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From Lifetime Jobs to Churning?^a

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1. Introduction

It is often taken for granted that lifetime jobs have become scarcer than in the past.¹ In Switzerland, the death of the “job for life” paradigm has received important coverage in the media since the late 1990s, although the picture was somewhat exaggerated by single but large and visible events like the privatization

- a Former versions of this paper were presented at the Interdisciplinary Congress on Research in Vocational Education and Training, Berne (Switzerland), March 2011; at the Spring Meeting of Young Economists (SMYE), Groningen (Netherlands), April 2011; at the IZA European Summer School in Labor Economics, Ammersee (Germany), May 2011; at the Annual Congress of the European Society for Population Economics (ESPE), Hangzhou (China), June 2011; and at the Annual Meeting of the Swiss Society of Economics and Statistics (SSES), Neuchâtel (Switzerland), June 2013. We thank Ulrich Blum, Martin Carnoy, Christian Dustmann, Jean-Marc Falter, Yves Flückiger, Jennifer Hunt, Marcelo Olarreaga, José Ramirez, Eric Verdier, and two anonymous referees for helpful suggestions that led to improvements. Financial support from the *Swiss Leading House “Economics of Education” (University of Geneva)* is gratefully acknowledged. The usual disclaimer applies.
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- 1 As the following few recent newspaper articles show, such feeling arise in many countries. *The Washington Post* (“The Scared Workers”, 2 September 2013): “lifetime employment at one company is dead.” *The New York Times* (“Young, Educated And Jobless In France”, 3 December 2012): “In our parents’ generation, you had a job for life; now we constantly have to change jobs, change companies, change regions. [...] The trend is especially evident in France, where 82 percent of people hired today are on temporary contracts.” *The Economist* (“Still Work to be Done”, 29 November 2007): “the traditional Japanese ‘lifetime employment’ model was deeply entrenched. It is often said that this model is now collapsing and that the era of ‘jobs for life’ has come to an end.” FARBER (1998) quotes similar newspaper articles published at the beginning of the 1990s, showing that such feelings are not new.

of the telecommunications sector or the bankruptcy of the Swiss national airlines.²

This paper investigates the evolution of job stability in Switzerland over the 1990s and 2000s. Understanding how job stability evolves is of high concern from both an academic and policy making perspective, as it helps shaping institutions like unemployment insurance schemes or employment protection legislation. It also allows to verify if the popular feeling of general increasing job instability is justified, and if so, why.

In the last two decades, technical change and globalization have both been pointed out as the main sources of change on wage inequality and unemployment (see e.g., AUTOR, KATZ, and KEARNEY, 2008). One could expect these shocks to affect the length of employment relationships as well. More precisely, skill-biased technical change can be expected to jeopardize the situation of low-skilled workers while at the same time fostering employment perspectives for high-skilled ones. Greater trade openness implies stronger competition on the product market, which might translate into changes in the structure of employment. For instance, BISCOURP and KRAMARZ (2007) show that increased imports of finished goods are negatively associated with unskilled labor and LO TURCO, MAGGIONI, and PICCHIO (2013) find that offshoring significantly reduces job stability in the Italian manufacturing sector, with blue collar workers being the most affected. However, governments strive to increase labor market security in order to cushion the shocks caused by greater openness (RODRIG, 1998).

A related factor to explain the possible drift away from lifelong jobs lies in the gradual shift from manufacturing to services.³ As mentioned by FARBER (2007), it seems that high-quality manufacturing jobs are lost, perhaps to import competition, and are being replaced by low-quality service sector jobs (so-called hamburger-flipping jobs). Manufacturing relies more heavily on specific human capital which requires a longer subsequent period of time to recoup the cost of

- 2 Swisscom went public in October 1998. In the newspaper *Le Temps*, 7 August 1998, labor union representatives mentioned “fear of dismissals” and “loss of confidence in the future”. It was planned that, in 2001, “the civil servant status [would] be removed and Swisscom [would] slash 4,000 jobs.” The Swiss national airlines went bankrupt in October 2001. In *Le Temps*, 4 April 2001, an employee was summarizing his situation as follows: “If I lose my job at Swissair, I’ll go elsewhere. I no longer believe in jobs for life.” (Translations from the French by the authors.)
- 3 In Switzerland, the industrial sector’s share of employment has fallen from 47% in 1960 to 23% in 2011. Over the same period, services’ share almost doubled, from 39% to 74% of the workforce (see Federal Statistical Office: Employment and income from employment: Panorama 2013).

investment. In services, employers demand more general and versatile skills from workers, with lower sunk costs of training.

Switzerland clearly belongs to the league of countries with moderate employment protection, its OECD EPL index being one of the lowest in the world.⁴ The Swiss labor market is in fact characterized by a substantial flexibility in the employment relationship, high openness to trade and migration, and mostly peaceful industrial relations and low unemployment. If technical change and trade do have an impact on the length of jobs, one should therefore be able to observe changes in the employment relationships in Switzerland.

The Swiss labor market is also of particular interest because of important changes that took place in the last two decades. Bilateral agreements with the EU members became effective in 1999, granting free movement of workers. The agreements were further extended to all new EU countries in 2005. Another important change that has taken place in 2005 is the introduction of compulsory maternity leave insurance, which makes it easier for women to keep their jobs in case of a child birth. It is well known that, until recently, Swiss women often interrupted their career in their mid-career for child care reasons. The tendency is now to maintain some attachment to the labor market (e.g., through part time employment).

While the literature mostly focuses on job (in)stability, job (in)security is a distinct concept. Job instability simply relates to the duration of jobs, whereas job insecurity refers to all forms of welfare-reducing uncertainty surrounding employment. Job insecurity therefore implies that job terminations have undesired consequences for workers, whereas job instability is a mere quantitative concept.

Job stability and job security do not necessarily evolve together. For example, if workers decide voluntarily to change jobs more often, job stability would decrease without affecting job security. Conversely, if (involuntary) layoffs become more frequent while (voluntary) quits become less frequent, job stability could remain constant while job security would definitely decrease.

Measuring job insecurity is problematic, and it has been investigated using workers' beliefs about their job loss likelihood (SCHMIDT, 1999). Subjective expectations about employment and unemployment, however, tend to over-estimate the risk of job loss (DICKERSON and GREEN, 2012). In this paper, we do not tackle the question of job insecurity directly, but our analysis might offer some insights. In our estimations, we use competing risks duration models to break down employment spells across destination states following their termination

4 Employment protection legislation (EPL) is discussed in OECD (2004), and updated statistics are available at <http://www.oecd.org/employment/protection>.

(new job, unemployment or inactivity) and the events that caused the latter (quits or layoffs), along the lines of BOOTH, FRANCESCONI, and GARCIA-SERRANO (1999), GOTTSCHALK and MOFFITT (1999), HIRSCH and SCHNABEL (2010), and BERGEMANN and MERTENS (2011). To some extent, transitions from employment to unemployment and the propensity of layoffs may well reflect job insecurity.

An additional hypothesis we investigate in details is that of possible differences across age groups. Our contention is that the general feeling of a more uncertain labor market might arise from decreasing job stability among older workers, given that job losses have serious consequences beyond a certain age. Even though no changes currently occur to them, young workers might anticipate potential troubles and begin to fear more for their career now. Previous research suggests that disadvantaged groups, such as youths and women, may benefit most from job market flexibility (see for example EUROPEAN COMMISSION, 2006). Therefore, different age groups may well be affected differentially over time. One could also argue that young individuals are quicker in adapting to new technologies. On the contrary, and potentially due to slower adoption of skills, older workers enjoy less inter-job mobility, but are also more likely to suffer from long term unemployment if made redundant. In order to test for these potential differences, we provide estimations that separate workers across age groups.

The remainder of the paper proceeds as follows. Section 2 provides an overview of the literature on job stability. Section 3 presents our dataset and provides descriptive statistics. Section 4 explains the difficulties encountered when analyzing tenure data and shows why duration analysis is appropriate for such a task. Section 5 presents the results of a series of Cox proportional hazards models. Section 6 summarizes the findings and concludes.

2. Literature on Job Stability

2.1 *Empirical Studies*

The literature on job stability and lifetime jobs finds its origins in the seminal papers by HALL (1982) and URETA (1992). Thereafter, numerous results on job stability have been produced. Table 1 provides the characteristics of a number of studies that investigate the evolution of job stability.⁵ Studies are organized by country and year of publication. These studies comprise different countries, time periods, and measures of job stability. However, they all find either constant or

5 GOTTSCHALK and MOFFITT (1999) display a similar table focusing on US studies.

falling job stability. From the countries where several studies are available, one might infer that data sources have a significant impact on the results: it appears that an analysis based on the same datasets, even over different time periods, yield similar conclusions.

Table 1 highlights how studies making use of duration models are still a minority so far, but all papers published since 2010 use this technique. Duration analysis is the most adequate technique to analyze job spells, since it readily handles right-censoring and left-truncation (see Section 4). These two difficulties are sidestepped by OLS regressions on elapsed tenure and by logit regressions on the probability of having held a job for less or more than some threshold. Indeed, such models do not account for the fact that ongoing job spells may last for many additional years (right-censoring) and that stock sample surveys suffer from an over-representation of long job spells (left-truncation).⁶

Job stability has already been studied in Switzerland by SOUSA-POZA (2004). Here, we use updated data in order to re-examine this question. We also use duration models, and split job separations into different destination states and different termination reasons. The case for this more complex analysis is made stronger by the fact that job-to-job moves have been shown to differ systematically from job-nonemployment-job moves (ROYALTY, 1998). We also account for right-censoring and left-truncation, so that we are able to retrieve information from all job spells observed in our survey, whether or not they are ongoing at the beginning of observation, and whether or not they are completed by the end of observation.

On the econometric side, the paper closest to ours is BERGEMANN and MERTENS (2011), as they estimate duration models that control for left-truncation and distinguish between destination states and job termination reasons. The other existing papers using duration models either (i) do not account for left-truncation (BOOTH, FRANCESCONI, and GARCIA-SERRANO, 1999; GOTTSCHALK and MOFFITT, 1999), (ii) do not study the evolution of job stability but compare different socio-economic groups (MARINESCU, 2009; HIRSCH and SCHNABEL,

6 A simulation exercise reported by ROKKANEN and UUSITALO (2010) demonstrates the superiority of duration analysis over other models to analyze job stability. Comparing average elapsed tenure to infer changes in job stability might in fact be misleading: average elapsed tenure might decrease simply as a consequence of an increase in the number of new labor market entrants. Moreover, average elapsed tenure is plagued by inertia: a temporary short-term increase in the risk of job loss leads to a long-term decrease in mean elapsed tenure. Duration analysis is based on the estimation of the hazard rate, i.e., the instantaneous risk of job separation given that the job spell has lasted up to current instant, and therefore detects labor market variations much faster.

Table 1: Comparison of Studies on the Evolution of Job Stability

Study	Country	Dataset	Analysis	Overall finding
SWINNERTON and WIAL (1995, 1996)	US	CPS supplements 1979–1991	4-year retention rates	Decrease in job stability
DIEBOLD, NEUMARK, and POLSKY (1996, 1997)	US	CPS supplements 1983–1991	4- and 10-year retention rates	No clear trend
BOISJOLY, DUNCAN, and SMEEDING (1998)	US	PSID 1968–1992	Annual rate of involuntary job loss	Increase in involuntary job losses
FARBER (1998)	US	CPS supplements 1973–1993	Changes in tenure quantiles, $Pr(T < 1)$, $Pr(T \geq 10)$, $Pr(T \geq 20)$	No clear trend
GOTTSCHALK and MOFFITT (1999)	US	SIPP 1984–1995, PSID 1981–1992	Duration models	No clear trend
JAEGER and STEVENS (1999)	US	PSID 1981–1996, CPS supplements 1973–1996	$Pr(T \leq 1)$, $Pr(T < 10)$	Increase in $Pr(T < 10)$ but no trend in $Pr(T \leq 1)$
FARBER (2009)	US	CPS supplements 1973–2008, DWS 1983–2008	OLS on elapsed tenure, $Pr(T < 1)$, $Pr(T \geq 10)$, $Pr(T \geq 20)$	Decrease of long-term employment
HEISZ (1999)	Canada	CLFS 1981–1996	Distribution of completed tenure	No clear trend
HEISZ (2005)	Canada	CLFS 1976–2001	1- and 4-year retention rates	No clear trend
GREGG and WADSWORTH (1995)	UK	BLFS 1975–1993	Distribution of elapsed tenure	Decrease in job stability and increase in inequality of job stability
BURGESS and REES (1996)	UK	GHS 1975–1992	Distribution of elapsed tenure	No clear trend
BURGESS and REES (1998)	UK	GHS 1975–1993	$Pr(T < 1)$, $Pr(T \geq 5)$	No clear trend

Study	Country	Dataset	Analysis	Overall finding
BOOTH, FRANCESCO, and GARCIA-SERRANO (1999)	UK	BHPS 1990 (retrospective information 1915–1990)	Duration models	Decrease in job stability
GREGG and WADSWORTH (2002)	UK	BLFS 1975–2000, GHS 1975–1998	Median elapsed tenure, $\Pr(T < 1)$, $\Pr(T > 5)$, $\Pr(T > 10)$	Decrease in job stability
WINKELMANN and ZIMMERMANN (1998)	Germany	GSOEP 1974–1994	Count data models for the number of job changes	If anything, increase in job stability
BERGEMANN and MERTENS (2011)	Germany	GSOEP 1984–1997	Duration models	Decrease in job stability, driven by an increasing hazard of being laid off
GIANNELLI, JAENICHEN, and VILLOSTO (2012)	Germany	IABS 1975–2004	Duration models	Decrease in job stability
ROTHE, GIANNELLI, and JAENICHEN (2013)	Germany	WHIP 1985–2003	Duration models	Decrease in job stability
MAHRINGER (2004)	Austria	IABS 1998–2010	Duration models	No trend for men, increase in job stability for women
GIVORD and MAURIN (2004)	France	ASSR 1993–2003	1- to 3-year retention rates	No clear trend
SOUSA-POZA (2004)	Switzerland	FLFS 1982–2002	Job loss rates	Increase in the risk of involuntary job loss
BRATBERG, SALVANES, and VAAGE (2010)	Norway	SLFS 1991–2001	1-year retention rates	No clear trend
ROKKANEN and UUSITALO (2010)	Finland	SNR 1986–2002	$\Pr(T < 1)$, $\Pr(T \geq 8)$, job separation rates	Very weak indications that job stability has decreased
			Duration models	No clear trend

Note: ASSR = Austrian Social Security Records, BHPS = British Household Panel Survey, BLFS = British Labor Force Survey, CLFS = Canadian Labor Force Survey, CPS = Current Population Survey, DWS = Displaced Workers Survey, FPIIS = Finnish Pension Insurance Scheme, FLFS = French Labor Force Survey, GHS = General Household Survey, GSOEP = German Socio-Economic Panel, IABS = (German) Institute for Employment Research Sample, PSID = Panel Study for Income Dynamics, SIPP = Survey of Income and Program Participation, SLFS = Swiss Labor Force Survey, SNR = Statistics Norway Registers, WHIP = Work Histories Italian Panel.

2010; ROKKANEN and UUSITALO, 2010; LO TURCO, MAGGIONI, and PICCHIO, 2013), or (iii) do not distinguish between several possible destination states (GIANNELLI, JAENICHEN, and VILLOSIO, 2012; ROTHE, GIANNELLI, and JAENICHEN, 2013).

2.2 Models Explaining Changes in Job Tenure

As highlighted in subsection 2.1, most recent research on job stability has focused on documenting trends empirically. Only a handful of papers build models to explain the observed changes in job tenure. To our knowledge, VALLETTA (1999) provides the first theoretical framework to explain the evolution of job security. In his model, firms choose the level of specific investment in worker productivity, producing either robust or fragile contracts.⁷ Job security is closely linked to the incidence of fragile contracts, which increases when: (i) the probability of negative economic shocks falls, (ii) the value of shirking (by the worker and the firm) increases, and (iii) the probability of finding an outside match increases. The development of the information and communications technology (ICT) has certainly increased the probability of finding a new job and facilitated on-the-job search. Due to rising migration flows, the probability for the firm of finding an outside match has also increased. In these circumstances, VALLETTA'S (1999) model predicts an increase in the incidence of fragile contracts, which is tantamount to an increase in job insecurity.

FRIEDBERG and OWYANG (2004) build a matching model with moral hazard to explain changes in job tenure and pension structure. The model incorporates the use of defined benefit (DB) pensions as a device to deter shirking by workers.⁸ According to their model, potential explanations for the decline in job duration and DB pensions are: (i) an increase in the frequency of shocks that reduce the

7 Under a robust contract, the level of investment and resulting productivity is high enough to preserve the job even in the case of an adverse shock on economic conditions. Under a fragile contract, the level of investment is inadequate to preserve the job pairing when an adverse shock arises.

8 Two different types of pensions are available in the US, and the system is similar in Switzerland. Defined benefit (DB) pensions (*Leistungsprimat, primauté des prestations*) offer a defined payout to workers after they leave an employer. The payout is fixed in percentage of the worker's wage. Defined contribution (DC) pensions (*Beitragsprimat, primauté des cotisations*) offer a payout that depends on employer and employee contributions, and these pension plans are portable across employers. Both in the US and in Switzerland, the evolution of pension structure followed a similar pattern: DB pensions are becoming less frequent, while DC pensions become the norm.

value of existing matches relative to new matches and (ii) an increase in uncertainty about future productivity. They also argue that new technologies appear to have reduced the value of existing jobs relative to new jobs and raised uncertainty, so that accelerating technological change might explain the joint decline in job tenure and DB pensions.

ZAVODNY (2003) provides detailed explanations of the possible theoretical relationship between technological progress and job separations. On the one hand, technology could be positively associated with job separations: when skills demanded by firms change, the latter might find it cheaper to hire new employees who are already familiar with the new technology than to retrain current workers. On the other hand, technological change may be negatively associated with involuntary job separations, because it may boost a firm's growth rate and technology-intensive firms may provide more training in specific skills than other firms. ZAVODNY's (2003) empirical results generally indicate a negative relationship between technology and job separation likelihood.

RODRIGUEZ and ZAVODNY (2003) discuss the theoretical effects of age and education on job displacements. Because older workers tend to be endowed with more specific human capital than younger workers, employers are less likely to lay off the former. However, older workers tend to have completed fewer years of education, and more educated workers are likely to have lower cost of or greater benefits from specific skills than less educated ones. The age-displacement relationship may shift adversely for older workers over time, because of increased demand for (young) skilled workers. Technological change may have created the need to retrain employees more often, and older workers are more costly to retrain because they are closer from retirement, leaving less time for firms to recoup their investment. Empirically, RODRIGUEZ and ZAVODNY (2003) find that the incidence of job displacement rose among workers aged 35 and older relative to younger workers, and among more educated workers relative to less educated workers (in contradiction with the skill-biased technological change hypothesis).

Finally, the decline in worker unions constitute a further explanation of decreasing job tenure. ABRAHAM and MEDOFF (1984) indeed provide evidence showing that unionized workers are better protected against permanent layoff than nonunionized ones, and figures for Switzerland display a steady decline in the number of unionized workers since 1980.

3. Data

This paper uses data from the Swiss Labor Force Survey (SLFS), which is carried out every year since 1991 by the Swiss Federal Statistical Office (SFSO). It contains detailed information about labor status, wages, and socioeconomic characteristics. The SLFS is a rotating panel, in which individuals remain (up to) five years in a row. Around 10,000 active persons were interviewed each year between 1991 and 2001. Since 2002, the SLFS was enlarged, and roughly 30,000 active persons are now interviewed each year. For our purposes, we restrict the sample to individuals aged 18 to 65 (62 for women), who are not self-employed. Further deleting all cases with missing information lead to final samples containing about 36,000 male workers (40,000 job spells) and 32,000 female workers (36,000 job spells).

Table 2 displays summary statistics of the covariates used in the analysis. Only one observation per individual is considered in the statistics, at the time of entry in the sample. There is no wide discrepancy between our sample and the active population in the Swiss labor market.

The central variable of our analysis is tenure. In the SLFS, the beginning of a job spell is obtained from responses to “In which year did you start working in this company?” and “Do you remember in which month?”. For all job spells ongoing at the time of an interview, we can thus retrieve the starting date and compute tenure on a monthly basis.

If an individual changed jobs twice between two interviews, we cannot take the “intermediate” job spell in account for two reasons: (i) if the individual moved directly from one job to the next, then no out of employment period is registered in the data and we are simply unable to retrieve the switching dates; (ii) even if we are able to identify a change because a nonemployment spell is observable, we have no information on the reason of the change nor on the termination reason of the “intermediate” job spell, which are crucial elements to our analysis.

Table 3 reports descriptive statistics on the number of observed spells in the final samples. For most individuals, we only observe one employment spell. Because of the structure of the SLFS, the maximum number of jobs that can be observed for a single individual is 4, even if each individual might be interviewed in 5 waves of the survey. Short job spells completed between two interview dates cannot be considered, and for statistical reasons, we have to discard spells started after the 4th interview and observed only at the 5th. Such spells are left-truncated and right-censored at the same point in time.

Analyzing tenure is tantamount to studying transitions out of employment, which can be performed efficiently using duration models (see Section 4). We

Table 2: Descriptive Statistics

Variable	Men		Women	
Age	40.63	(11.16)	39.75	(10.98)
Education level: low	0.15	(0.36)	0.21	(0.41)
Education level: medium	0.56	(0.50)	0.54	(0.50)
Education level: high	0.28	(0.45)	0.25	(0.43)
Foreign nationality	0.36	(0.48)	0.29	(0.45)
Number of children	0.70	(1.00)	0.60	(0.93)
Marital status: married	0.60	(0.49)	0.50	(0.50)
Part time employment	0.06	(0.23)	0.43	(0.50)
Firm size \geq 100 co-workers	0.40	(0.49)	0.32	(0.47)
Cantonal unemployment rate	3.28	(1.61)	3.30	(1.62)
Cantonal vacancy rate (deseasonalized)	0.27	(0.20)	0.27	(0.20)
Migration to population ratio (\times 1000)	5.59	(3.46)	5.57	(3.45)
# individuals	35,986		32,028	

Notes: Standard deviations in parentheses. For each individual appearing in the final sample, only the first observation is used.

Table 3: Number of Job Spells in Final Sample

# job spells	Men		Women	
	Absolute frequency ^a	Relative frequency ^b	Absolute frequency ^a	Relative frequency ^b
1	32,650	90.73	28,770	89.83
2	2,926	8.13	2,890	9.02
3	383	1.06	346	1.08
4	27	0.08	22	0.07
# observed individuals	35,986		32,028	
# observed spells	39,759		35,676	

a Number of individuals who experience the corresponding number of job spells.

b Ratio of the number of individuals who experience the corresponding number of job spells to the total number of individuals in the sample.

start our analysis by pooling all job spells in a single regression. We then refine our model and break down exits into several mutually exclusive destination states: *new job* when the individual transits directly from his old job to a new one (job-to-job move), *unemployment* when the individual suffers a period of unemployment after his job terminates, and *inactivity* if the labor market is left permanently for retirement or for no defined length.⁹

Beside this classification of destination states, we consider the motives for job terminations: separations initiated by the workers are labelled as *quits*, whereas those initiated by the firms are *layoffs*. A residual group, labelled *other reasons*, contains several possible exits that cannot be unambiguously classified as either worker or firm initiated separations.

3.1 Employment Tenure through Time

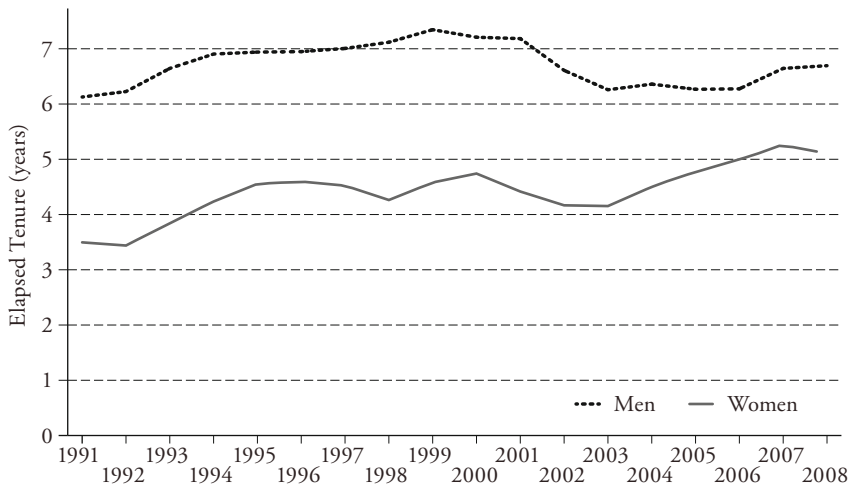
Figure 1 displays the evolution of median elapsed (incomplete) tenure for our final male and female samples, showing an expected gap between genders. With a difference of more than 3 years in the median tenure of men and women in 1991 but of less than 2 years in 2008, the spread however seems to be decreasing.

Median tenure is clearly countercyclical. It increased in the early 1990s when unemployment rose and GDP growth was weak. When the economy recovered in the late 1990s, median tenure decreased. Finally, after 2000, median tenure rose again (especially for women), mirroring the rise in unemployment. This countercyclical of median tenure is explained by the evolution of hires and separations along the business cycle (see GREGG and WADSWORTH, 1995, or AUER and CAZES, 2000).

Figure 2 displays the evolution of median elapsed tenure by age groups. Unsurprisingly, tenure rises with age. More interesting is the evolution of tenure over time: between 1991 and 2008, median tenure is more or less constant for the younger categories of men and women. However, median tenure is clearly on the decline for older men, with five years less in 2008 than in 1991 for workers over 45. For older women, conversely, median tenure seems on the rise. Although

9 Transitions to unemployment and inactivity are both job-to-nonemployment moves. We separate these two types of transitions as it makes the empirical analysis more sophisticated. If we kept them together, the broader destination state *job-to-nonemployment* would in fact reflect a weighted average of *unemployment* and *inactivity*. The unemployment category contains both registered and unregistered unemployed, while homemakers, retired and other inactive persons are considered inactive. Students are not classified in either category, but are ascribed to a separate destination state (*training*), whose estimations are not reported due to the very low number of transitions it encompasses.

Figure 1: Median Elapsed Tenure in Switzerland



Notes: Authors' calculations based on SLFS 1991-2008. Median elapsed tenure is computed on final samples and with sampling weights.

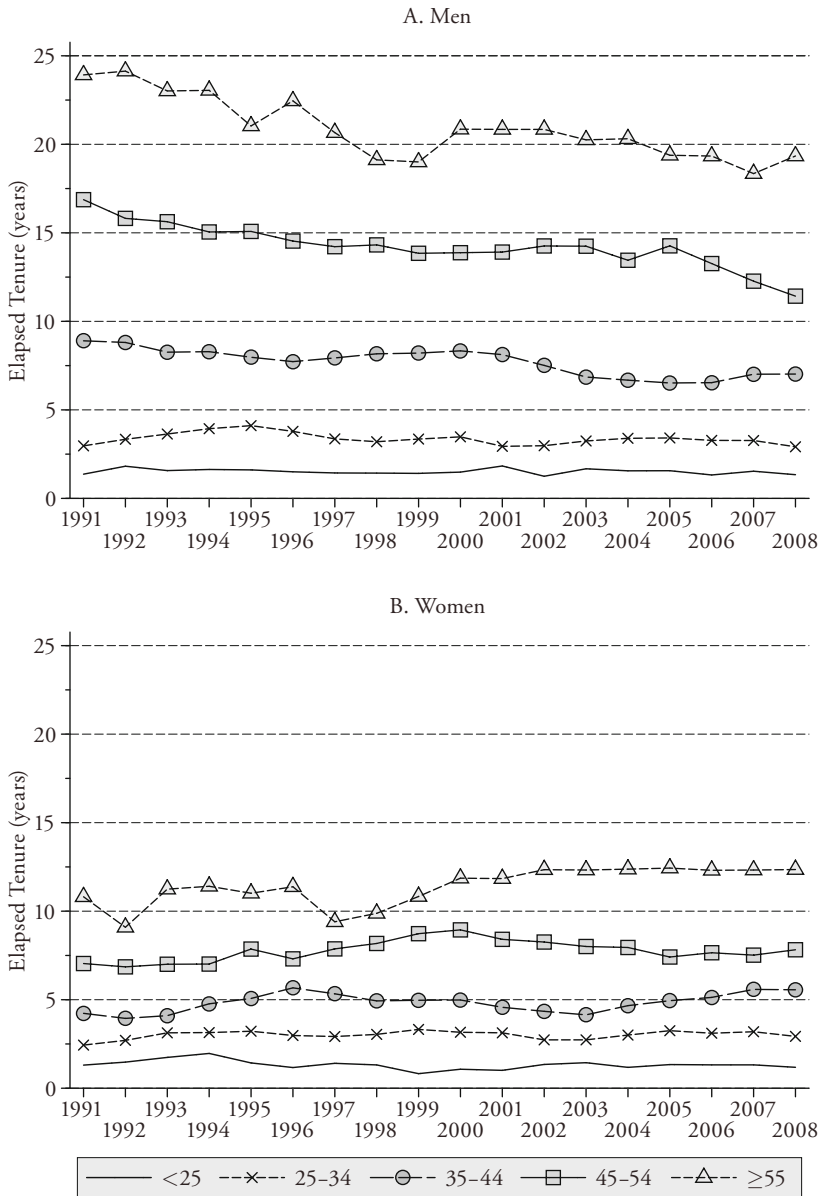
overall median tenure has not changed much during the last two decades, different groups of workers may have been affected more or less severely by reduced job stability.

4. Modeling Tenure Data

Some statistical issues must be addressed to handle panel tenure data efficiently. First, stock sampling gives rise to *left-truncation*: long job spells are over-represented in the sample. Second, most job spells are still ongoing when observation ends, generating *right-censoring*. Both issues are readily handled by duration analysis, which is therefore the most natural technique to analyze job tenure data.

Among different duration models, the semi-parametric Cox proportional hazards model (Cox, 1972, 1975) is appealing for our purposes. It leaves the baseline hazard unspecified and the duration dependence is therefore free to evolve non-monotonically over the job spell. The duration dependence of the hazard of

Figure 2: Median Elapsed Tenure, by Age Groups



Note: see Figure 1.

job termination is certainly non-monotonic and we do not want to impose any a priori restriction on its shape, which will be determined by the data alone.¹⁰

In what follows, we denote the duration of a job spell by t . $S(t)$ is the survivor function, $f(t) = dS(t) / dt$ is the probability density function, and $h(t) = f(t) / S(t)$ is the hazard function. The Cox model specifies the hazard function as:

$$h(t | x) = h_0(t) \cdot \exp(x' \beta) \quad (1)$$

where $h_0(t)$ is the baseline hazard function, x is a vector of characteristics, and β are the parameters of interest. This model is said to be semi-parametric since the baseline hazard function $h_0(t)$ is left unparameterized while the covariates enter the model log-linearly and multiplicatively.

Consider now individuals $i = 1, \dots, n$ with trivariate response $(t_{0i}; t_i; \delta_i)$, representing a period of observation $(t_{0i}; t_i]$, ending in either failure ($\delta_i = 1$) or right-censoring ($\delta_i = 0$). If individual i is known to fail (i.e., leave or lose his job) at time t_i , he contributes to the likelihood function the value of the density at time t_i , conditional on entry time t_{0i} , $f(t_i | x_i) / S(t_{0i} | x_i)$.

A right-censored observation, known to survive at least up to time t_i , contributes $S(t_i | x_i) / S(t_{0i} | x_i)$, which is the probability of surviving beyond time t_i conditional on entry time t_{0i} . Log-likelihood for the sample is thus:

$$\log L = \sum_{i=1}^n \delta_i \log h(t_i | x_i) + \log S(t_i | x_i) - \log S(t_{0i} | x_i) \quad (2)$$

For individuals under observation when their job spell starts, $t_{0i} = 0$ and $S(t_{0i} | x_i) = 1$. In our data though, virtually all job spells have already started when the individuals are interviewed. These spells are left-truncated, and the solution consists in analyzing only the part of the duration that comes after the date of interview (GUO, 1993). The period before the first interview and between two interviews in the case of a job change must not be considered as a period at risk since, had the job ended, we would never have known it. Starting dates, being asked retrospectively, allow to condition on time spent on the job but not at risk. With the exception of BERGEMANN and MERTENS (2011), this approach using information on all observed job spells has not been used to analyze job tenure.

10 Another possible choice would be the piece-wise exponential model. Results obtained with this alternative methodology are similar to those we report in Section 5 and are available on request. For a detailed presentation of duration models, see for example KIEFER (1988), LANCASTER (1990), BLOSSFELD and ROHWER (2002), or KALBFLEISCH and PRENTICE (2002).

Since we consider several possible exits from a job, competing risks models are used. The methodology is the same as the one just described, except that a specific hazard rate is specified for each mutually exclusive exit e :

$$h_e(t | x) = h_{0e}(t) \cdot \exp(x' \beta_e), \quad e = 1, \dots, m. \quad (3)$$

The overall log-likelihood of the model is then given by:

$$\log L = \sum_{e=1}^m \log L_e \quad (4)$$

The estimation of competing risks model is achieved by estimating a separate equation for each possible exit. For each exit-specific estimation, spells ending in a different exit than the one under study are considered as right-censored.¹¹

5. Results

This section investigates the evolution of job stability in Switzerland over the period 1991–2008. As a first step, we analyze overall job stability by pooling all job spells, ignoring destination states and reasons for job termination. Such regressions indicate how the probability of job termination is quantitatively affected by the covariates, but does not allow to assess whether workers' welfare is better or worse. The results are displayed in Table 4.

The coefficient of the year variable provides information on the trend of job stability. In estimations (1a) and (1b), it is slightly positive for both genders (but only significant for men), which implies an increasing hazard of job termination over time. Job spells have become slightly shorter between 1991 and 2008 and this points to some reduced job stability.^{12,13} This finding confirms SOUSA-POZA's (2004) results of no substantial changes in job stability in the 1990s in Switzerland.

- 11 For the sake of completeness, it should be mentioned that estimates for the Cox model are not, strictly speaking, obtained through maximizing the likelihood function. The reason is that it not only contains the β parameters but also the baseline hazard $h_0(t)$, which is unspecified. Instead, the likelihood function is broken down to exclude the baseline hazard, which makes it possible to maximize a partial likelihood function (Cox, 1972, 1975). BRESLOW's (1974) method is used to handle duration ties.
- 12 Quantitatively speaking, the point estimate of 0.007 in estimation (1a) indicates that the hazard rate of job termination has increased at an average rate of 0.7% (that is, precisely:

Coefficients estimated for age groups indicate that job stability is an inverted U-shaped function of age. Individuals between 45 and 55 are found to have the most stable employment relationships. This result was expected as those individuals are the most likely to have family duties, so that changing job would be risky for them. At the same time, these individuals have acquired a substantial amount of specific human capital, which makes employers reluctant to dismiss them. A similar structure of age group coefficients can also be found in HIRSCH and SCHNABEL (2010) and BERGEMANN and MERTENS (2011).

The hazard functions based on estimations (1a) and (1b) of Table 4 are plotted in Figure 3, revealing the duration dependence of the risk of job termination through a job spell. Both hazard functions are drawn for the mode of the overall sample covariates distribution, so as to allow for meaningful comparisons across genders. Time axis is cut at 30 years of tenure because longer job spells are scarce and trajectories fluctuate strongly afterwards.

Figure 3 first confirms that women have shorter job spells than men, their hazard rate being always larger (even though characteristics are held constant across the two groups). The difference however is weakly significant, with confidence intervals sometimes overlapping.

The duration dependence of the job termination hazard peaks within the first few years of a job spell, then decreases until 10 years of tenure, and becomes virtually constant thereafter. Beyond 30 years of tenure (not shown in the graph), the hazard rate would eventually increase very sharply, because jobs end mechanically as workers reach retirement age. This shape of the hazard function is perfectly consistent with several other studies estimating duration models for tenure (BOOTH, FRANCESCO, and GARCIA-SERRANO, 1999; MARINESCU, 2009; HIRSCH and SCHNABEL, 2010; BERGEMANN and MERTENS, 2011). It indicates that job relationships have a high risk of ending early, and many will last no more than a few years or even a few months, but those exceeding some years

$e^{0.007} - 1$) per year over the period 1991–2008, everything else being constant. The effect on expected survival time cannot however be computed in a simple fashion, as it depends on the value of all covariates. In our case, for a worker with characteristics at the mode of all covariates, this translates into an expected median job spell that decreased by 6 months between 1991 and 2008, from 3 years and 11 months to 3 years and 5 months.

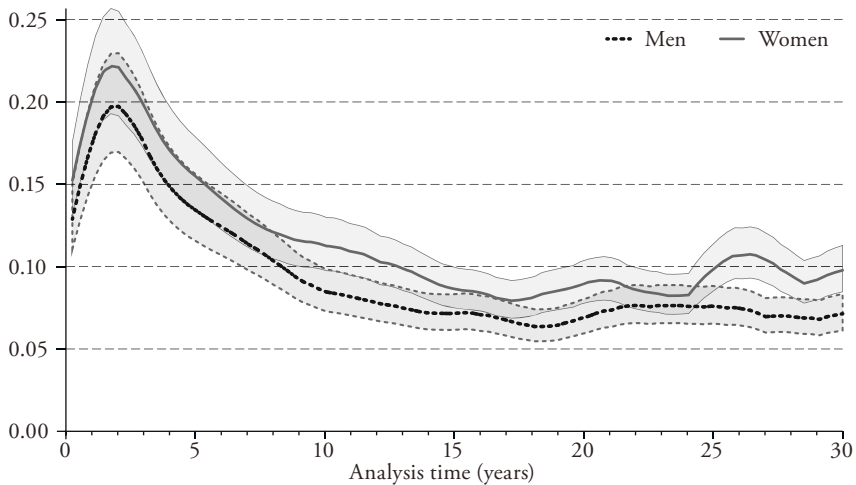
- 13 To account for potential non-linearities, we re-estimated the model by replacing the continuous trend variable by year fixed effects. The coefficients estimated for the year fixed effects are displayed in Appendix Figure A.1. No clear pattern emerges, with increases and decreases over the observation period, and few of these coefficients being significantly different from zero. As a whole, only a blurred image crops up from the results on the evolution of job stability in Switzerland.

Table 4: Hazard of Job Termination, by Gender

	Men		Women	
	(1a)	(2a)	(1b)	(2b)
Year	0.007 ^{***} (0.002)	—	0.003 (0.002)	—
Age 25–35 years	-0.302 ^{***} (0.035)	-0.306 ^{***} (0.036)	-0.302 ^{***} (0.032)	-0.281 ^{***} (0.033)
Age 35–45 years	-0.541 ^{***} (0.039)	-0.553 ^{***} (0.042)	-0.587 ^{***} (0.035)	-0.587 ^{***} (0.038)
Age 45–55 years	-0.723 ^{***} (0.046)	-0.805 ^{***} (0.051)	-0.808 ^{***} (0.041)	-0.816 ^{***} (0.044)
Age > 55 years	-0.086 [*] (0.044)	-0.136 ^{***} (0.048)	-0.459 ^{***} (0.043)	-0.427 ^{***} (0.045)
Year × Age 18–25	—	-0.003 (0.006)	—	0.014 ^{***} (0.005)
Year × Age 25–35	—	-0.000 (0.004)	—	-0.003 (0.003)
Year × Age 35–45	—	0.006 (0.004)	—	0.007 (0.004)
Year × Age 45–55	—	0.028 ^{***} (0.006)	—	0.010 [*] (0.006)
Year × Age > 55	—	0.017 ^{***} (0.005)	—	-0.005 (0.006)
# spells	39,759	39,759	35,676	35,676
# individuals	35,986	35,986	32,028	32,028
# failures	10,060	10,060	10,456	10,456
LogL	-81,430	-81,419	-86,754	-86,748
AIC	162,971	162,956	173,618	173,614
BIC	163,621	163,654	174,260	174,302

Notes: Data: Swiss Labor Force Survey, 1991–2008. Robust clustered standard errors in parentheses. ^{***}/^{**}/^{*}: Significant at the 0.01/0.05/0.10 level. Base category for age groups: 18–25 years. Unreported covariates: education groups, foreign nationality, marital status, number of children, part-time work, firm with more than 100 co-workers, cantonal unemployment rate, cantonal vacancy rate, canton dummies, sector dummies, cantonal migration to population ratio. Complete results are available on request.

Figure 3: Hazard Rate of Job Termination, by Gender



Notes: Hazard rates drawn for the mode of the covariates distribution, i.e., for a Swiss individual aged 40, married, without children, with medium education, working full-time in the manufacturing sector, in a firm with less than 100 co-workers, in the canton of Zürich, in 2007 (last complete year under observation). Shaded areas are 90% confidence intervals.

become very unlikely to be dissolved. This pattern also matches perfectly with central facts about job mobility stated by FARBER (1999): 1) long-term relationships are common, 2) most new jobs end early, and 3) the probability of a job ending declines with tenure.

Such a duration dependence of the hazard rate can be explained if jobs are “experience goods”.¹⁴ In this case, JOVANOVIĆ (1979) shows that it pays to remain and collect information on a new job. Before dissolving their match, both worker and employer must accumulate some critical amount of information to determine whether it is worth maintaining their employment relationship. This results in a separation rate that rises at the beginning of a job, and drops afterwards.

Even if job stability did not clearly decrease for the entire active population, it could be the case for some specific groups, as hinted by Figure 2 in Section 3. To explore this issue, we add interactions terms between the year and age groups in

14 A job is an “experience good” if its characteristics are difficult to observe ex ante. The only way to assess the quality of a particular job match is to form the match and “experience it”.

estimations (2a) and (2b) of Table 4. We find that employment has become significantly less stable for male workers over 45 years old. Among women, however, it appears that workers under 25 suffer a higher risk of job termination than in the past, whereas older female workers have seen less change in their hazard rate.

In order to better grasp the determinants of job stability and its evolution, we now split estimations by destination state (new job, unemployment, or inactivity) and by termination reason (quit, layoff, or other reasons).

The results of Table 5 show that, as expected, job-to-job mobility decreases with age for both genders.¹⁵ Towards unemployment, hazard rates are higher for men over 55, but no significant pattern is observed among women. Towards inactivity, the hazard is a U-shaped function of age, being low for middle-aged individuals and increasing at older ages, certainly under the effect of retirement. Unsurprisingly, we also find a significantly larger transition rate towards inactivity for women between 25 and 35, which corresponds to the period where most of them give birth to their first child.¹⁶

These results are largely consistent with those provided by HIRSCH and SCHNABEL (2010). Their *employment* state is directly comparable to *new job* here, and their *non-employment* can be considered as a merger of our *unemployment* and *inactivity*. HIRSCH and SCHNABEL (2010) find a clearly negative effect of age on the separation rate towards employment but a separation rate towards non-employment that decreases at young ages and increases at old ages.

Over time, these hazard functions seem to have evolved differently. The intensity of transitions to a new job remained relatively stable. If anything, it increased for male workers between 45 and 55 and decreased for female workers between 25 and 35. For both genders, transitions towards unemployment have become more frequent for the youngest age groups, and to a lesser extent for the oldest age groups. For middle-aged workers, unemployment risk does not seem to have changed. Finally, as regards transitions towards inactivity, it appears that the hazard rate has significantly increased for men older than 35 and women aged 35 to 45, but it decreased for women over 55.

Summing up the broad picture of our results in Table 5, one is tempted to conclude that occurrences of adverse labor market events have risen for the youngest workers, but less so for all other workers. A possible explanation for the decline in the median tenure observed for male workers over 45 is that a growing share

15 For the sake of space, only estimations encompassing interaction terms between year and age groups are displayed. Results without interaction terms are available from the authors.

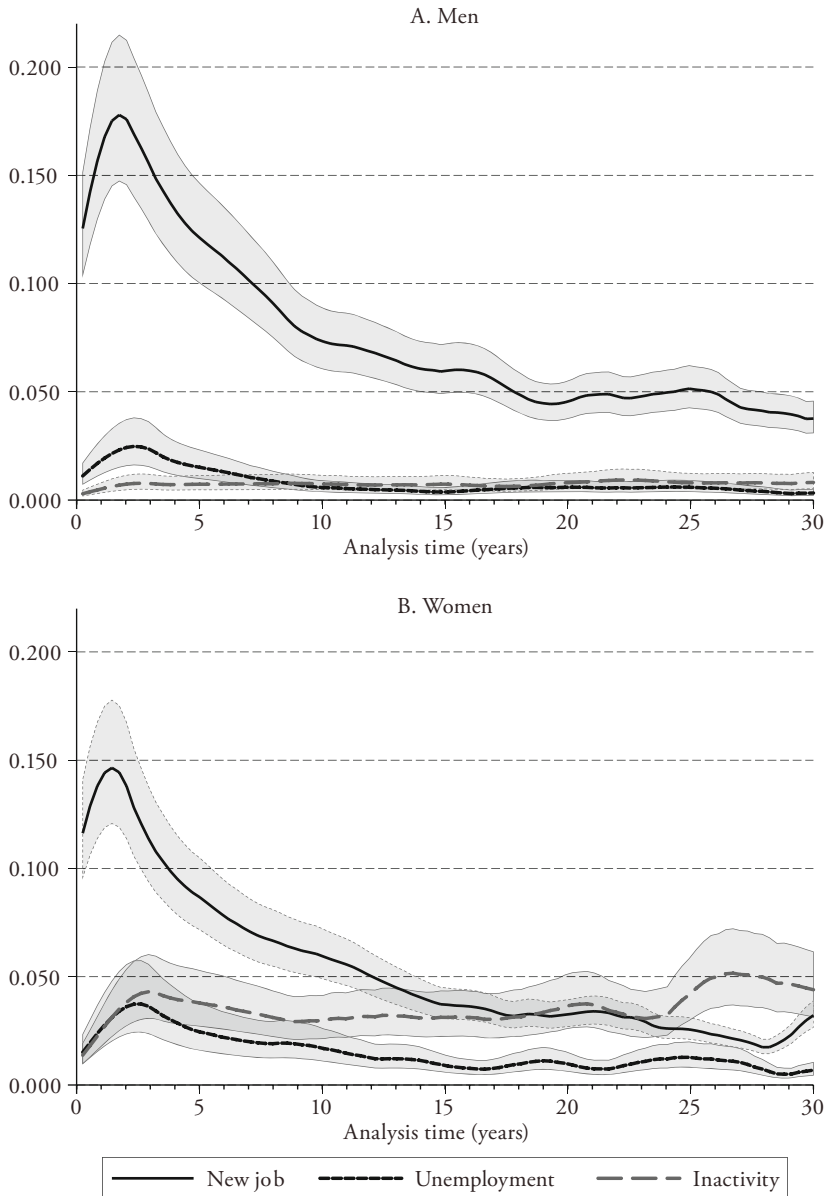
16 Additional estimations accounting for the introduction of compulsory maternity leave insurance in 2005 led to unaltered results.

**Table 5: Competing Risks Model for the Hazard of Job Termination,
by Destination State**

	Men			Women		
	(1a) New job	(2a) Unem- ployment	(3a) Inactivity	(1b) New job	(2b) Unem- ployment	(3b) Inactivity
Age 25–35 years	-0.148*** (0.045)	-0.189 (0.116)	-0.858*** (0.179)	-0.227*** (0.041)	-0.049 (0.125)	0.387*** (0.111)
Age 35–45 years	-0.394*** (0.051)	-0.051 (0.128)	-0.976*** (0.193)	-0.418*** (0.047)	0.047 (0.131)	-0.306** (0.120)
Age 45–55 years	-0.710*** (0.062)	0.197 (0.133)	-0.708*** (0.191)	-0.707*** (0.057)	0.182 (0.133)	-0.412*** (0.124)
Age > 55 years	-1.057*** (0.084)	0.501*** (0.145)	1.525*** (0.158)	-1.161*** (0.084)	0.120 (0.168)	0.852*** (0.120)
Year × Age 18–25	-0.009 (0.007)	0.034 [†] (0.019)	-0.019 (0.029)	0.002 (0.006)	0.079*** (0.018)	0.006 (0.018)
Year × Age 25–35	-0.000 (0.004)	-0.002 (0.011)	0.031 (0.019)	-0.008 [†] (0.004)	0.049*** (0.012)	-0.009 (0.007)
Year × Age 35–45	0.003 (0.005)	0.007 (0.013)	0.055*** (0.020)	0.001 (0.005)	0.010 (0.012)	0.025*** (0.009)
Year × Age 45–55	0.017** (0.008)	0.013 (0.014)	0.089*** (0.019)	0.008 (0.008)	0.012 (0.013)	0.011 (0.010)
Year × Age > 55	-0.010 (0.012)	0.023 (0.017)	0.018*** (0.006)	0.002 (0.013)	0.071*** (0.022)	-0.028*** (0.007)
# spells	39,759	39,759	39,759	35,676	35,676	35,676
# individuals	35,986	35,986	35,986	32,028	32,028	32,028
# failures	6,590	1,320	1,801	6,018	1,325	2,744
LogL	-54,683	-10,770	-11,799	-50,625	-10,907	-21,156
AIC	109,484	21,657	23,715	101,369	21,929	42,430
BIC	110,182	22,355	24,413	102,057	22,594	43,118

Notes: Data: Swiss Labor Force Survey, 1991–2008. Robust clustered standard errors in parentheses. ***/**/*: Significant at the 0.01/0.05/0.10 level. Base category for age groups: 18–25 years. Unreported covariates: education groups, foreign nationality, marital status, number of children, part-time work, firm with more than 100 co-workers, cantonal unemployment rate, cantonal vacancy rate, canton dummies, sector dummies, cantonal migration to population ratio. The total number of failures does not correspond to what appears in Table 4, because estimations for the destination state “training” are not reported (349 transitions for men and 369 for women).

Figure 4: Hazard Rates of Job Termination, by Destination State



Note: see Figure 3.

of them has taken up early retirement schemes, as reflected in the coefficients on transitions from employment to inactivity. The results for female workers are in line with their observed increasing participation rate and increasing attachment to the labor market.

Figure 4 draws the hazard functions (evaluated at the mode of the covariates) towards the different destination states. The hazard rate for new jobs is by far the largest, at least for men. It peaks during the first few years of a job spell and then decreases, sharply up to 10 years of tenure and more gradually thereafter. Transition rates towards unemployment are comparatively low and globally decrease. The most noticeable difference between men and women is found for hazard rates towards inactivity: the hazard rate is virtually nil for men while it is sizeable for women. Further, there is no apparent duration dependence for this destination state. Again, other papers in the literature display very similar hazard rates, in terms of both levels and duration dependence (see HIRSCH and SCHNABEL, 2010; BERGEMANN and MERTENS, 2011).

In Table 6, exits are separated according to termination reasons: quits, layoffs, and other reasons. As in the previous set of estimations, the coefficients for all age groups are in line with expectations. Older workers quit much less frequently. As workers age, they do not want to take the risk of changing jobs, because they likely have family and children they need to take care of. At the same time, older workers are also more likely to be laid off. These findings are very similar to those in MARINESCU (2009) and BERGEMANN and MERTENS (2011), who consider the effects of age on layoff probabilities.¹⁷

Only few of the interaction terms are significant in this set of estimations, maybe because of the shorter observation period on which it is based.¹⁸ The quit rate has decreased among young female workers, which is consistent with the increase in the average age of women at the birth of their first child. Among men, even though coefficients cannot be precisely estimated, it seems that the hazard of quit declined among the young whereas it increased for the older. The contrary seems true for the layoff hazard, which might be on the rise for young male workers, as their coefficient is positive and on the edge of significance. For all women, the layoff rate is increasing, but only significantly so for the oldest group.

Figure 5 displays the hazard rates for the different termination reasons. The largest transition risk is found for quits as in BERGEMANN and MERTENS (2011). This hazard rate is of a similar magnitude for men and women, and it once again

17 The coefficients on age groups are not displayed in Ioanna Marinescu's published paper. We thank her for kindly providing us with complete results tables.

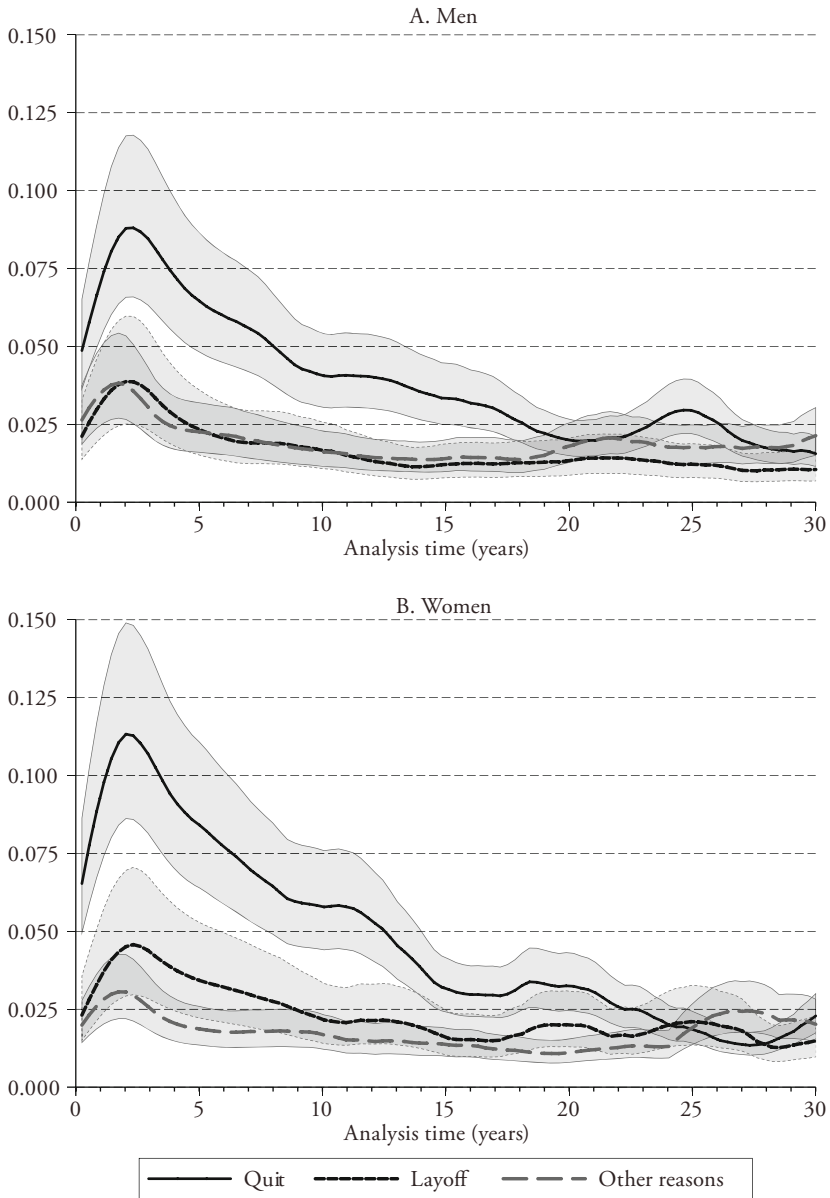
18 The SLFS provides information on job termination reasons only since 1996.

**Table 6: Competing Risks Model for the Hazard of Job Termination,
by Termination Reason**

	Men			Women		
	(1a) Quit	(2a) Layoff	(3a) Other reasons	(1b) Quit	(2b) Layoff	(3b) Other reasons
Age 25–35 years	-0.198** (0.091)	0.153 (0.193)	-0.357*** (0.117)	-0.511*** (0.086)	-0.104 (0.175)	-0.233** (0.094)
Age 35–45 years	-0.666*** (0.103)	0.483** (0.196)	-0.730*** (0.132)	-0.784*** (0.097)	0.131 (0.179)	-0.763*** (0.107)
Age 45–55 years	-1.220*** (0.144)	0.699*** (0.200)	-1.240*** (0.166)	-1.261*** (0.118)	0.436** (0.180)	-1.187*** (0.129)
Age > 55 years	-2.154*** (0.233)	0.855*** (0.214)	0.249** (0.127)	-1.806*** (0.187)	0.410 [†] (0.209)	-0.021 (0.117)
Year × Age 18–25	-0.015 (0.019)	0.047 (0.034)	-0.031 (0.022)	-0.042*** (0.016)	0.018 (0.032)	0.020 (0.018)
Year × Age 25–35	-0.018 (0.011)	-0.001 (0.018)	-0.036** (0.014)	-0.008 (0.012)	0.027 (0.020)	-0.012 (0.013)
Year × Age 35–45	0.004 (0.013)	-0.024 (0.017)	-0.019 (0.016)	-0.004 (0.014)	0.007 (0.019)	0.004 (0.015)
Year × Age 45–55	0.029 (0.022)	-0.004 (0.018)	0.034 (0.023)	-0.000 (0.018)	0.024 (0.018)	-0.003 (0.021)
Year × Age > 55	0.052 (0.038)	0.019 (0.022)	0.029** (0.013)	0.014 (0.032)	0.049 [†] (0.026)	-0.025 (0.016)
# spells	34,271	34,271	34,271	30,629	30,629	30,629
# individuals	31,283	31,283	31,283	27,767	27,767	27,767
# failures	3,150	1,766	2,572	3,215	1,548	2,616
LogL	-25,355	-13,736	-19,206	-26,202	-12,174	-20,731
AIC	50,828	27,591	38,531	52,521	24,467	41,580
BIC	51,514	28,276	39,217	53,198	25,144	42,257

Notes: Data: Swiss Labor Force Survey, 1996–2008. Robust clustered standard errors in parentheses. ***/**/*: Significant at the 0.01/0.05/0.10 level. Base category for age groups: 18–25 years. Unreported covariates: education groups, foreign nationality, marital status, number of children, part-time work, firm with more than 100 co-workers, cantonal unemployment rate, cantonal vacancy rate, canton dummies, sector dummies, cantonal migration to population ratio. Number of observations are smaller than in Tables 4 and 5 because information about job termination reason is only available since 1996.

Figure 5: Hazard Rates of Job Termination, by Termination Reason



Note: see Figure 3.

exhibits a peak early in the job spell and a continuous decrease thereafter. The layoff hazard seems slightly higher for women than for men (even though the difference is not statistically significant).

As a robustness check, we finally estimate all our models separately by cohorts.¹⁹ This alternative approach allows a comparison of trends for older and younger cohorts, which also captures the idea of changes in lifetime jobs. The results are displayed in Table A.1 of the appendix. Even though statistical significance is generally low, the conclusions of our previous analyses are confirmed. Hazard rates towards unemployment and layoff increased mostly for the youngest cohorts (but only significantly so for female workers towards unemployment). The older cohorts of women were adversely affected as well. Among the older male cohorts however, only transitions towards inactivity did significantly increase.

6. Conclusions

This paper investigates the evolution of job stability through the estimation of a series of Cox proportional hazards models on tenure. Job stability has been extensively studied in the literature, though not often with the appropriate econometric tools. In line with recent literature, we make use of competing risks models, allowing individuals to move towards several destination states and jobs to terminate for various reasons.

Our estimates do not show any clear tendency towards either a general increase or decrease in the duration of job spells. For male workers over 45, we do however find a decrease in job stability. In any case, shorter job spells do not necessarily indicate a worsening situation for the workers. If they decide to change jobs more often, job stability would indeed decrease but that would not necessarily prove that their welfare has diminished.

In refined sets of estimations, workers' destination states (new job, unemployment, or inactivity) and reasons for job termination (quits, layoffs, or other reasons) are taken into account. Our results then point to a worsening situation for the youngest workers, and to a lesser extent for workers over 55. Unemployment and layoff hazards have both increased for workers below 25. These negative outcomes seem to have increased for older women as well. For older male workers, however, no such evolution is observed, and their decreasing job stability is explained by a larger transition rate towards inactivity, which is at least

19 We are grateful to an anonymous referee for this suggestion.

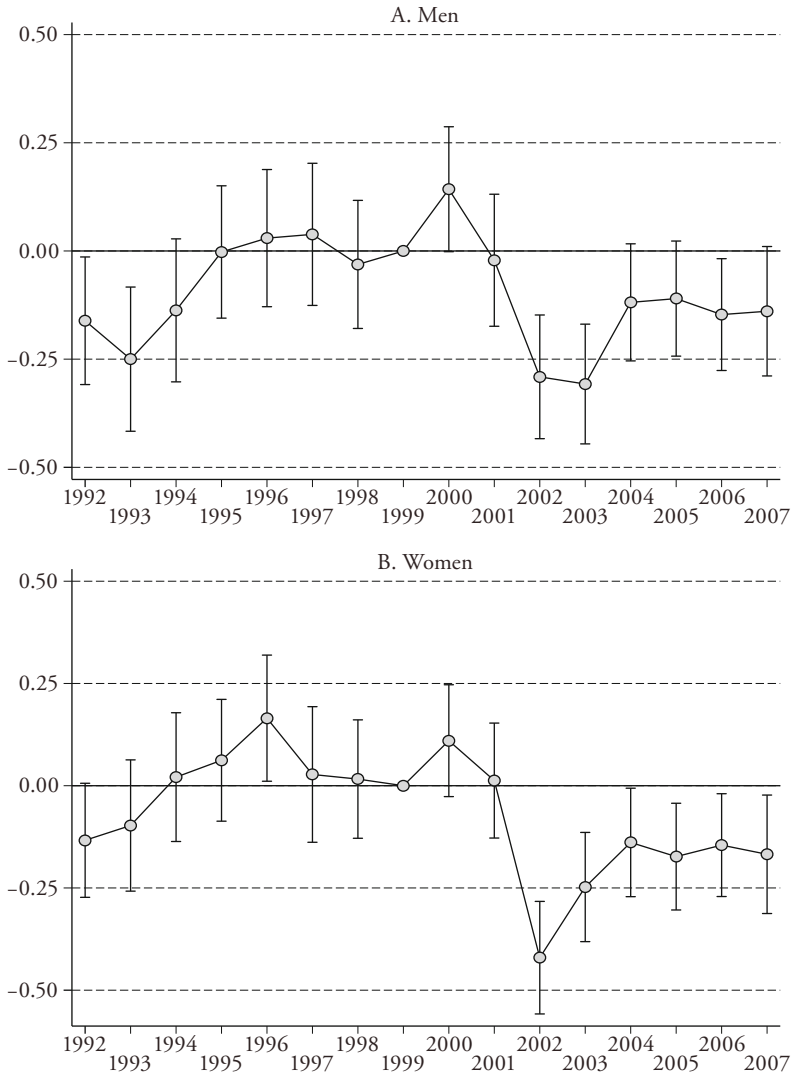
partly explained by the increasing number of early retirements. Even if some of the moves towards inactivity might be considered as disguised layoffs, our data-set does not enable us to provide a more precise assessment.

Finally, let us mention that it would be interesting to combine both destination states and termination reasons, since they are complementary decompositions. One could indeed imagine that the probabilities of quitting towards a new job and quitting towards unemployment or inactivity are different. The transition rate from a quit to a new job could also differ from the transition rate from a layoff to a new job. Our data set does unfortunately not contain enough observations to robustly estimate such models.

What can be said as to the causes of these changes? The major labor market shock that has hit Switzerland during our observation period originates from the complete removal of barriers to migration flows from EU countries in 1999. One can therefore not exclude that the stronger competition on the supply side has reduced the security of the younger cohorts, without affecting much the rate of dismissal of older generations (as long as early retirement schemes were not forced). Employers may have gradually shifted to more conservative policies in terms of “lifetime employment promises” since they can rely on a larger pool of applicants. Older workers, on the other hand, would still benefit from past promises or be granted generous retirement schemes. Such a conclusion is only indicative at best and does not rely on a direct test of causality between the corresponding events. Confounding factors like technological change, demographic changes or the shift from manufacturing to services could also bring about similar results, and additional research is required to test these assumptions.

Appendix

Figure A.1: Estimated Coefficients for the Evolution of the Hazard of Job Termination



Notes: The coefficients are from an estimation similar to (1a) and (1b) in Table 4, where the variable year has been replaced by time fixed effects with 1999 as a reference. Whiskers are 95% confidence intervals.

Table A.1: Coefficients on Year, by Cohort

	Men			
	Cohort 1	Cohort 2	Cohort 3	Cohort 4
All destinations	-0.004 (0.005)	0.011 (0.010)	0.012 [†] (0.007)	-0.011 [†] (0.007)
New job	-0.012 (0.011)	0.004 (0.012)	0.010 (0.008)	-0.018 ^{**} (0.008)
Unemployment	0.023 (0.016)	0.027 (0.025)	0.018 (0.020)	0.021 (0.022)
Inactivity	0.026 ^{***} (0.007)	0.078 ^{**} (0.034)	0.015 (0.029)	-0.039 (0.030)
# spells 1991–2008	8,049	9,148	13,463	9,099
# individuals 1991–2008	7,733	8,478	11,851	7,924
Quit	-0.027 (0.053)	-0.010 (0.033)	-0.008 (0.018)	-0.003 (0.015)
Layoff	-0.001 (0.029)	-0.032 (0.031)	-0.004 (0.026)	0.025 (0.028)
Other reasons	0.022 (0.015)	0.012 (0.035)	-0.051 ^{**} (0.026)	-0.024 (0.021)
# spells 1996–2008	6,130	7,647	11,669	8,825
# individuals 1996–2008	5,929	7,175	10,452	7,727
	Women			
	Cohort 1	Cohort 2	Cohort 3	Cohort 4
All destinations	-0.014 ^{**} (0.006)	0.009 (0.009)	0.004 (0.007)	0.009 (0.006)
New job	-0.002 (0.012)	-0.005 (0.013)	-0.001 (0.009)	-0.004 (0.007)
Unemployment	0.058 ^{***} (0.020)	0.044 [†] (0.025)	0.019 (0.023)	0.065 ^{***} (0.019)
Inactivity	-0.025 ^{***} (0.007)	0.011 (0.019)	0.013 (0.015)	0.053 ^{***} (0.020)
# spells 1991–2008	6,717	8,079	11,682	9,198
# individuals 1991–2008	6,367	7,406	10,452	7,727

	Women			
	Cohort 1	Cohort 2	Cohort 3	Cohort 4
Quit	0.028 (0.042)	-0.055* (0.029)	-0.011 (0.019)	0.010 (0.015)
Layoff	0.070** (0.032)	0.052* (0.031)	-0.019 (0.030)	0.040 (0.030)
Other reasons	-0.059** (0.020)	0.006 (0.035)	-0.014 (0.023)	-0.001 (0.017)
# spells 1996–2008	4,995	6,824	9,989	8,821
# individuals 1996–2008	4,787	6,327	8,954	7,699

Notes: Data: Swiss Labor Force Survey. Robust clustered standard errors in parentheses. ***/**/*: Significant at the 0.01/0.05/0.10 level.

Each number in this table is the coefficient on year obtained in an estimation similar to (1a) and (1b) in Table 4, conducted separately on each cohort, and including age as a continuous control variable.

Cohort 1: individuals born before 1950. Cohort 2: individuals born 1950–1959. Cohort 3: individuals born 1960–1969. Cohort 4: individuals born 1970 or later.

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SUMMARY

Using data over 1991–2008 for Switzerland, we investigate job stability through a series of Cox proportional hazards models. Our baseline results show that employment has become less stable for older male workers, with less noticeable change for other groups. However, when destination states are considered in the model, results indicate that younger workers face more transitions towards unemployment than before, whereas older male workers' greater instability is caused by an increase in transitions to inactivity. It thus appears that the situation of young workers has deteriorated, while the evolution of older men's job stability is at least partly explained by the increasing number of early retirements. For women, our results are largely consistent with their increasing participation rate and attachment to the labor market.