From Lifetime Jobs to Churning?*

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Abstract

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 workers, with no noticeable change for younger ones. However, when destination states are included in the model, our results indicate that younger workers face more transitions towards unemployment than before, whereas older workers' greater instability is caused by an increase in transitions to inactivity. Finally, when measures of occupational skill intensity are included, our results provide supporting evidence for the skill-biased technological change hypothesis.

Keywords: Job stability, Job security, Tenure, Lifetime jobs, Duration models, Switzerland.

JEL Classification: J63, J62, C41.

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1 Introduction

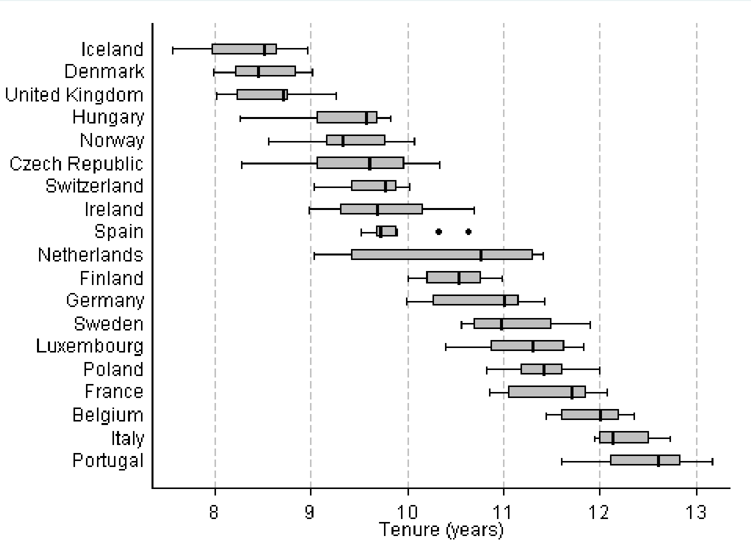
It is often taken for granted that lifetime jobs have become scarcer than in the past. Definite proofs of this popular belief are however hard to come by, and the literature provides only little evidence in its favor. In Switzerland, the death of the "job for life" paradigm has received important coverage in the media since the late 1990s, although the picture was somewhat exaggerated by single but large and visible events like the privatization of the telecommunications sector or the bankruptcy of the Swiss national airlines. 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 stability is justified, and if so, why. A previous study of job stability in Switzerland is provided by Sousa-Poza (2004), but it is essentially descriptive. Here, we make use of duration analysis, which is the most relevant econometric technique in this context, and look for the reasons of variations in job stability.

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 *et al.*, 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) have shown that increased imports of finished goods are negatively associated with unskilled labor. However, countries might tend to increase labor market security in order to cushion the shocks caused by greater openness (Rodrik, 1998).

Switzerland clearly belongs to the league of countries with moderate employment protection, its OECD EPL index being one of the lowest in the world.[1] 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

[1]Employment protection legislation (EPL) is discussed in OECD (2004). See `http://www.oecd.org/employment/protection` for updated measures.

Figure 1: Distribution of annual average tenure for some European countries, 1996-2010



Notes: Data come from OECD.StatExtracts and are based on the European Labour Force Survey. Annual average tenure is for all persons (men and women) in dependent employment or self-employed. Years 1996-2010 are retained because Swiss data is available over this period only. Countries are included only if at least 10 yearly observations are available and are ranked by overall (unweighted) average tenure over the years 1996-2010.

in Switzerland.

The Swiss labor market is also of particular interest because of important changes that took place in the last 2 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) and it cannot be excluded that the new law has helped in this regard.

Figure 1 shows the distribution of average tenure for some European countries. With an average tenure between 9 and 10 years over the period 1996-2010, Switzerland lies slightly below the average of all European countries. In terms of variability in tenure, Switzerland appears as a typical country. Comparisons based on OECD (1997) would lead to similar remarks.

While the literature mostly focuses on job (in)stability, job (in)security is a distinct concept, though an implicitly quite relevant one. 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 clearly 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 still offer some insights. In our estimations, we use competing risks duration models to break down employment spells across destination states following their termination (new job, unemployment or inactivity) and the events that caused the latter (quits or layoffs), along the lines of Booth *et al.* (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. In fact, the discrepancy between the popular feeling of growing job instability and the lack of evidence observed in the statistics might arise from a relative stability in the overall population, but variations that compensate between subpopulations. In particular, if the situation of young workers improves to the detriment of older workers, this might well be felt as growing job instability, given that a job loss has serious consequences beyond a certain age. Every worker is bound to grow older, so that he might anticipate potential troubles regardless of his current age. In order to test for this hypothesis, we provide estimations that separate workers across age groups.

The remainder of the paper is organized 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

The literature on job stability and lifetime jobs finds its origins in the seminal paper by Hall (1982) and in Ureta's (1992) extensions. Thereafter, the literature on job stability has produced controversial results. Table 1 provides some characteristics of a number of studies that investigate the evolution of job stability.[2]

Admittedly, it is not easy to discern a general consensus. The evolution of job stability has been particularly debated in the US, and results for European countries are not clearcut either. Moreover, we only report the findings on overall job stability in the population studied. The variety of results is much larger when considering subpopulations within each study.

Table 1 highlights how studies making use of duration models are still a minority so far. Duration analysis is however the most adequate technique to analyze job spells, since it readily handles right-censoring and left-truncation. These two difficulties are left out 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 might last for many additional years (right-censoring) and that stock sample survey suffer from an over-representation of long job spells (left-truncation).[3]

In this paper, we use duration models to investigate the evolution of job stability in Switzerland. Even though job stability has already been studied in this country (Sousa-Poza, 2004), we provide a more careful analysis by splitting 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-non employment-job moves (Jung and Winkelmann, 1993; Royalty, 1998). We also account for right-censoring and left-truncation, so that we

[2]Gottschalk and Moffitt (1999) display a similar table focusing on US studies.

[3]A simple 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.

are able to retrieve information from all job spells observed in our survey, whether or not they are already ongoing at the beginning of observation, and whether or not they are completed by the end of observation. As some previous papers use a similar methodology, our work might be considered as scientific replication in the sense of Hamermesh (2007). 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. Other papers estimating duration models either do not account for left-truncation Booth *et al.* (1999); Gottschalk and Moffitt (1999), or do not study the evolution of job stability but compare different socio-economic groups (Marinescu, 2009; Hirsch and Schnabel, 2010; Rokkanen and Uusitalo, 2010).

Table 1: Comparison of studies on 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 *et al.* (1996, 1997)	US	CPS supplements 1983-1991	4 and 10-year retention rates	No clear trend
Boisjoly *et al.* (1998)	US	PSID 1968-1992	Annual rate of involuntary job loss	Increase in involuntary job losses
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≤1), Pr(T<10)	Increase in Pr(T<10) but no trend in Pr(T≤1)
Farber (2009)	US	CPS supplements 1973-2008 + DWS 1983-2008	OLS on elapsed tenure, Pr(T≥10), Pr(T≥20), Pr(T<1)	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≥5)	No clear trend
Booth *et al.* (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
Mahringer (2004)	Austria	ASSR 1993-2003	1- to 3-year retention rates	No clear trend
Givord and Maurin (2004)	France	FLFS 1982-2002	Job loss rates	Increase in the risk of involuntary job loss
Sousa-Poza (2004)	Switzerland	SLFS 1991-2001	1-year retention rates	No clear trend
Bratberg *et al.* (2010)	Norway	SNR 1986-2002	Pr(T<1), Pr(T≥8), job separation rates	Very weak indications that job stability has decreased
Rokkanen and Uusitalo (2010)	Finland	FPIS 1963-2004	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, FPIS = Finnish Pension Insurance Scheme, FLFS = French Labor Force Survey, GHS = General Household Survey, GSOEP = German Socio-Economic Panel, PSID = Panel Study for Income Dynamics, SIPP = Survey of Income and Program Participation, SLFS = Swiss Labor Force Survey, SNR = Statistics Norway Registers.

3 Data

We use 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 37,000 men (43,000 job spells) and 33,000 women (40,000 job spells).

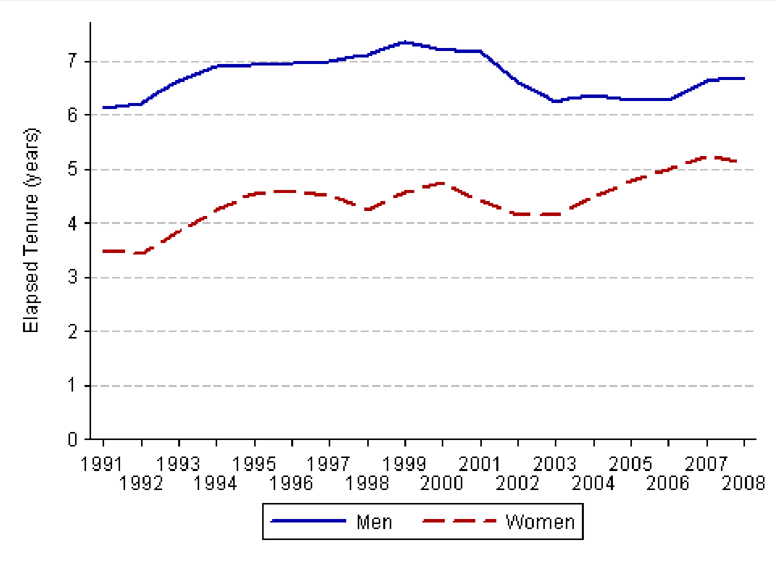
The central variable of our analysis is tenure. In the SLFS, the beginning of a job spell is obtained from responses to "In which year did you start working in this company?" and "Do you remember in which month?". In addition, the respondent's labor status is provided for each of the 12 previous months if he changed or left his job over this period. Both starting and ending dates of a job spell are thus known to fall in a given month, so that tenure can be computed on a monthly basis.

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

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

[4] 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". In concrete terms, the distinction between unemployment and inactivity is based on the individual declaration at the time of interview. The unemployment category contains both registered and unregistered unemployed, while homemakers, retired and other inactive persons are considered inactive. Students are not classed 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.

Figure 2: 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.

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 2 displays the evolution of median elapsed (incomplete) tenure for our final male and female sample, showing an expected gap between genders.[5] With a difference of more than 3 years in the median tenure of men and women in 1991 but less than 2 years in 2008, the spread however seems to be decreasing.

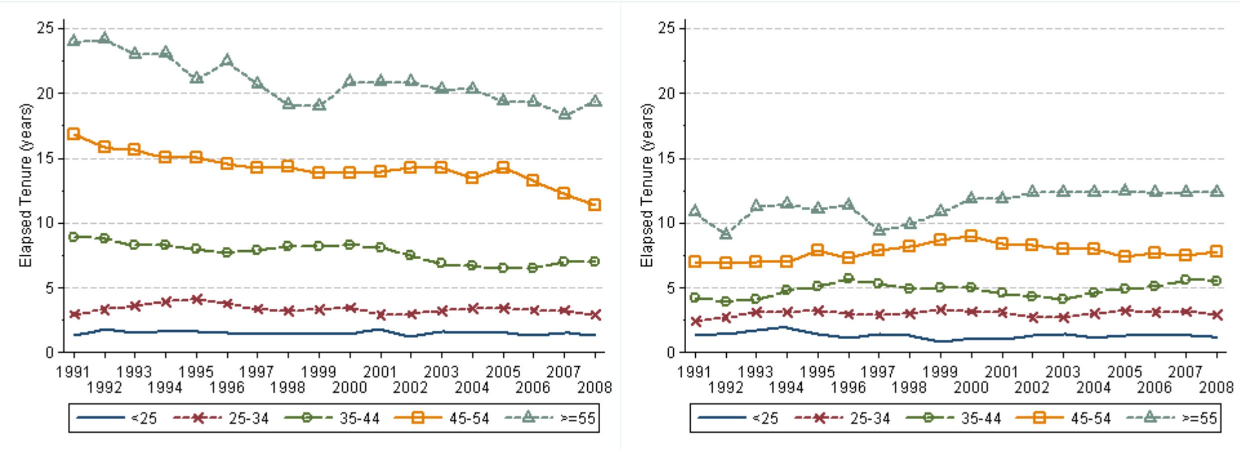
Median tenure is clearly countercyclical. It increased in the early 1990s when unemployment rose and GDP growth was weak. When the economy recovered in the late 1990s, median tenure decreased. Finally, after 2000, median tenure rose again (especially for women), mirroring the rise in unemployment. The countercyclicality 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 3 displays the evolution of median elapsed tenure by age groups.

[5]The values displayed in Figure 2 (median tenure) are much smaller than those in Figure 1 (average tenure) because tenure distribution is right-skewed. This is consistent with data in OECD (1997).

Figure 3: Median elapsed tenure, by age groups

A. Men B. Women



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. Our contention is that although the 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.

The general feeling in the public of a more uncertain labor market might arise from decreasing job tenure among older workers. Foreseeing themselves in the future, young workers might begin to fear more for their career now, even though there are no measured changes for them.

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 investigate job stability.

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 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.[6]

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:

$$h(t|x) = h_0(t) \cdot \exp(x'\beta) \tag{1}$$

where $h_0(t)$ is the baseline hazard function, x is a matrix 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 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 of the sample is thus:

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

For individuals under observation when their job spell starts, $t_{0i} = 0$ and $S(t_{0i}|x_i) = 1$. In our data though, most job spells have already started when the individuals enter the panel survey and these spells are left-truncated, and the solution consists in analyzing only the part of the duration that reaches into the observation window (Guo, 1993). The period before the first interview 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, then allows one to condition on time spent on the job but not in the panel. With the exception of Bergemann and Mertens (2011), this approach using information on all job spells has not been used to analyze job tenure.

[6]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:

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

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

$$\log L = \sum_{e=1}^{m} \log L_e \tag{4}$$

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.[7]

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 simply indicate how the probability of job termination is affected by the covariates. The results are displayed in Table 2.

The coefficient of the year variable provides information on the trend of job stability. In estimations (1), 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.[8,9]

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[7]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 unknown and 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). We use Breslow's (1974) method to handle duration ties.

[8]Quantitatively speaking, the point estimate of 0.006 in estimation (1) for men indicates that the hazard rate of job termination has increased at an average rate of 0.6% (that is, very precisely: $e^{0.006} - 1$) 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 values of the other covariates. For a worker with characteristics at the mode all covariates, this translates into an expected median job spell that decreased by 7 months between 1991 and 2008 (from 4 years and 9 months to 4 years and 2 months).

[9]To account for potential non-linearities, we re-estimated the model by replacing the

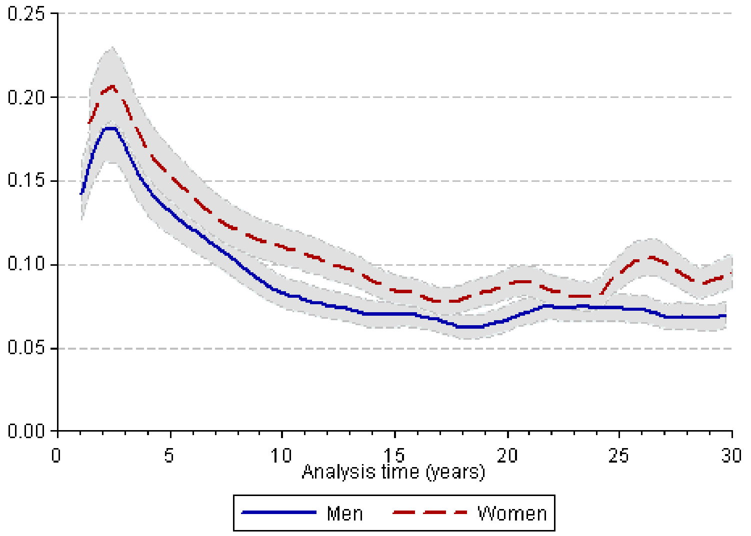
Table 2: Hazard of job termination, by gender

	Men		Women	
	(1)	(2)	(1)	(2)
Year	0.006***	—	0.003	—
	(0.002)		(0.002)	
Age 25-35 years	−0.311***	−0.315***	−0.300***	−0.279***
	(0.035)	(0.036)	(0.032)	(0.033)
Age 35-45 years	−0.550***	−0.563***	−0.586***	−0.587***
	(0.039)	(0.042)	(0.035)	(0.038)
Age 45-55 years	−0.731***	−0.814***	−0.808***	−0.817***
	(0.045)	(0.051)	(0.040)	(0.044)
Age > 55 years	−0.094**	−0.147***	−0.456***	−0.426***
	(0.044)	(0.048)	(0.043)	(0.045)
Year × Age 18-25	—	−0.005	—	0.012**
		(0.006)		(0.005)
Year × Age 25-35	—	−0.002	—	−0.004
		(0.003)		(0.003)
Year × Age 35-45	—	0.005	—	0.006
		(0.004)		(0.004)
Year × Age 45-55	—	0.027***	—	0.010*
		(0.006)		(0.005)
Year × Age > 55	—	0.016***	—	−0.005
		(0.005)		(0.006)
# spells	43,224	43,224	39,536	39,536
# individuals	36,795	36,795	33,328	33,328
# failures	10,061	10,061	10,456	10,456
LogL	−81,827	−81,815	−87,244	−87,238
AIC	163,762	163,746	174,596	174,592
BIC	164,402	164,434	175,229	175,272

Data: Swiss Labor Force Survey, 1991-2008. Robust 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, regional unemployment rate, regional vacancy rate, canton dummies, sector dummies. Complete results are available on request.

Figure 4: Hazard rate of job termination, by gender



Notes: Hazard rate drawn for the mode of the covariates distribution, i.e., for a Swiss individual aged 35-45, married, without children, with medium education, working full-time in a firm of the manufacturing sector, with less than 100 co-workers, in the canton of Zürich, in 2007 (the last year entirely under observation).
Shaded areas are 90% confidence intervals.

The hazard functions corresponding to estimations (1) in Table 2 are plotted in Figure 4, showing how the risk of job termination evolves through a job spell. All 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 seniority because longer job spells are scarce and the trajectories fluctuate strongly afterwards.

Figure 4 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 evolution of the hazard of job termination, though, is similar across gender: it 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

continuous year variable by time fixed effects. The results are presented in Appendix Figure A.1, which displays the coefficients of the time fixed effects. 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.

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 *et al.*, 1999; Marinescu, 2009; Hirsch and Schnabel, 2010; Bergemann and Mertens, 2011). It indicates that jobs 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 become very unlikely to be dissolved.

This shape of the hazard rate can be explained if jobs are "experience goods".[10] In this case, Jovanovic (1979) shows that the probability of leaving a job may initially rise with tenure, because 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 3 in Section 3. To explore this issue, we replace the single year variable by interactions terms between year and age groups in estimations (2) of Table 2. 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 the regressions of Table 2, coefficients of age groups are difficult to interpret meaningfully. For instance, individuals between 45 and 55 have the most stable employment relationships. Whether their position is better than that of other workers cannot be assessed solely from these results, as some confounding factors may be at play. In order to better grasp the determinants of job stability and its evolution, we split estimations by destination state (new job, unemployment, or inactivity) in Table 3 and by termination reason (quit, layoff, or other reasons) in Table 4.

In the set of estimations displayed in Table 3, all coefficients on age groups can be reasonably interpreted. Older workers exhibit less job-to-job mobility on the labor market. Hazard rates towards unemployment 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, probably under the effect of retirement. Unsurprisingly, we also find a significant increase in the transition

[10]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".

rate towards inactivity for women between 25 and 35, which corresponds to the period where most of them give birth to their first child.[11]

These results are largely consistent with those provided by Hirsch and Schnabel (2010) for separation rates towards employment and non-employment. Their state "employment" is directly comparable to "new job", and their state "non-employment" can be considered as a merger of 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 decreased for young workers and increased for those between 45 and 55. 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 aged 35 to 55, but it decreased for women over 55.

Summing up the broad picture of our results in Table 3, one is tempted to conclude that job insecurity has risen for the youngest workers, but not 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 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 5 draws the hazard functions 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. Hirsch and Schnabel (2010) find hazard rates towards employment and non-employment that look very similar to ours. Moreover, their separation rate towards non-employment is lower than towards employment at any time.

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

Table 3: Competing risks model for the hazard of job termination, by destination state

	Men			Women		
	(1) New job	(2) Unemployment	(3) Inactivity	(1) New job	(2) Unemployment	(3) Inactivity
Age 25-35 years	-0.160^{***}	-0.189	-0.863^{***}	-0.221^{***}	-0.055	0.386^{***}
	(0.044)	(0.116)	(0.180)	(0.040)	(0.126)	(0.111)
Age 35-45 years	-0.405^{***}	-0.051	-0.988^{***}	-0.415^{***}	0.044	-0.307^{**}
	(0.050)	(0.129)	(0.194)	(0.046)	(0.132)	(0.120)
Age 45-55 years	-0.720^{***}	0.197	-0.723^{***}	-0.706^{***}	0.179	-0.413^{***}
	(0.062)	(0.134)	(0.192)	(0.056)	(0.134)	(0.124)
Age > 55 years	-1.070^{***}	0.499^{***}	1.517^{***}	-1.154^{***}	0.112	0.851^{***}
	(0.084)	(0.146)	(0.159)	(0.084)	(0.170)	(0.120)
Year × Age 18-25	-0.012^{*}	0.035^{*}	-0.015	-0.001	0.086^{***}	0.007
	(0.007)	(0.019)	(0.029)	(0.006)	(0.018)	(0.018)
Year × Age 25-35	-0.003	-0.001	0.037^{*}	-0.011^{***}	0.056^{***}	-0.009
	(0.004)	(0.011)	(0.019)	(0.004)	(0.012)	(0.007)
Year × Age 35-45	-0.000	0.008	0.063^{***}	-0.003	0.017	0.026^{***}
	(0.005)	(0.013)	(0.020)	(0.005)	(0.012)	(0.009)
Year × Age 45-55	0.013^{*}	0.015	0.098^{***}	0.005	0.019	0.012
	(0.008)	(0.014)	(0.019)	(0.008)	(0.013)	(0.010)
Year × Age > 55	-0.013	0.025	0.025^{***}	-0.002	0.079^{***}	-0.027^{***}
	(0.012)	(0.017)	(0.006)	(0.013)	(0.022)	(0.007)
# spells	43,224	43,224	43,224	39,536	39,536	39,536
# individuals	36,795	36,795	36,795	33,328	33,328	33,328
# failures	6,591	1,320	1,801	6,018	1,325	2,744
LogL	−55,068	−10,770	−11,802	−51,107	−10,910	−21,157
AIC	110,252	21,657	23,719	102,330	21,933	42,429
BIC	110,940	22,345	24,408	103,010	22,589	43,109

Data: Swiss Labor Force Survey, 1991-2008. Robust 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, regional unemployment rate, regional vacancy rate, canton dummies, sector dummies.

The total number of failures does not correspond to what appears in Table 2, because estimations for the destination state "training" are not reported (349 transitions for men and 369 for women).

Figure 5: Hazard rates of job termination, by destination state

Note: see Figure 4.

In Table 4, 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 consistent and results are in line with expectations. Older workers quit much less frequently, but they are more likely to be laid off. These findings are very similar to those in Marinescu (2009)[12] and Bergemann and Mertens (2011).

As shown by the interaction terms, the layoff hazard rate is on the rise for young male workers. For older male age groups, the evolution is less pronounced. It is also worth of notice that women of all age groups have coefficients for layoffs which are similar in value but of opposite sign to the ones for other reasons. This points to some sort of trade-off between these two exits, which is once again consistent with the increasing attachment of female workers to the labor market.

Figure 6 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 its duration dependence once again 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).

[12]The coefficients on age groups do not appear in the published paper, and we thank Ioana Marinescu for kindly providing us with complete results tables.

Table 4: Competing risks model for the hazard of job termination, by termination reason

	Men			Women		
	(1) Quit	(2) Layoff	(3) Other reasons	(1) Quit	(2) Layoff	(3) Other reasons
Age 25-35 years	-0.207^{**}	0.145	-0.367^{***}	-0.514^{***}	-0.107	-0.229^{**}
	(0.091)	(0.192)	(0.117)	(0.086)	(0.175)	(0.093)
Age 35-45 years	-0.679^{***}	0.475^{**}	-0.743^{***}	-0.786^{***}	0.128	-0.764^{***}
	(0.102)	(0.195)	(0.132)	(0.097)	(0.179)	(0.107)
Age 45-55 years	-1.231^{***}	0.690^{***}	-1.254^{***}	-1.268^{***}	0.431^{**}	-1.189^{***}
	(0.143)	(0.199)	(0.166)	(0.118)	(0.179)	(0.129)
Age > 55 years	-2.166^{***}	0.847^{***}	0.232^{*}	-1.806^{***}	0.406^{*}	-0.024
	(0.233)	(0.213)	(0.127)	(0.187)	(0.209)	(0.117)
Year × Age 18-25	-0.016	0.061^{*}	-0.016	-0.019	0.023	-0.003
	(0.017)	(0.033)	(0.021)	(0.014)	(0.030)	(0.017)
Year × Age 25-35	-0.019^{**}	0.014	-0.020^{*}	0.016^{*}	0.033^{*}	-0.036^{***}
	(0.009)	(0.015)	(0.012)	(0.010)	(0.017)	(0.010)
Year × Age 35-45	0.003	-0.009	-0.003	0.020^{*}	0.013	-0.019
	(0.011)	(0.015)	(0.015)	(0.012)	(0.016)	(0.013)
Year × Age 45-55	0.028	0.011	0.051^{**}	0.024	0.029^{*}	-0.026
	(0.020)	(0.016)	(0.022)	(0.017)	(0.016)	(0.020)
Year × Age > 55	0.052	0.034^{*}	0.046^{***}	0.038	0.055^{**}	-0.048^{***}
	(0.037)	(0.020)	(0.011)	(0.032)	(0.024)	(0.015)
# spells	37,303	37,303	37,303	33,973	33,973	33,973
# individuals	32,042	32,042	32,042	28,972	28,972	28,972
# failures	3,151	1,766	2,572	3,215	1,548	2,616
LogL	−25,482	−13,777	−19,295	−26,352	−12,205	−20,851
AIC	51,079	27,669	38,707	52,820	24,526	41,817
BIC	51,756	28,346	39,384	53,489	25,195	42,486

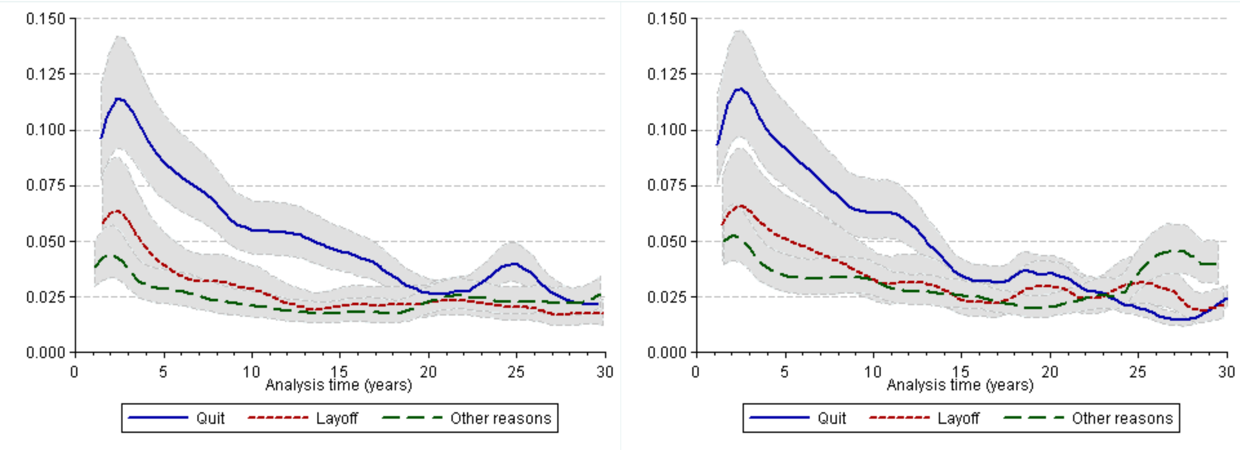
Data: Swiss Labor Force Survey, 1996-2008. Robust standard errors in parentheses. $^{***}/^{**}/^{*}$: Significant at the 0.01/0.05/0.10 level. Base category for age groups: 18-25 years.

Unreported covariates: education groups, foreign nationality, marital status, number of children, part-time work, firm with more than 100 co-workers, regional unemployment rate, regional vacancy rate, canton dummies, sector dummies.

Number of observations are smaller than in Tables 2 and 3 because information about job termination reason is only available since 1996.

Figure 6: Hazard rates of job termination, by termination reason

A. Men B. Women



Note: see Figure 4.

5.1 The effects of skills

In line with Bergemann and Mertens (2011), we finally test whether skill-biased technological change (SBTC) might be responsible for changes in job stability. Under the SBTC assumption, workers with advanced skills should benefit from better job prospects over time, whereas workers with mostly simple skills should see their situation worsen (see e.g., Autor *et al.*, 2008). In our framework, changes should in particular be observed on the hazard of layoff.

Using the procedure proposed by D'Amuri and Peri (2011), we build measures of occupational skill intensity using the *O*NET* data from the US Department of Labor.[13] This survey assigns values summarizing the importance of different skills to occupations (according to the Standard Occupation Classification, SOC2010). Starting from 78 tasks measured in *O*NET*, we create 5 measures of skill intensity in each occupation: communication, complex, mental, manual and routine. The first 3 skills can be labelled as "advanced", whereas the last 2 are "simple" skills.[14]

In the SLFS, occupations are coded using the International Standard Classification of Occupation (ISCO88), and we use crosswalks (SOC2010 to

[13]Version 17.0, available at `http://www.onetcenter.org/database.html`

[14]D'Amuri and Peri (2011, Appendix Table A2) provide a complete list of the 78 tasks in question and their distribution across skills. We here re-labelled the first broad group of skills as "advanced" instead of "complex" in order to avoid any confusion between the broad group and the skill itself.

SOC2000 and SOC2000 to ISCO88) to merge *O*NET* data with our main dataset.[15] Because codes do not match one to one during the process, we average SOC2010 occupations into a unique ISCO88 occupation using US levels of employment as weights.[16]

The estimations encompassing skills measures are reported in Table 5. Even though few coefficients are significant, the results broadly support the SBTC hypothesis. As mentioned before, we expect to observe most changes to occur for layoffs. Accordingly, we find an increasingly negative effect of communication skills on the hazard rate of layoff, its coefficient being especially significant for women. For women, the hazard rate towards unemployment is also evolving as expected under the SBTC hypothesis: in occupations using mental skills, transitions towards unemployment become less frequent, whereas the opposite takes place in occupations using more heavily manual skills. For other destination states and termination reasons, the results are less easily interpretable, but these are not as closely linked to the SBTC hypothesis.

[15] The crosswalks are available at `http://www.bls.gov/soc/soccrosswalks.htm` and `http://www.xwalkcenter.org/index.php`, respectively.

[16] Number of workers by occupations come from the May 2011 Occupational Employment Statistics, see `http://www.bls.gov/oes/`.

Table 5: Competing risks model for the hazard of job termination, with skills measures

	Men					
	(1) New job	(2) Unemployment	(3) Inactivity	(4) Quit	(5) Layoff	(6) Other reasons
Year	−0.054	0.066	0.084	−0.110**	0.234***	−0.022
	(0.034)	(0.075)	(0.064)	(0.053)	(0.073)	(0.057)
Year × Communication Skills	0.035**	−0.033	0.009	0.061***	−0.061*	0.012
	(0.015)	(0.036)	(0.028)	(0.023)	(0.034)	(0.025)
Year × Complex Skills	−0.005	−0.009	−0.002	0.014	−0.020	−0.012
	(0.008)	(0.019)	(0.016)	(0.014)	(0.017)	(0.015)
Year × Mental Skills	−0.011	0.028	−0.034	−0.063***	0.042	0.011
	(0.014)	(0.035)	(0.028)	(0.022)	(0.033)	(0.024)
Year × Manual Skills	0.019**	−0.008	0.026*	0.031**	0.008	0.003
	(0.009)	(0.021)	(0.016)	(0.014)	(0.019)	(0.015)
Year × Routine Skills	−0.016*	0.007	−0.006	−0.007	−0.025	0.001
	(0.009)	(0.020)	(0.016)	(0.014)	(0.018)	(0.015)
# spells	32,815	32,815	32,815	28,260	28,260	28,260
# individuals	28,457	28,457	28,457	24,642	24,642	24,642
# failures	4,976	972	1,285	2,440	1,292	1,900
LogL	−40,059	−7,602	−8,066	−19,000	−9,660	−13,749
AIC	80,236	15,318	16,249	38,117	19,437	27,614
BIC	80,917	15,977	16,931	38,788	20,096	28,273

	Women					
	(1) New job	(2) Unemployment	(3) Inactivity	(4) Quit	(5) Layoff	(6) Other reasons
Year	−0.051	0.104	0.091*	−0.060	0.300***	−0.126**
	(0.034)	(0.073)	(0.050)	(0.054)	(0.072)	(0.061)
Year × Communication Skills	0.007	0.033	−0.053**	−0.002	−0.093***	0.042
	(0.016)	(0.036)	(0.025)	(0.025)	(0.036)	(0.028)
Year × Complex Skills	−0.005	0.004	0.002	−0.004	−0.001	−0.002
	(0.008)	(0.018)	(0.013)	(0.013)	(0.017)	(0.014)
Year × Mental Skills	0.021	−0.068*	0.046*	0.045*	0.041	−0.013
	(0.016)	(0.035)	(0.026)	(0.024)	(0.036)	(0.027)
Year × Manual Skills	−0.008	0.051**	−0.018	−0.006	−0.026	0.018
	(0.011)	(0.024)	(0.017)	(0.017)	(0.024)	(0.020)
Year × Routine Skills	−0.003	−0.030	−0.002	−0.006	−0.005	−0.011
	(0.010)	(0.021)	(0.015)	(0.015)	(0.021)	(0.018)
# spells	32,513	32,513	32,513	27,683	27,683	27,683
# individuals	27,770	27,770	27,770	23,881	23,881	23,881
# failures	5,002	1,078	2,177	2,665	1,217	2,110
LogL	−41,486	−8,646	−16,366	−21,267	−9,334	−16,438
AIC	83,089	17,404	32,849	42,651	18,783	32,992
BIC	83,769	18,049	33,529	43,319	19,428	33,648

Data: Swiss Labor Force Survey, 1991-2008 for estimations (1) to (3) and 1996-2008 for estimations (4) to (6). Robust standard errors in parentheses. ***/**/*: Significant at the 0.01/0.05/0.10 level.

Unreported covariates: education groups, foreign nationality, marital status, number of children, part-time work, firm with more than 100 co-workers, regional unemployment rate, regional vacancy rate, canton dummies, sector dummies.

Number of observations are smaller than in Tables 3 and 4 because some industries of the SLFS could not be assigned *O*NET* tasks measures.

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 often without the appropriate econometric tools. In line with recent literature, we make use of competing risks models, which allows 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 workers over 45, we do however find a decrease in job stability.

The analysis of job stability, however, remains of rather limited interest, as it simply shows that job spells are becoming shorter without giving any indications about the underlying mechanisms at play. If workers decide to change jobs more often, job stability would indeed decrease but that would not necessarily prove that their welfare has diminished.

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

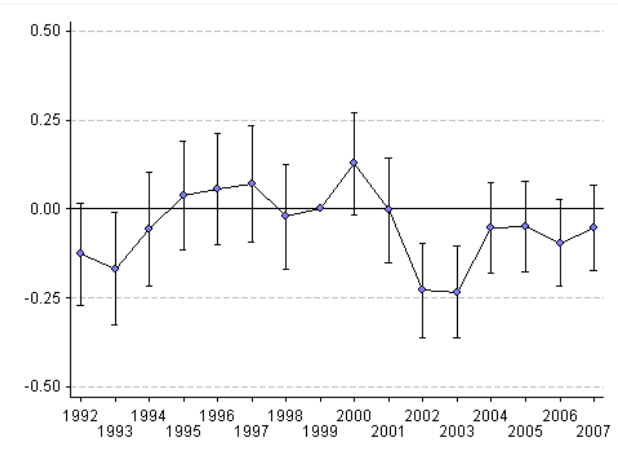
Finally, let us mention that it would be interesting to combine both destination states and termination reasons, since they are complementary decompositions. One could indeed imagine that the probabilities of quitting towards a new job and quitting towards unemployment or inactivity are different. The 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 sufficient observations to estimate such models, which we leave in the agenda of future research.

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 much 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 more research is needed to test these assumptions.

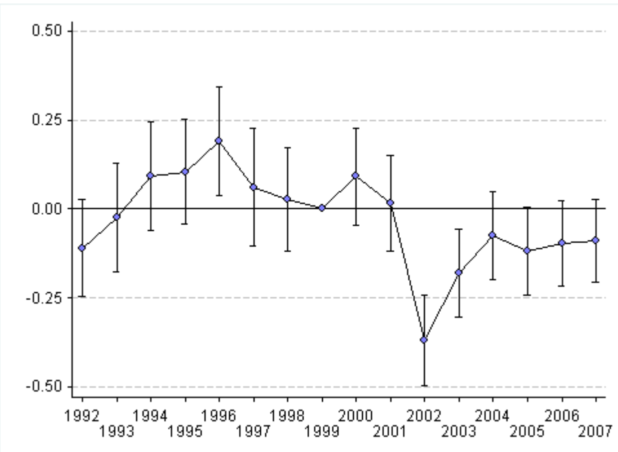
Appendix A: Evolution of Job Stability

Figure A.1: Estimated coefficients for the evolution of the hazard of job termination

A. Men



B. Women



Notes: The coefficients are from an estimation similar to (1) in Table 2, where the variable year has been replaced by time fixed effects with 1999 as reference. Whiskers are 95% confidence intervals.

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