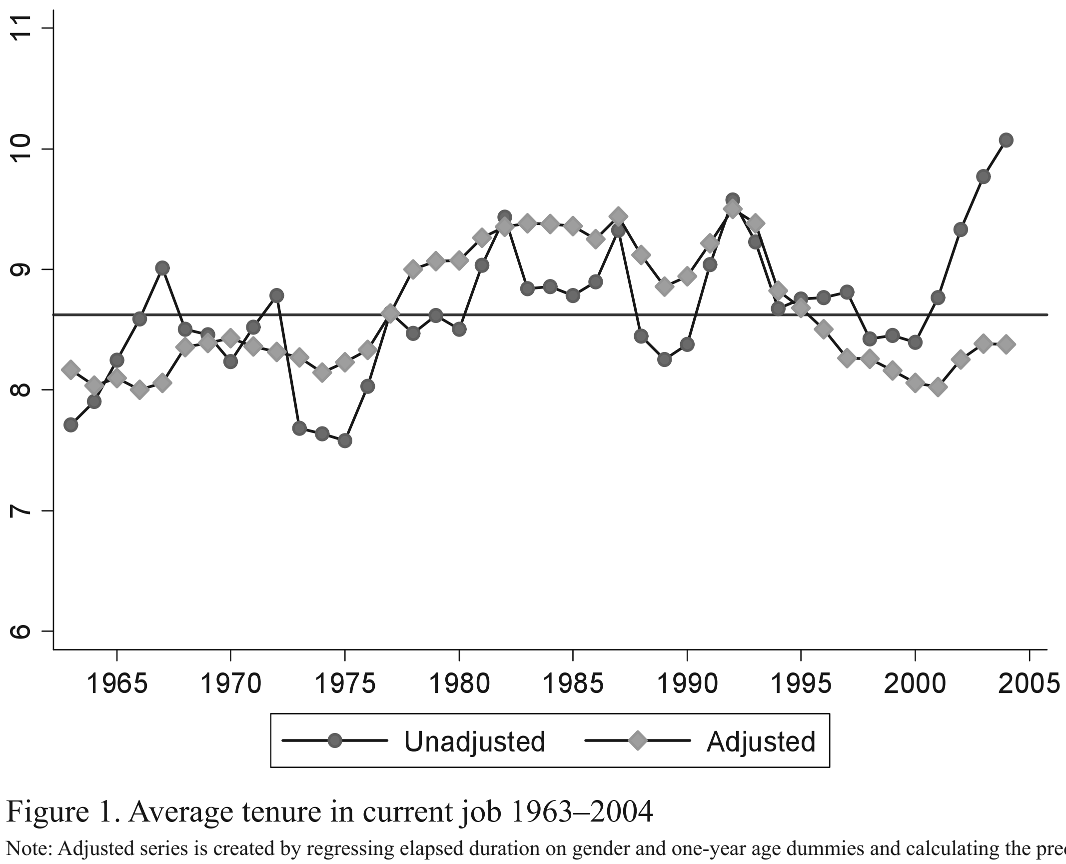

Figure 1. Average tenure in current job 1963–2004

Note: Adjusted series is created by regressing elapsed duration on gender and one-year age dummies and calculating the predicted values when gender and age distributions are set equal to the sample average.

adjusted averages that account for changes in the age and gender distribution. The horizontal line in the figure refers to the overall arithmetic average, 8.7 years.

The unadjusted numbers are trending slightly upwards over time. This partly reflects population aging. Older workers have accumulated on average longer tenure, and an increase in their share in the data increases the average tenure. It is not clear whether any adjustments for population aging should be made; an increase in average tenure is a real phenomenon even if it is caused by a change in the age structure. However, our sample is not exactly representative of the current population in any single year; it is only representative for the cohorts included in the data. Since a new cohort is added to data every fifth year, the sample gets successively older for five years and then suddenly younger as a new cohort enters in the data. The effect of sampling scheme can be seen as a modest five-year cycle in the unadjusted average tenure. The age effect is particularly strong in the end of the data. The youngest cohort born in 1975 is included in the data in 1998, but after that no new cohorts enter and the sample gets gradually older.

The largest changes in the adjusted average tenure series occur in the end of 1970s, when the average tenure increases, and in the 1990s, when the average tenure starts a gradual decline.

4.2. Comparison of register vs. survey data

To verify that changes occurring in the register data are not due to changes in reporting procedures or changes in insurance coverage, we compared the tenure distribution in the register data to the figures calculated from the Labor Force Survey[2].

Comprehensive micro-level data from the Labor Force Survey are not available for research purposes in Finland. Therefore, we cannot calculate simple statistics such as the mean tenure for years before 1997. However, Statistics Finland regularly publishes monthly data on new jobs, defined as jobs with tenure less than a year. Statistics Finland also provided us with unpublished tabulations on the tenure distribution that allow the calculation of the fraction of jobs with current tenure of more than ten years in a consistent way.

In Figure 2 we plot the fraction of workers with elapsed tenure over one year and in Figure 3 fraction of workers with elapsed tenure over ten years based on both the pension records and on the survey data. We have made the sources as comparable as possible. Since the annual interview of the LFS was held between September and December in the 1980s we use data from the last quarter of the year also after 1997. After 1980s the coverage of the register data is better so that, in contrast to the other analyses in this paper, also the public sector employees and workers under the age of 23 can be included in this comparison.

According to Figures 2 and 3 the evolution of the tenure distribution is very similar in the survey and register data. According to both sources, on average, about 80% of the workers have been in their jobs for more than a year. Also the changes in tenure distribution seem similar. According to both sources there was a large increase in the fraction of workers with more than one year of tenure in the beginning of the 1990s. This does not imply that job markets were more stable during those years, rather the opposite. Finland experienced a major recession in the beginning of the 1990s

[2] Questions on the current tenure were first added to the annual interview of the Labor Force Survey in the fall of 1982. The question on current tenure was included in the survey every year until 1987 and then every other year between 1987 and 1993. In 1995 and 1996, the tenure question was included in the EU Labor Force Survey conducted in spring. From 1997 onwards the question on elapsed tenure has been included in the monthly Labor Force Survey. In addition to changes in survey dates, also the survey question has been changed slightly which may make the LFS time-series less consistent over time.

Figure 2. Share of workers with elapsed tenure > 1 year in register vs. survey data

Figure 3. Share of workers with elapsed tenure > 10 years in register vs. survey data

Notes to Figures 2 and 3: Elapsed tenure in register data is calculated based on jobs ongoing on October 15th each year. Data includes workers between 18 and 64 years of age and cover both public and private sector employees. Self-employed are excluded. Survey data up to 1993 is based on annual interview of the Labor Force Survey. Data from 1995 and 1996 are from EU Labor Force Survey. From 1997 onwards data is based on monthly Labor Force Survey. To avoid inconsistencies due to differences in survey dates we have used the numbers from the last quarter of each year.

and very few workers were recruited during those years. Hence, the fraction of new workers with short tenure declined and average tenure increased.

Survey and register data seem to produce similar numbers also in Figure 3 that displays the fraction of workers with more than ten years of tenure. According to either source about 35 to 40% of workers have been working for their current employer for more than ten years. There is also a slight increase over time, again potentially explained by population aging.

5. Effects of the changes in entry rates and the risk of job loss on average tenure

Though commonly used, the change in mean elapsed duration of ongoing employment spells may not be a particularly informative statistic for measuring the changes in job stability.3 To illustrate this we simulate the dynamic effects of changes in job stability defined as the change in the risk of job ending on average tenure.

We follow Lancaster (1990) and start by noting that the number of persons employed at time *t* is

$$(1) \qquad \int_0^\infty n(-y)S_{-y}(y)dy,$$

which depends on the number of entrants *y* periods ago *n*(–*y*) and their historical *y*-period survival rates $S_{-y}(y)$. Hence, the mean elapsed duration of ongoing spells at time *t* is

$$(2) \quad \mu_t = \frac{\int_0^\infty n(-y)S_{-y}(y)y\,dy}{\int_0^\infty n(-y)S_{-y}(y)dy}.$$

[3] *See, for instance, Farber (1997, 1998, 2007), Jaeger and Stevens (2000) and Gregg and Wadsworth (1995, 2002).*

In general, expression (2) depends on all past entry rates and all past survival rates. For example Ureta (1992) notes that calculating average completed tenure based on survival rates computed from a cross-section data, as was done in the seminal paper by Hall (1982), leads to a bias if the arrival rates are not constant. Ureta's example of non-constant arrival rates had to do with increased labor force participation rates by women. Similar effects could be caused by large scale changes in immigration, changes in school-leaving age or major swings in the business cycle.

However, even if the entry rates were constant, the mean elapsed duration of ongoing jobs depends on all past survival rates and not only on the recent changes that the measure of changes in job stability should capture. Hence, for example a major recession that causes a temporary shock to the job exit rates affects average tenure long after the recession has ended.

In Figure 4 we illustrate the effect of an increase in the labor market entry rate on the average duration of ongoing jobs. We start from a stationary state where the labor market entry rate is constant and equals the labor market exit rate. We estimate a Kaplan-Meyer survival function from our pension record data and use expression (2) to calculate average duration of ongoing jobs. To simplify calculations, we assume that there is no unemployment so that the workers whose contracts end immediately find a new job where their tenure is naturally initially zero.

We then increase the entry rate with a constant number of new entrants each year so that total employment increases by 20 percent in 20 years and remains constant thereafter. We use equation (2) to calculate average tenure in ongoing jobs each year from t = 0 to t = 100. As illustrated in Figure 4 the increase in labor market entry decreases the average tenure because there are more recent entrants with short tenure. Interestingly, the effect of an increase in the entry rate has a long-lasting effect. Even at time t = 30 the average tenure is substantially below the initial level even

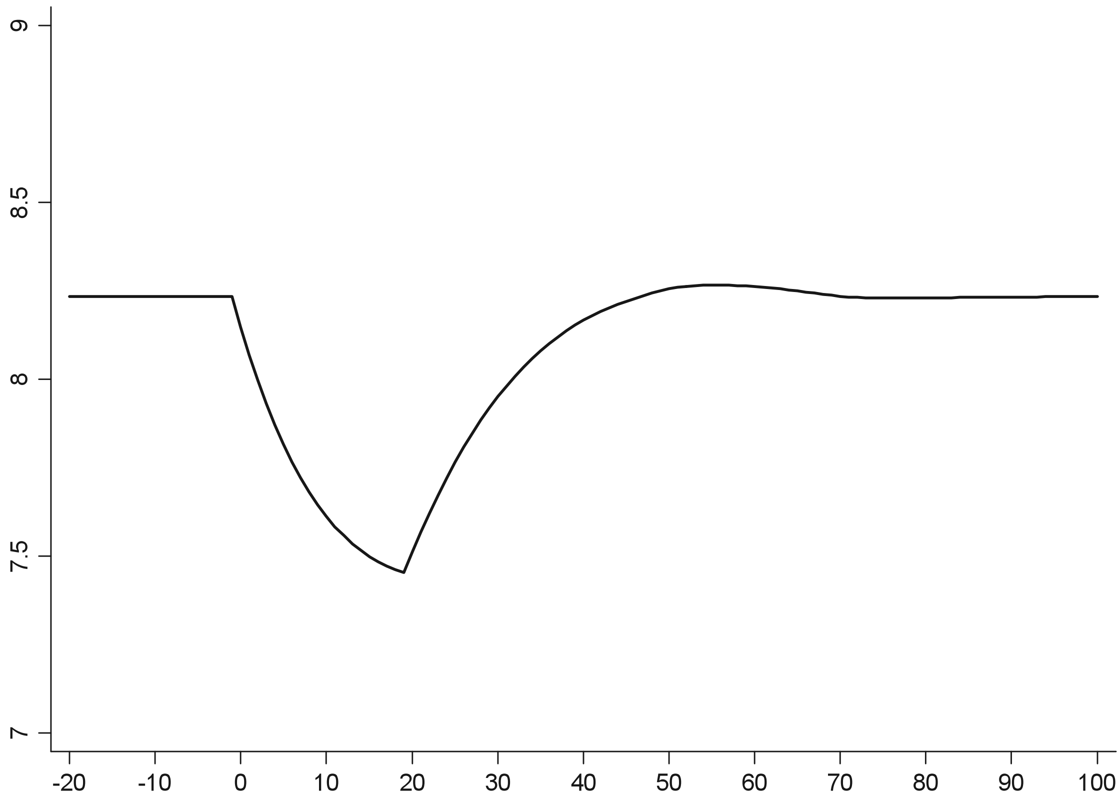

Figure 4. Simulation results: Effect of an increase in entry rate on average tenure

though the increase in the entry rate had ceased ten years ago. The average tenure stabilizes to the initial level after about forty years.

A researcher comparing average tenure generated by our simulation between, for example, years $t = 0$ and $t = 30$ might well conclude that average tenure has declined and jobs become less stable. The conclusion would obviously be misleading. The survival function of new jobs, and hence the average eventual completed tenure, is constant in our simulations. The decrease in average elapsed tenure in a cross-section is entirely due to an increase in the number of recent entrants.

In Figure 5 we present results from another simulation more closely related to changes in job stability. We keep the number of new entrants constant but simulate the effects of a temporary increase in the hazard of job ending. We start again with survival function estimated from our pension records data but increase the hazard of job loss by 50 percent for five years for all cohorts that are in the labor market between years $t = 0$ and $t = 5$. After five years we return the hazard rate to its initial level. Average elapsed tenure is again calculated based on equation (2) for time periods from $t = 0$ to $t = 100$.

As shown in Figure 5, the increase in the hazard of job ending decreases the average tenure. Note that this is partly due to our simplifying assumption according to which the workers who lose their jobs are immediately re-employed (with zero tenure). This would also imply that hiring rate would have to increase. If an increase in the job ending rate led to an increase in unemployment, the change in average tenure would depend on the changes in re-employment rates. In a typical recession the re-employment rate decreases and leads to an increase in average tenure (Burgess and Rees 1996). The number of new (low tenure) jobs typically starts to increase only when recession is over. It is therefore better to interpret our simulation as an impact of temporarily higher volatility rather than as an effect of recession.

The most interesting feature of simulation results presented in Figure 5 is that a temporary increase in volatility has long-lasting effects on average tenure. Even though the shock to the job ending rates only lasts for five years the average tenure is substantially below the initial level several decades afterwards. Again, data on the mean elapsed duration of ongoing jobs in consecutive cross-sections would give a misleading picture on the changes in uncertainty.

6. Changes in the hazard of job ending – Results from a duration model

The most natural measure of labor market uncertainty is probably the risk that a job ends. A standard way of modeling this risk is the conditional probability that a job ends given that it has lasted for a given time, i.e. the hazard rate. Modeling the effects of exogenous covariates on the hazard rate – instead of on e.g. the completed duration of a job – also provides a simple way to account for censoring due to a finite observation period, duration dependence, and the effects of covariates that vary over time.

Our data is a sample of jobs ongoing on January 1st, 1963 and of jobs beginning between 1963 and 2004. It is therefore a mixture of stock and flow samples. Jobs that began before 1963 are observed only if they are still ongoing in 1963 creating a left-truncated sample. Left truncation arises also because we observe only employment spells ongoing after the 23rd birthday, and because jobs lasting less than 6 months until 1965, 4 months until 1972 and a month thereafter are not included in the data. The starting date is known also for the stock-sampled jobs. The ending dates of jobs that are still ongoing at the end of 2004 are unknown, and the data is hence right-censored at this

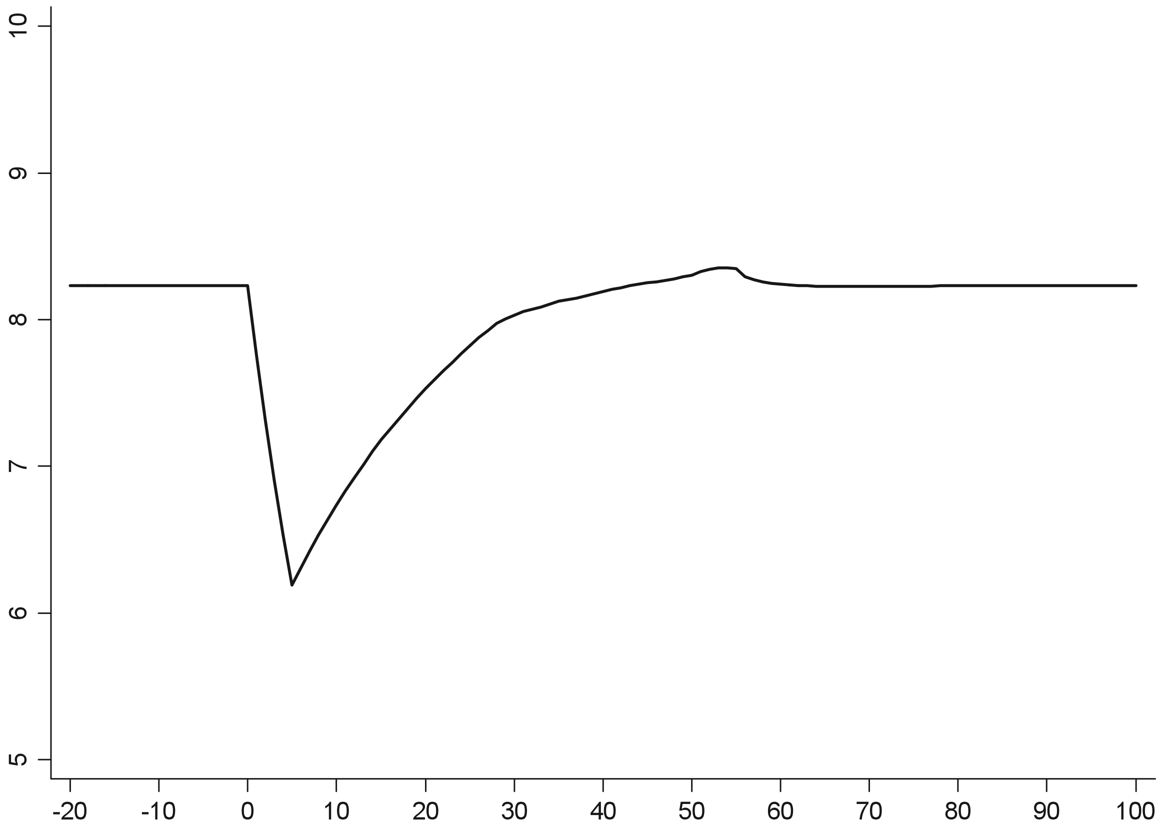

Figure 5. Simulation results: effect of an increase in the hazard of job ending on average tenure

point. We also censor jobs that are still ongoing when the worker turns 65 and becomes eligible for the old-age pension.

To model the changes in the hazard of job ending we use a competing risks model with two possible destinations: non-employment and a new job. Our definition is based on whether we observe a new employment spell within two weeks after the previous spell ends. We also experimented with a time limit of four weeks for this definition, but this had practically no effect on the results.

A large number of jobs end soon after they begin. Hazard rates decline rapidly during the first years in the job after which the rate of decline is much slower. To allow flexible forms of duration dependence we specify a piecewise constant baseline hazard function where the hazard stays constant for four month periods during the first year and for one year periods thereafter.

We explain the changes in the hazard rates by gender, age and time. Again, we aim at maximum flexibility and include the time varying covariates as a set of one-year age and time dummies. We assume that the covariates have proportional effects on the hazard rates. We also assume that there are no interactions so that, for example, duration dependence is independent of age. While these assumptions might be questioned, we would argue that they have little impact on our primary objective; consistent estimation of the time effects that capture the changes in the hazard of job ending over time.

We account for individual-specific unobserved heterogeneity by specifying a mixed proportional hazard model with three discrete points of support following the approach by Heckman and Singer (1984). The choice of the number of support points is somewhat arbitrary, but experimentation with different numbers of discrete points revealed that the other parameter estimates remained practically unchanged with two, three, or four mass points. However, imposing a restriction that there is no unobserved heterogeneity would have substantial effects on the estimates.

We use the partial likelihood method (Lancaster 1979) that makes accounting for left truncation relatively easy. Jobs that are ongoing at the start of the observation period contribute to the hazard estimates only from the entry date onwards. The jobs that end within the observation period contribute to the likelihood function through both the hazard and the survival function, but the jobs that are still ongoing at the end of the observation period contribute only through the survival function.

Our model specification can be formalized as follows. We define the hazard function related to destination s for job j of individual i conditional on the vectors of observed covariates x_{ij} and individual-specific unobserved heterogeneity terms $\nu_i = (\nu_{i1}, \nu_{i2})$ as

$$(3) \quad h_s(t_{ij} | x_{ij}, \nu_i) = \lambda_s(t_{ij}) exp(x'_{ij}\beta_s + \nu_{is})$$

We specify the baseline hazard function $\lambda_s(t_{ij})$ for destination s as

$$(4) \quad \lambda_s(t_{ij}) = \sum_{k=1}^{K} \lambda_{sk} I_{t_{ij} \in [d_{k-1}, d_k)}$$

where I is an indicator function splitting each job spell into K episodes.

Furthermore, the survival function related to destination s for job j of individual i given the vectors of observed covariates x_{ij} and individual-specific unobserved heterogeneity terms v_i is defined as

$$(5) \quad S_s(t_{ij} | x_{ij}, \nu_i) = exp\left[-\int_0^{t_{ij}} h_s(u | x_{ij}, \nu_i) du \right].$$

Let θ denote the parameter vector to be estimated and w_i yhe sampling weight of individual i. In addition, let c_{ijs} denote a dummy variable indicating whether job j of individual i is censored. The pseudo log-likelihood function can now be written as

(6)

$$log\, L(\theta) = \sum_{i=1}^{N} w_i\, log \left[\sum_{m=1}^{3} \pi_m \prod_{j=1}^{J_i} \prod_{s=1}^{2} \frac{h_s(t_{ij} | x_{ij}, \nu_m)^{1-c_{ijs}} S_s(t_{ij} | x_{ij}, \nu_m)}{S_s(e_{ij} | x_{ij}, \nu_m)} \right],$$

where the term between the brackets is the marginal likelihood contribution of individual *i*. The parameters v_m and π_m denote the mass point vectors and corresponding probabilities of the unobserved heterogeneity distribution that are estimated along with the other parameters of the model.

We parametrize the probabilities π_m of the unobserved heterogeneity distribution as

$$\pi_m = \frac{\exp(p_m)}{\sum_{m'=1}^{M} \exp(p_{m'})} \tag{7}$$

and normalize p_1 to zero. This parametrization takes automatically care of the requirement that the estimated probabilities must lie between zero and one.

The role of the division by $S_s\left(e_{ij} \middle| x_{ij}, v_m\right)$ in the pseudo log-likelihood function is to take into account the left truncation in our data. That is, we condition the likelihood contribution of each job on the fact that it must have survived until e_{ij} which denotes the elapsed duration at which job *j* of individual *i* enters the data.

Our approach produces a large number of parameters. For expositional reasons we prefer presenting the estimated hazard ratios and their 95% confidence intervals in a graphical way. The parameter estimates and their standard errors are reported in the appendix. We start by plotting the baseline hazard function in Figure 6. We omit the first category with elapsed duration of 0–4 months. As shown in the figure, the hazard of job loss declines rapidly during the first few years. After having lasted for seven years, the hazard is about 10 percent of the hazard during the first four months. After that the hazard of job loss remains approximately constant for twenty years. The hazard of job change declines also, but the decline is clearly slower. At very high durations the estimates of job change hazard get less precise due to a small number of job changes.

Figure 7 presents the effect of age on the hazard rates. We have chosen age 23 as the reference category and omit age 64 from the figure for expositional reasons. The risk of job loss first declines as the workers get older. From about age 52 onwards the hazard increases rapidly with age reflecting the effect of early retirement and gets very high after age 60. Hazard of job change declines almost linearly as the workers get older.

Finally in Figure 8 we plot the time effects using 1963 as a reference period. We find that the hazard of job change is much more volatile than the hazard of job loss. Early 1970s, late 1980s and late 1990s were years of particularly rapid job-to-job mobility. In contrast, there is only one peak in the hazard to non-employment that coincides with the recession in the early 1990s. If we interprete the hazard to non-employment as an indicator of uncertainty, we can also note that the uncertainty after year 2000 is approximately at the same level as it was forty years earlier, in the beginning of the 1960s.

A comparison of the job-to-job and job-to-non-employment hazards to the mean elapsed duration of ongoing jobs presented in Figure 1 reveals that these series appear to be almost unrelated to each other. The increase in average tenure observed around 1980 could be partially explained by lower job-to-job mobility in the end of the 1970s and the decrease in average tenure after 1993 by an increase in job-to-job mobility in the mid-1990s, but clearly something else is going on also. In addition, two possible measures of job security mean elapsed duration of ongoing jobs and hazard of job ending, point to very different time pattern in the changes of uncertainty prevailing in the labor market.

As we noted in the previous chapter, the mean elapsed duration of ongoing jobs depends not only on the current hazard of job ending, but also on past hazards and the variation in entry rates. To quantify these effects we used the estimates from a slightly simplified version of our duration model4 together with the number of new entrants each year and the initial

[4] *The model used for the simulation exercise includes only duration dependence and year effects with no unobserved heterogeneity.*

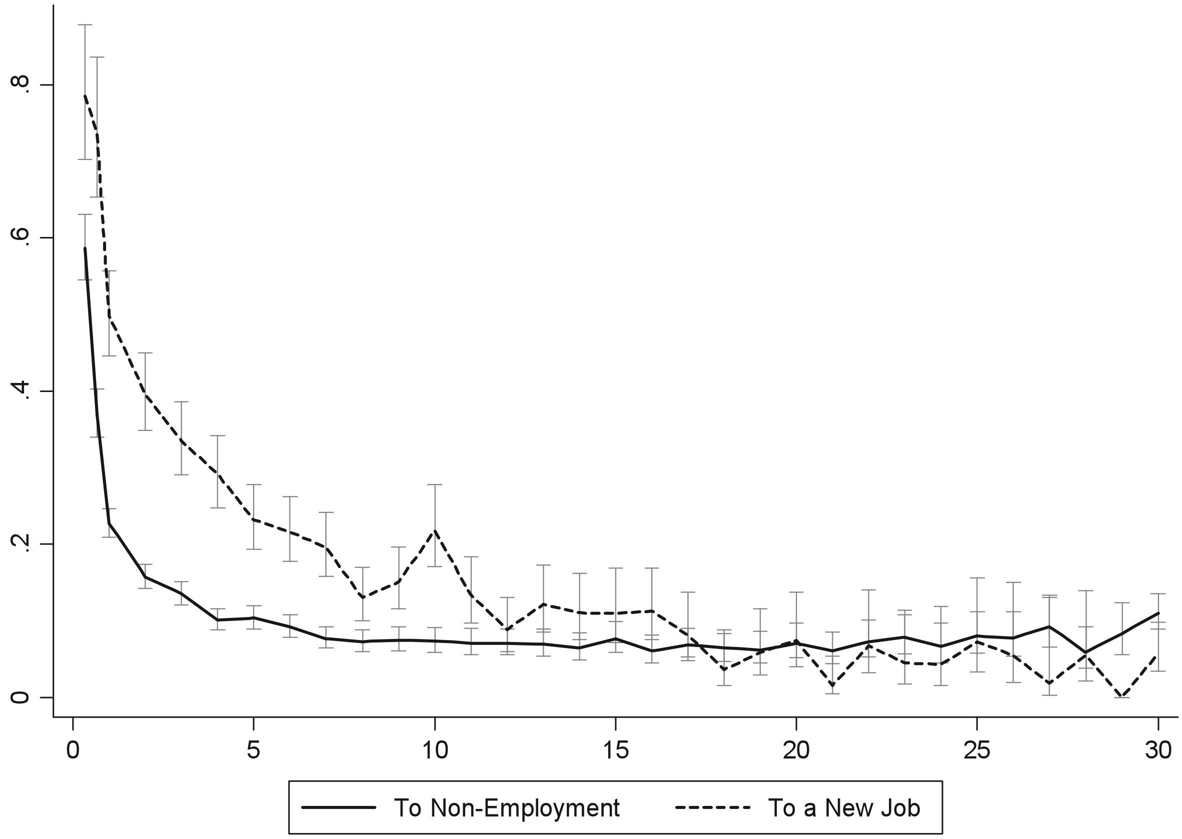

Figure 6. Hazard of job ending as a function of elapsed duration

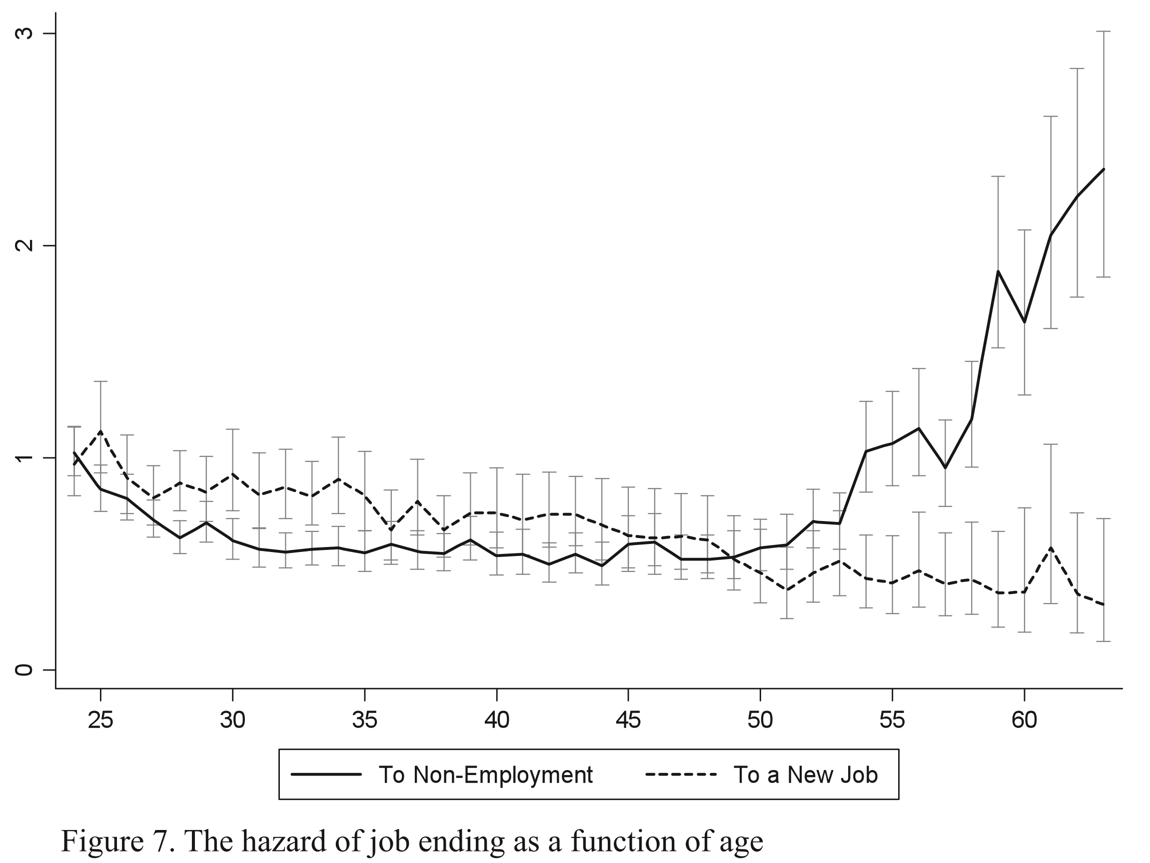

Figure 7. The hazard of job ending as a function of age

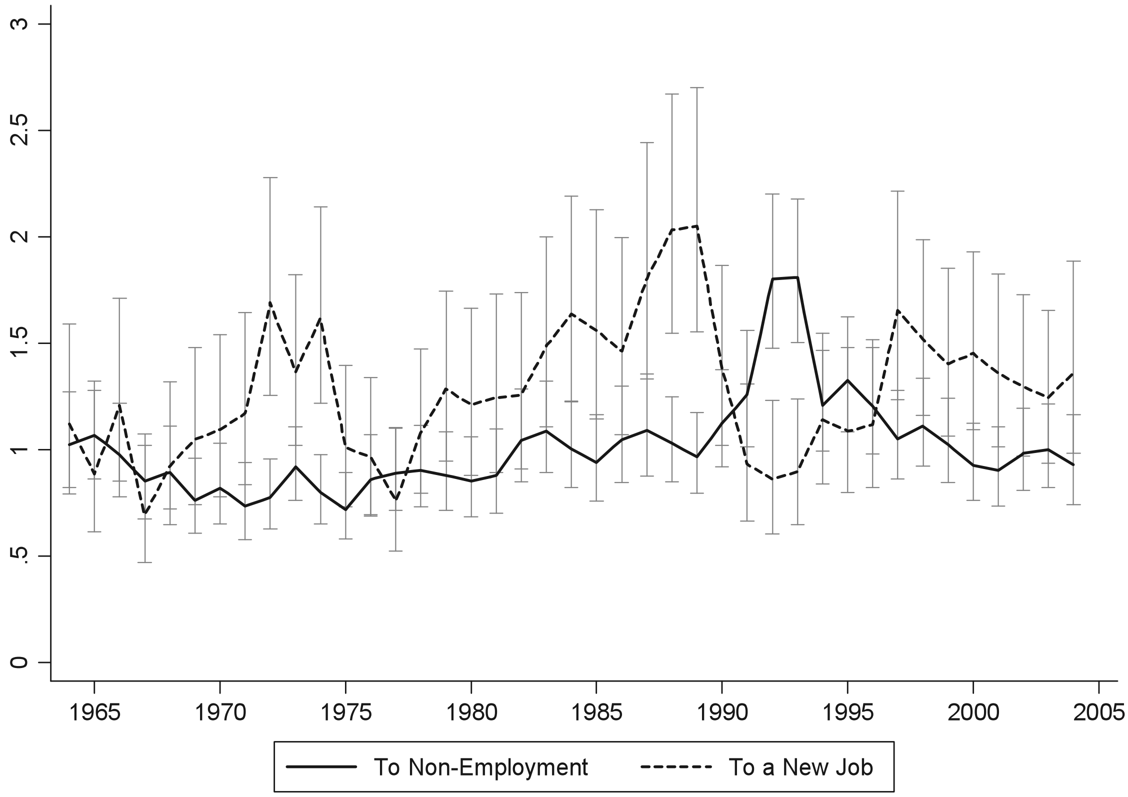

Figure 8. The hazard of job ending as a function of calendar year

duration distribution that prevailed in 1963 to simulate the average tenure in each year. As shown in the lower right corner of Figure 9 the simulated average tenure corresponds relatively well to the observed average tenure in any given year. Only major deviation between the observed and the simulated series occurs in the end of the series, and this can be explained by the changes in the age distribution of our sample. To quantify the effects of changes in job ending hazards on average tenure, we then restricted both the job-to-job and job-to-non-employment hazards to their sample averages and simulated again the implied mean elapsed tenure for each year. As shown in the top right corner of Figure 9 this has surprisingly little effect on the time pattern of the average tenure. However, as shown in the top and bottom left corner of Figure 9, if we also restrict the number of new entrants to a constant the pattern is very different irrespective of whether we keep the hazard rates constant or not. Note that even if both the entry rate and the hazard rates are constant, average elapsed duration generally changes over time since there is no reason to assume that the tenure distribution in 1963 reflected a steady state. In fact, the mean elapsed duration of ongoing jobs in 1963 was rather low, possibly due to earlier labor demand or supply shocks.

8. Conclusions

The estimates presented in this paper suggest that there has been no long-term trend in the mean elapsed duration of ongoing private sector jobs in Finland. Average tenure has still varied considerably over time, but this variation has been mainly due to the variation in the number of new entrants into the labor market. If, instead, we measure job market stability by the hazard of job loss – or more precisely by

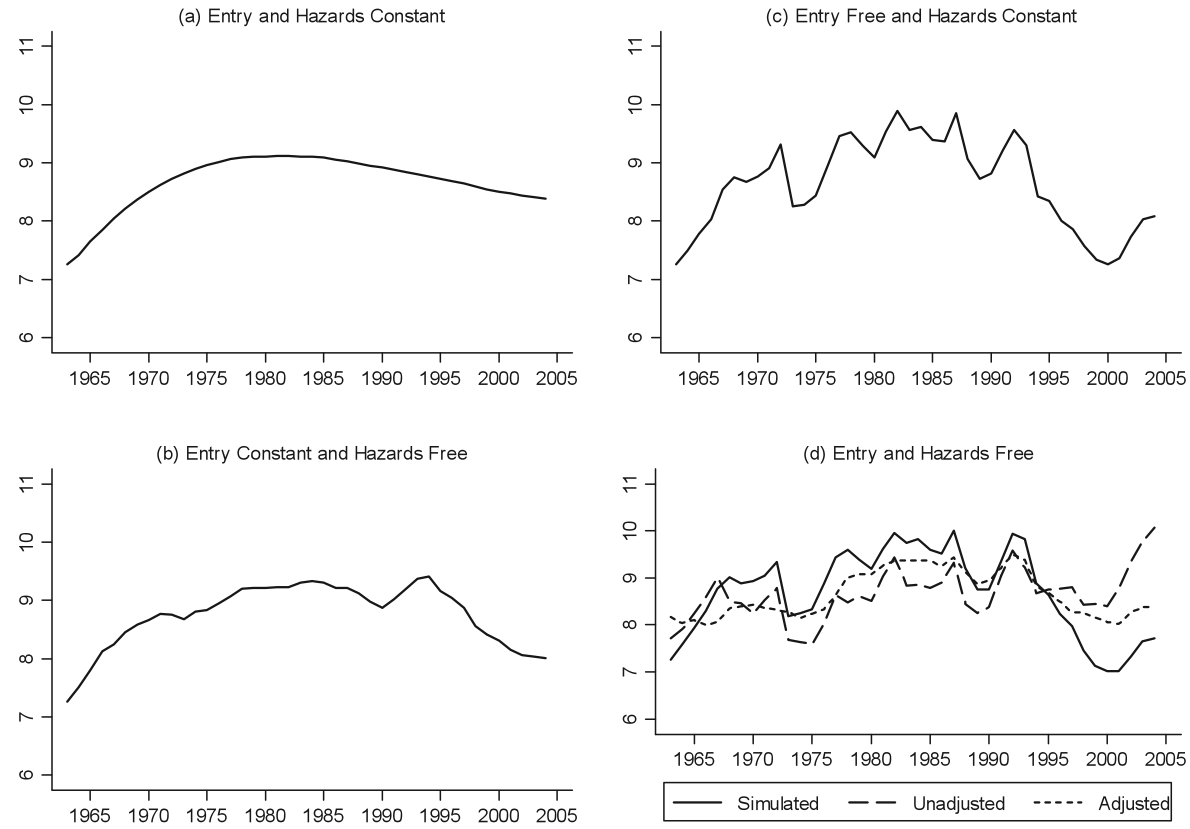

Figure 9. Simulation results

the hazard of job ending that leads into non-employment – using a standard duration model specification, we find that job stability declined during a major recession but that the risk of job loss after the year 2000 is at the level observed forty years ago.

In addition to demonstrating that job security has not fundamentally decreased over the long-term, our results also suggest that there are important caveats to be kept in mind when using the elapsed duration of ongoing jobs to measure changes in job stability. In our data the changes in average tenure appear to be mainly related to the changes in the number of new jobs rather than changes in stability of existing jobs. This implies that results from earlier studies on job stability that focus on the average elapsed duration of ongoing jobs may be quite misleading.

This paper also illustrates the benefits of using register data originally collected for administrative purposes. Pension registers are particularly useful since calculating pensions typically requires information about entire careers. Compared with previous research on job stability, we can both cover a much longer time period, starting from the early 1960s to present, and observe entire lifetime job histories.

Appendix

Table A1. Parameter estimates for the mixed proportional hazard competing risks model of job endings

	To non-employment			To a new job		
	Estimate	SE	P-value	Estimate	SE	P-value
Constant	1.626	0.147	0.000	-0.938	0.203	0.000
Male	-0.195	0.033	0.000	-0.023	0.035	0.258
Duration depencence (reference category 0)						
4 months	-0.534	0.037	0.000	-0.242	0.057	0.000
8 months	-0.995	0.043	0.000	-0.303	0.063	0.000
1 year	-1.482	0.042	0.000	-0.697	0.057	0.000
2 years	-1.848	0.051	0.000	-0.927	0.065	0.000
3 years	-2.002	0.057	0.000	-1.094	0.073	0.000
4 years	-2.291	0.069	0.000	-1.234	0.082	0.000
5 years	-2.268	0.073	0.000	-1.462	0.092	0.000
6 years	-2.385	0.079	0.000	-1.532	0.099	0.000
7 years	-2.567	0.091	0.000	-1.631	0.108	0.000
8 years	-2.620	0.097	0.000	-2.038	0.135	0.000
9 years	-2.591	0.104	0.000	-1.892	0.136	0.000
10 years	-2.613	0.109	0.000	-1.525	0.125	0.000
11 years	-2.646	0.125	0.000	-2.012	0.162	0.000
12 years	-2.646	0.120	0.000	-2.424	0.199	0.000
13 years	-2.664	0.127	0.000	-2.108	0.179	0.000
14 years	-2.742	0.135	0.000	-2.203	0.195	0.000
15 years	-2.570	0.131	0.000	-2.208	0.219	0.000
16 years	-2.799	0.151	0.000	-2.182	0.205	0.000
17 years	-2.673	0.139	0.000	-2.508	0.266	0.000
18 years	-2.743	0.159	0.000	-3.313	0.423	0.000
19 years	-2.778	0.167	0.000	-2.836	0.348	0.000
20 years	-2.647	0.160	0.000	-2.598	0.314	0.000
21 years	-2.794	0.172	0.000	-4.132	0.618	0.000
22 years	-2.618	0.167	0.000	-2.695	0.374	0.000
23 years	-2.544	0.164	0.000	-3.104	0.477	0.000
24 years	-2.711	0.195	0.000	-3.147	0.517	0.000
25 years	-2.522	0.171	0.000	-2.625	0.393	0.000
26 years	-2.556	0.188	0.000	-2.905	0.515	0.000
27 years	-2.379	0.175	0.000	-3.975	1.002	0.000
28 years	-2.830	0.228	0.000	-2.901	0.474	0.000
29 years	-2.484	0.201	0.000	-11.314	0.096	0.000
30 years	-2.209	0.106	0.000	-2.857	0.271	0.000
Year effects (reference category 1963)						
1964	0.022	0.111	0.421	0.115	0.178	0.259
1965	0.066	0.109	0.273	-0.121	0.187	0.258
1966	-0.025	0.114	0.415	0.189	0.178	0.144
1967	-0.161	0.118	0.085	-0.367	0.198	0.032
1968	-0.112	0.110	0.153	-0.079	0.182	0.332

1969	-0.271	0.117	0.010	0.047	0.176	0.395
1970	-0.199	0.117	0.044	0.091	0.174	0.300
1971	-0.307	0.124	0.007	0.158	0.173	0.179
1972	-0.256	0.108	0.009	0.526	0.152	0.000
1973	-0.085	0.096	0.186	0.311	0.148	0.018
1974	-0.225	0.103	0.014	0.480	0.144	0.000
1975	-0.329	0.109	0.001	0.011	0.165	0.473
1976	-0.152	0.113	0.090	-0.035	0.167	0.416
1977	-0.118	0.111	0.145	-0.275	0.190	0.074
1978	-0.103	0.107	0.168	0.079	0.157	0.306
1979	-0.128	0.106	0.115	0.252	0.156	0.053
1980	-0.161	0.112	0.075	0.191	0.163	0.121
1981	-0.130	0.114	0.127	0.219	0.169	0.098
1982	0.043	0.106	0.342	0.229	0.165	0.083
1983	0.084	0.100	0.202	0.398	0.151	0.004
1984	0.005	0.103	0.482	0.494	0.148	0.000
1985	-0.061	0.109	0.286	0.445	0.158	0.002
1986	0.047	0.110	0.335	0.380	0.159	0.008
1987	0.086	0.111	0.219	0.590	0.155	0.000
1988	0.029	0.098	0.384	0.710	0.139	0.000
1989	-0.035	0.100	0.363	0.718	0.141	0.000
1990	0.117	0.103	0.128	0.323	0.154	0.018
1991	0.229	0.110	0.018	-0.071	0.173	0.341
1992	0.590	0.102	0.000	-0.147	0.181	0.207
1993	0.593	0.095	0.000	-0.110	0.165	0.253
1994	0.188	0.099	0.030	0.131	0.156	0.201
1995	0.283	0.103	0.003	0.084	0.157	0.296
1996	0.187	0.105	0.038	0.110	0.156	0.242
1997	0.048	0.101	0.317	0.503	0.149	0.000
1998	0.105	0.094	0.131	0.418	0.137	0.001
1999	0.024	0.098	0.404	0.339	0.142	0.008
2000	-0.077	0.099	0.220	0.373	0.145	0.005
2001	-0.104	0.105	0.162	0.308	0.150	0.020
2002	-0.018	0.100	0.429	0.259	0.147	0.039
2003	0.000	0.099	0.498	0.219	0.145	0.065
2004	-0.073	0.115	0.264	0.309	0.166	0.031
Age effects (reference category 23)						
24	0.024	0.057	0.334	-0.029	0.086	0.367
25	-0.162	0.066	0.007	0.118	0.097	0.111
26	-0.213	0.068	0.001	-0.100	0.104	0.169
27	-0.345	0.063	0.000	-0.210	0.088	0.009
28	-0.476	0.063	0.000	-0.126	0.081	0.061
29	-0.368	0.070	0.000	-0.175	0.093	0.030
30	-0.494	0.081	0.000	-0.081	0.106	0.220
31	-0.563	0.084	0.000	-0.191	0.110	0.042
32	-0.585	0.075	0.000	-0.150	0.096	0.060
33	-0.564	0.070	0.000	-0.200	0.093	0.016
34	-0.552	0.082	0.000	-0.106	0.101	0.148

35	-0.593	0.089	0.000	-0.196	0.116	0.045
36	-0.526	0.086	0.000	-0.413	0.126	0.001
37	-0.584	0.082	0.000	-0.230	0.114	0.022
38	-0.599	0.081	0.000	-0.414	0.110	0.000
39	-0.488	0.085	0.000	-0.302	0.117	0.005
40	-0.618	0.095	0.000	-0.300	0.128	0.010
41	-0.604	0.098	0.000	-0.345	0.135	0.005
42	-0.694	0.094	0.000	-0.307	0.122	0.006
43	-0.609	0.089	0.000	-0.312	0.113	0.003
44	-0.710	0.104	0.000	-0.381	0.142	0.004
45	-0.523	0.105	0.000	-0.456	0.157	0.002
46	-0.507	0.104	0.000	-0.476	0.164	0.002
47	-0.650	0.101	0.000	-0.463	0.143	0.001
48	-0.649	0.100	0.000	-0.489	0.150	0.001
49	-0.632	0.109	0.000	-0.649	0.168	0.000
50	-0.552	0.107	0.000	-0.782	0.190	0.000
51	-0.527	0.110	0.000	-0.982	0.221	0.000
52	-0.358	0.100	0.000	-0.783	0.185	0.000
53	-0.373	0.099	0.000	-0.666	0.194	0.000
54	0.031	0.105	0.384	-0.842	0.198	0.000
55	0.066	0.106	0.267	-0.891	0.221	0.000
56	0.131	0.112	0.122	-0.760	0.236	0.001
57	-0.047	0.108	0.334	-0.903	0.237	0.000
58	0.166	0.107	0.060	-0.849	0.250	0.000
59	0.631	0.109	0.000	-1.015	0.301	0.000
60	0.495	0.120	0.000	-1.001	0.373	0.004
61	0.718	0.123	0.000	-0.550	0.313	0.039
62	0.803	0.122	0.000	-1.025	0.369	0.003
63	0.859	0.124	0.000	-1.177	0.429	0.003
64	2.394	0.106	0.000	-1.437	0.715	0.022
Unobserved heterogeneity						
m2	-0.330	0.215	0.062	-1.915	0.139	0.000
m3	-1.034	0.171	0.000	-1.016	0.074	0.000
p2	1.737	0.413	0.000			
p3	1.465	0.275	0.000			
Pseudo log-likelihood						-4385653

Notes: Maximum likelihood estimates using sampling weights. Standard errors based on Huber-White sandwich estimator.

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