NEW JOBS OR RECALLS ?

Flow Dynamics in Austrian Unemployment
Reconsidered

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Abstract

The results of this study demonstrate the importance of explicitly accounting for the possibility of recalls in the analysis of unemployment composition and the determinants of unemployment spell durations in Austria. In our sample covering unemployment spells in 1985 we find that recalls accounted for nearly one half of the employment to unemployment to re–employment transitions with the probability of recall being mainly dependent on industry and job characteristics related to seasonal work. In particular, the estimation results from a simple logistic model suggest that temporary layoffs may constitute a regular pattern in the work life of individuals affected by that type of unemployment. We then analyze unemplyoment spell durations in a competing risks framework and, indeed, find significantly different hazards for the two types of risks, new jobs and recalls. While both exit rates exhibit positive duration dependence according to our estimates, the new job hazard is considerably flatter than the recall hazard. The estimated effects of the covariates are also quite different across new jobs and recalls. Thus, failure to distinguish between different types of layoffs may lead to a serious misperception of unemployment dynamics in Austria. This assertion conforms well with previous findings for the U.S and the Danish labour market and suggests a reinspection of re–employment patterns in other European labour markets as well.

1. Introduction

Temporary layoffs (rehires, recalls) have been widely recognized as an important feature of North America's labour markets. Feldstein (1975) and Lilien (1980) derive estimates of rehires amounting to over 70 per cent of the laid off workers in US manufacturing. Topel and Welch (1980) report that in the period 1973–1976 approximately one third of the Unemployed had been recalled by their former employer. Katz and Meyer (1988) find that over 30 per cent of the total weeks of unemployment in a sample of unemployment insurance recipients in Missouri and Pennsylvania were attributable to unemployment spells ending in recall. Similarly, a study by Robertson (1989) comes up with pretty much the same numbers for Canada. It indicates that in 1984 about 50 per cent of all employment separations and 60 per cent of layoffs ended with individuals returning to their former employer; workers on temporary layoff thereby accounted for about 35 per cent of the Unemployed.

According to the seminal contributions by Feldstein (1976,1978) and Baily (1977) the subsidy element contained in the unemployment insurance benefit system is responsible for an excessive extent of temporary layoff unemployment in the US. From this point of view one might expect an even higher amount of temporary layoff unemployment in Europe since most European UI benefit systems do not contain any element of "experience rating" of UI contributions. However, several observers have suggested that tighter regulatory impediments to the recruitment and dismissal of employees are responsible for temporary layoff unemployment being virtually non-existent in Europe (see e.g. Moy, Sorrentino (1981); for Austria: Gutierrez-Rieger, Podczeck (1981); for legal restraints on layoffs in general: Emerson (1988)).

However, firm specific human capital, as well as efficient risk shifting and, last not least, possibly collusive behaviour against the unemployment insurance fund may make implicit contracts regarding temporary separations attractive for both workers and firms when facing temporary downturns in production. And indeed, recent empirical evidence for two European countries suggests that temporary layoffs may constitute an important feature of labour markets even in the absence of any institutionalized regulations concerning this type of unemployment. Jensen and Westergard-Nielsen (1989) show that at least 40 per cent of all unemployment spells in Denmark in the period 1979–1984 were due to temporary layoffs and that these spells accounted for at least 16 per cent of total unemployment. Fischer and Pichelmann (1991) report that about one third of all unemployment spells in Austria in 1985 ended with the individual returning to the former employer; like in Denmark temporary layoff unemployment was found to be on average of shorter duration but, nevertheless, accounting for some 20 per cent of total weeks of unemployment.

In this paper we use the data set from Austria to demonstrate the importance of explicitly accounting for the possibility of recalls in the analysis of unemployment composition and the determinants of unemployment spell durations in Austria. Drawn from administrative records this unique sample of unemployed registered in 1985 allows to construct a longitudinal data set covering not only individual employment and unemployment patterns with a broad variety of single spell characteristics (including an annonymous employer's identification number) but also various states out of the labour force such as maternity, sickness, pension and a few others of minor importance.

We first focus on the estimation of a simple logit model to empirically determine the impact of several socio–economic variables which theory predicts to affect the likelihood of an individual spell ending in recall rather than in finding a new job. The empirical findings suggest that recall probabilities are mainly dependent on industry and job characteristics related to seasonal work. In particular, the estimation results indicate that temporary separations may constitute a regular pattern in the work life of individuals affected by that type of unemployment.

Second, recall and new job exit probabilities may exhibit quite different time patterns and may also be affected by explanatory variables in a different way (Katz (1986); Katz, Meyer (1988)); in particular, empirical results on duration dependence in the transition rates from unemployment to employment may no longer hold when temporary layoffs are accounted for (Han, Hausman (1990); Jensen, Westergard–Nielsen (1989)). Therefore, we investigate the robustness of single risk specifications typically used in most studies of unemployment spell durations (see e.g. Frühstück, Pichelmann (1987), Winter–Ebmer (1990), Steiner (1990) for applications of duration models of this type to Austrian data) by alternatively estimating a competing risks model of unemployment duration in which re–employment can be either with a new employer or with the previous employer. In the empirical analysis we, indeed, find significantly different hazards for the two types of risks, new jobs and recalls. While both exit rates exhibit positive duration dependence according to our estimates, the new job hazard is considerably flatter over time than the recall hazard. The estimated effects of the covariates are also quite different across new jobs and recalls. This again demonstrates the importance of explicitly accounting for different types of layoffs.

The remainder of the paper is organized as follows: Section 2 provides information on the data and the measurement procedures to identify temporary layoff unemployment in Austria. Section 3 presents the findings regarding the factors affecting the probability of individual unemployment spells ending with recall rather than with exit to a new job. In Section 4 standard econometric duration models are used to analyze the transition rates from unemployment to employment in a competing risks framework. Section 5 concludes.

2. Data and Measurement Design

The data set consists of a random sample of all individuals who had been registered as unemployed by the Labour Offices in Austria in the course of the year 1985. The number of workers experiencing a (registered) spell of unemployment amounted to about 450 000 in 1985, about 16 per cent of the Austrian labour force; the average rate of unemployment stood at 4.5 per cent. The sample covers 2499 individuals, approximately 0.5 per cent of the total population under consideration. Various comparisons with official data indicate that the sample properly reflects the socio-economic composition of the registered unemployed in 1985.

Table 1

SAMPLE CHARACTERISTICS: SEX AND AGE

	UNDER 25	25-44	OVER 44	Row Total
MALES	535 35.1 56.9 21.4	721 47.3 63.1 28.9	268 17.6 64.3 10.7	1524 61.0
FEMALES	405 41.5 43.1 16.2	421 43.2 36.9 16.8	149 15.3 35.7 6.0	975 39.0
Column Total	940 37.6	1142 45.7	417 16.7	2499 100.0

Combining the unemployment data with Social Insurance Records data allows to construct a longitudinal data set covering not only individual employment and unemployment patterns with a broad variety of single spell characteristics but also various states out of the labour force such as maternity, sickness, pension and a few others of minor importance. The observation period dates back till 1972,January, and ends in 1988, August.

Preliminary inspection of the data showed that 72 persons in the sample (2.9% of the total) had only registered as searching for a job while still being employed but never actually had entered unemployment. These cases have been eliminated from the subsequent analysis, which focusses on unemployment spells dating in 1985, but not necessarily starting or ending in that year. If an individual had experienced more than one spell of unemployment in 1985, one of the spells was randomly selected. Following administrative procedures, interruptions of unemployment due to sickness, work on a daily basis etc. that lasted no longer than 28 days have been ignored.

Since each firm is issued an (anonnymous) identification number, which is separately recorded for every single spell of employment, temporary separations can be identified. The measurement design is as follows:

starting from the unemployment spell under consideration the next previous and the next subsequent spell of employment and – if encountered – its characteristics were determined. Regarding the spell inflow origin states we allow for a possible time interval classified as out of the labour force between the termination of the last job and the beginning of the unemployment spell. If the latter time interval was less or equal 28 days an employment–unemployment transition is said to have occurred. The analogous definition is applied with respect to transitions from unemployment to employment. Temporary layoff unemployment then is simply defined as the situation where the firm's identification numbers of the two employment spells are identical. It should be noted that this procedure tends to result in a slightly conservative estimate of temporary layoffs, because rehires occurring later than four weeks after the outflow from registered unemployment have been neglected. The category 'no next job' consists predominantly of transitions from unemployment to (early) retirement and withdrawal from the labour force for unknown reasons, but also contains a few censored cases, i.e. unemployment spells still in progress at the end of the observation period.

Table 2

UNEMPLOYMENT INFLOW ORIGIN BY OUTFLOW DESTINATION STATES
(in per cent; sum of cells = 100)

	SPELL OUTFLOW DESTINATION				
SPELL INFLOW ORIGIN	EMPLOYMENT NEW JOB	EMPLOYMENT RECALL	> 1 MONTH OUT OF LF	NO NEXT JOB	Row Total
EMPLOYMENT	31.3	28.4	8.5	7.0	75.3
OUT OF LF ≤ 6 MONTH	5.1	1.2	2.4	1.4	10.2
OUT OF LF > 6 MONTH	5.6	.7	2.6	4.0	12.9
NO PREVIOUS JOB	1.0	-	.3	.3	1.6
Column Total	43.0	30.3	13.9	12.8	100.0

Number of Observations = 2427

Table 2 clearly shows that temporary layoffs (recalls, rehires) constitute an important element of unemployment in Austria. Three out of ten workers experiencing a registered spell of unemployment in 1985 returned to work with their previous employer. The picture gets even more impressive, if one restricts attention to the employment–unemployment–employment transitions with no interim period out of the labour force lasting for more than four weeks. The subsequent analysis will be confined to this sub-group of workers with uninterrupted labour market attachment. A transition pattern of this type was encountered in approximately sixty per cent of all unemployment spells, and nearly one half of these spells were of the temporary layoff type.

Last not least, in order to check whether workers were repeatedly affected by temporary layoffs, we have also searched for another spell of unemployment both backwards and forewards; if a previous or a subsequent spell of unemployment was encountered, the same procedure to identify temporary separations was applied. The full details regarding the sample design and measurement procedures are spelled out in Beidl et al.(1990).

3. New Jobs or Recalls: Incidence Analysis

What determines whether an individual unemployment spell ends through rehire to the former employer rather than through exit to a new job? In this section we provide a tentative answer to this question by estimation of simple logistic regression model of the form

$$(1) \qquad \text{Prob (Recall; X)} = 1 / (1 + e^{-X\beta})$$

where X denotes a vector of explanatory variables and the unknown coefficients β are estimated from the data using the maximum likelihood method.

The variables in vector X may be classified into several groups. First, a priori considerations suggest to control for the impact of seasonality on layoff decisions. Seasonal unemployment is clearly related to temporary layoff unemployment by its usually passing and short term character; moreover, seasonal downturns in demand and production are easily predictable. Despite the fact that there exists no institutionalized regulatory framework with regard to temporary layoffs in Austria, employers may, therefore, turn to implicit temporary layoffs as a response to seasonal demand variations. In order to capture the conjectured seasonality in temporary layoff spells we include an industry variable and a dummy variable on whether the unemployment spell started in winter; furthermore, the tenure of the previous employment spell may serve as an indicator for seasonal fluctuations in employment, e.g. one might expect seasonal workers in agriculture or construction to be employed for about nine months.

The previous employment duration may, on the other hand, also pick up an effect from experience which may positively affect the recall probability (Haltiwanger (1984)). Unfortunately, no direct observations on experience and/or qualification are available; therefore only age and, to some extent, earnings in the previous job can serve as a proxy for experience. Higher earnings may also be positively correlated with the probability of recall via a wage compensation mechanism for more temporary layoff unemployment.

Table 3

DESCRIPTIVE STATISTICS: MEANS OF VARIABLES

VARIABLE	NEW JOB	RECALL	TOTAL
TOTAL	.524	.476	1.000
SEX			
FEMALE	.388	.323	.357
AGE	30.2	33.4	31.7
BELOW 25	.246	.158	.204
25 - 45 YEARS	.580	.564	.572
OVER 45	.174	.278	.223
MARRIED			
YES	.412	.541	.473
HARD TO PLACE			
YES	.172	.120	.148
CHILDREN			
YES	.388	.455	.420
UNEMPLOYMENT DURATION			
- 1 MONTH	.186	.113	.151
1 - 3 MONTHS	.361	.487	.421
3 - 6 MONTHS	.104	.201	.150
6 -12 MONTHS	.297	.190	.246
OVER 12 MONTHS	.053	.009	.032
INDUSTRIES			
AGRICULTURE	.016	.052	.033
ENERGY	.001	.000	.001
MINING	.005	.019	.012
MANUFACTURING	.287	.148	.221
CONSTRUCTION	.163	.364	.259
TRADE	.168	.058	.116
TOURISM	.167	.184	.175
TRANSPORTATION	.033	.038	.035
FINANCIAL SERVICES	.042	.017	.030
SOC. & GOV. SERVICES	.117	.120	.119
PREVIOUS EMPLOYMENT DURATION			
- 1 MONTH	.159	.055	.110
1 - 3 MONTHS	.208	.110	.161
3 - 6 MONTHS	.104	.087	.096
6 - 12 MONTHS	.266	.599	.424
OVER 12 MONTHS	.263	.149	.209
PREVIOUS EARNINGS			
SCHILLING PER DAY	409.4	434.5	421.4
PREVIOUS UNEMPLOYMENT			
YES	.658	.836	.743
PREVIOUS TEMPORARY LAYOFF			
YES, TOTAL	.167	.639	.392
IN AGE GROUP BELOW 25	.024	.049	.036
IN AGE GROUP 25 - 45	.116	.371	.237
IN AGE GROUP OVER 45	.028	.219	.119
UNEMPLOYMENT SPELL			
START IN WINTER	.397	.669	.526

As regards individual socio–economic characteristics we have included the usual set of variables: sex, age, marital status, the presence of children and placement restrictions may all affect search behaviour and thus recall probabilities.

Finally, not all instances of rehire by the former employer indicate collusion between the employer and employee. Even in the absence of any implicit contract extending into the period of unemployment, rehiring may occur thanks to a continuing geographical or occupational match between employee characteristics and the employer's needs. However, this interpretation is less plausible when temporary layoffs become a regular pattern in the work life of the individuals affected by that type of unemployment. To control for this effect we include an indicator variable for a previous unemployment spell and age–specific indicator variables for a previous unemployment spell of the temporary layoff type. Descriptive summary statistics are supplied in Table 3.

The estimation results and some test statistics of model (1) are summarized in Table 4. A positive B coefficient implies that the corresponding category of the variable raises the probability of a recall spell outcome. Put precisely, the sum of the B coefficients for a certain combination of variables gives the logarithmic odds of being rehired for a person with that characteristics. The overall test statistics suggest that the model fits the data reasonably well. Note, though, that insignificant variables have not been removed from the model.

The general picture that emerges from the empirical analysis is that the probability of recall is mainly dependent on industry and job characteristics related to seasonal work, whereas all of the socio–economic background variables of the individuals as well as previous earnings turn out to be of no significant importance.

Regarding industries we find (at the 5 per cent level) significant increases in the log–odds for being rehired – relative to manufacturing – in agriculture, construction, tourism and, somewhat surprisingly, social and public services. The estimates do not imply, however, that non–negligible chances of recall are confined to industries one might a priori consider as being prone to seasonal fluctuations in production. Evaluated at the means of the continous variables and the base level of the categorical variables the probability of recall in manufacturing, for instance, is estimated to be 0.19; the corresponding number for trade is 0.15 and for social and public services 0.29. In the interpretation of these results one should keep in mind, though, that economic activity might contain a seasonal component probably related to agriculture, construction or tourism in several other industries ; e.g. food–processing, transportation, entertainment and sport, tourism–related local community services and so forth.

Table 4

LOGISTIC REGRESSION: PROBABILITY OF RECALL BY FORMER EMPLOYER

VARIABLE	B	S.E.	WALD	SIG	EXP(B)
CONSTANT	-1.0680	.3467	9.4882	.0021	
SEX					
FEMALE	.0357	.1646	.0023	.8283	1.0363
AGE	-.0079	.0082	.9278	.3352	.9921
MARRIED					
YES	-.0082	.1688	.0023	.9614	.9919
HARD TO PLACE					
YES	-.2487	.1865	1.7781	.1824	.7798
CHILDREN					
YES	.1800	.1628	1.2223	.2689	1.1972
INDUSTRIES			36.1687	.0000	
AGRICULTURE	.9526	.4226	5.0816	.0242	2.5924
ENERGY	-4.0335	13.5012	.0893	.7651	.0177
MINING	.8950	.6638	1.8180	.1775	2.4474
MANUFACTURING	base				
CONSTRUCTION	.7474	.2005	13.9029	.0002	2.1115
TRADE	-.2663	.2479	1.1543	.2826	.7662
TOURISM	.6496	.2174	8.9329	.0028	1.9149
TRANSPORTATION	.6283	.3599	3.0486	.0808	1.8745
FINANCIAL SERVICES	-.4698	.4129	1.2949	.2551	.6251
SOC. & GOV. SERVICES	.5506	.2295	5.7540	.0165	1.7343
PREVIOUS EMPLOYMENT DURATION			40.8477	.0000	
- 1 MONTH	-.7776	.2550	9.3013	.0023	.4595
1 - 3 MONTHS	-.5601	.2240	6.2516	.0124	.5712
3 - 6 MONTHS	-.0770	.2532	.0925	.7610	.9259
6 - 12 MONTHS	.3930	.1863	4.4494	.0349	1.4813
OVER 12 MONTHS	base				
PREVIOUS EARNINGS					
SCHILLING PER DAY	-.0003	.0005	.5468	.4596	.9997
PREVIOUS UNEMPLOYMENT					
YES	-.2572	.1696	2.2986	.1295	.7732
PREVIOUS TEMPORARY LAYOFF					
YES, AGE GROUP BELOW 25	1.3883	.3378	16.8918	.0000	4.0080
YES, AGE GROUP 25 - 45	1.6841	.1758	91.7431	.0000	5.3878
YES, AGE GROUP OVER 45	2.6264	.3059	73.7354	.0000	13.8245
UNEMPLOYMENT SPELL					
START IN WINTER	.7417	.1421	27.2258	.0000	2.0995

	Chi-Square	df	Significance
-2 Log Likelihood	1494.470	1425	.0980
Model Chi-Square	512.276	24	.0000
Goodness of Fit	1474.993	1425	.1741

Classification Table for Temporary Layoff Unemployment

Observed	Predicted NEW JOB	Predicted RECALL	Correct
NEW JOB	628	132	82.63%
RECALL	199	491	71.16%
		Overall	77.17%

The importance of temporary layoffs to accomodate seasonal fluctuations, in particular in the construction industry, can clearly be seen when the combined effects of the other proxies for seasonality are taken into account. The mere fact of an unemployment spell starting in winter ceteris paribus approximately doubles the odds of being rehired. The average male construction worker when being laid off in winter and having experienced a previous employment duration of more than 3 months faces a 50 – 60 per cent chance of being recalled by his former employer. If, additionally, this construction worker had already experienced a previous spell of temporary layoff unemployment, the estimated recall probability increases to 0.8 – 0.9 . The size of the coefficients on previous temporary separations suggest generally that temporary layoffs may constitute a regular pattern in the workers' employment careers. Concluding this exemplary discussion of estimation results it may be interesting to note that in tourism recalls are on average somewhat less likely than in construction given that in tourism unemployment spells starting in winter occur less frequent and the tenure of the previous job may well be below 3 months.

4. New Jobs or Recalls: Unemployment Duration Analysis

Unemployment spells ending in recall are on average of significantly shorter duration than spells that result in new jobs. In our sample this difference amounted to approximately one month; furthermore, the variance of spell durations is much less for spells ending in recall; see table 5.

In this section we focus on the duration determinants in terms of explanatory variables and duration dependence effects in a proportional hazard model. Neglecting the different types of exit routes into re-employment, however, may lead to a serious bias in results and conclusions. Therefore, the analysis is based on the estimation of a simple competing risk model, with 'recall' and 'new job' as the distinct risks. Compared to a single risk specification this procedure allows to check whether the effects of the explanatory variables and duration dependence patterns differ considerably for these two risks.

Table 5

MEAN DURATION OF UNEMPLOYMENT SPELLS (IN DAYS)

	MEAN	STD DEV	CASES
FOR ENTIRE POPULATION	110	131.95	1450
NEW JOB	128	166.37	760
RECALL	91	73.43	690

The basic tool in the analysis of duration models is the hazard function which determines the probability of leaving a state conditional on the duration in this state and a set of explanatory variables. In case of competing risks models, the relation

(2) $h_e\ (t; Z) = h_n\ (t; Z) + h_r\ (t; Z)$

holds, which says that the total re–employment hazard, $h_e\ (t; Z)$, equals the sum of the recall exit rate $h_r\ (t; Z)$, and the hazard rate of finding a new job $h_n\ (t; Z)$. Z denotes the vector of explanatory variables. Competing risks models determine separately a hazard function for each of the destination states in question. In the estimation of the recall hazard, spells ending in the finding of a new job are treated as censored at the date of new job finding; spells ending in recall are analogously treated as censored at the recall date in the estimation of the new job hazard.

In the empirical analysis we employ a parametric estimator for the exit rates. The specific functional form chosen is the Weibull model which is the most popular specification in unemployment duration studies (Kiefer (1988)). The Weibull model is given by

(3) $h\ (t; Z) = \alpha\ t^{\alpha-1} \exp\ (\beta'Z)^{\alpha}$

where $\alpha\ t^{\alpha-1}$ is the baseline hazard function assumed identical for all individuals and β is a vector of unknown coefficients. In the absence of time–varying covariates included in Z the shape parameter α solely determines the evolution of the exit rates over time: for $\alpha > 1$ the hazard rate monotonically increases, $0 < \alpha < 1$ indicates a falling hazard rate, and for $\alpha = 1$ the baseline hazard is a constant and the Weibull model specializes to the exponential case. The Weibull model thus offers a convenient way to test for duration dependence mechanisms in the transition to re–employment.

Obviously, the imposition of a monotonic downward or upward form of the baseline hazard which is implied by the specification of one–parameter member of the Weibull family may be too restrictive and thus yield inconsistent parameter estimates. However, some experimentation with a different parametric (Log–Logistic model) and a semi–parametric approach (Cox model) seems to indicate that the Weibull specification captures the main features of the data sufficiently well.

The estimation was performed by a ML procedure using the program RATC 1.1 (Rohwer 1990). The results for the recall and the new job exit rates in the competing risk framework and for the corresponding compressed single risk model are reported in Table 6; for a separate estimation for the male and female populationn in the sample see Tables A1 and A2 in the appendix. A positive β coefficient implies that the corresponding variable raises an individual's exit rate relative to the baseline hazard. Note again, though, that like in the previous section insignificant variables have not been removed from the estimation.

We start our discussion of empirical results with the observation that the estimated shape parameter α is significantly greater than unity both for the recall and the new job hazard indicating positive duration dependence in both exit ways to re-employment. The size of the estimated α's, however, differs considerably between the two groups with the recall hazard function rising faster than the new job exit rate. It may be interesting to note that these time patterns of exit rates are at variance with the results reported by Katz and Meyer (1988) for the U.S. and Jensen and Westergard-Nielsen (1989) for the Danish labour market; they consistently find downward sloping recall hazard rates. Our estimates indicating increasing re-employment hazard rates in the compressed single risk model, however, are compatible with previous analyses of Austrian unemployment duration data by Winter-Ebmer (1990) and Steiner (1990) who have obtained strikingly similar results.

Apart from the distinct time patterns of the baseline hazards the compressed single risk estimates also mask many large differences between the effects of the covariates on the recall and the new job exit rates. A previous spell of unemployment, for instance, significantly incrases unemployment duration for new jobs, but tends to reduce unemployment duration for recalls. Previous long-term unemployment, on the other hand, reduces both exit rates quite substantially, though with the negative impact being much stronger on the recall hazard than on the new job exit rate. Having held a job in construction, to take another example, lowers the new job hazard but increases the recall hazard with the latter effect being dominant in the single risk estimation. Or consider previous wages; higher pre-unemployment wages significantly increase the new job exit rate but have no significant impact on the recall exit rate according to our estimates. Being classified as hard-to-place severly impairs new job exits, but the negative impact on the recall hazard rate is even more accentuated. Lastly note that the negative effect of increasing age on the re-employment rates is much larger for new job exits than for recalls. Thus, the estimated effects of the covariates are quite different across new jobs and recalls. This again demonstrates the value of the competing risk specification which allows the disentangling of the two processes which produce the aggregated re-employment hazard. For a visual impression of these empirical results the reader is referred to Charts A1 & A2 in the appendix which give a plot of the estimated new job and recall hazard for two different sets of covariates.

Obviously, our estimates suffer from the potential problem of omitted variables in the list of covariates. It is well known that in hazard rate models uncontrolled heterogeneity biases parameter estimates towards zero, which may result in spurious findings of negative duration dependence. Note, though, that a bias in the opposite direction may arise in a competing risk framework (Katz, Meyer (1988)). If uncontrolled factors that raise the recall hazard also lower the new job hazard, then one can in theory find spurious positive duration dependence in the new job hazard. In practice, however, the assumption of zero correlation among the unobserved heterogeneity factors in the new job and recall hazards does not seem to be too implausible (Han, Hausman (1990)).

Table 6

COMPETING RISKS ESTIMATION OF RE-EMPLOYMENT PROBABILITIES

	NEW JOB		RECALL		SINGLE RISK	
VARIABLE	B	T-VALUE	B	T-VALUE	B	T-VALUE
CONSTANT	-4.4029	30.1911	-5.6395	34.5406	-4.4160	41.1533
SEX						
FEMALE	-0.0228	0.3083	0.0338	0.4193	0.0069	0.1269
AGE	-0.0216	6.0313	-0.0079	2.5948	-0.0139	5.9728
MARRIED						
YES	0.0790	1.0191	0.2779	3.6868	0.1841	3.4124
HARD TO PLACE						
YES	-0.2776	3.3306	-0.4877	5.1395	-0.3779	6.0556
CHILDREN						
YES	-0.2347	3.1616	-0.2356	3.3720	-0.2410	4.7118
INDUSTRIES						
MANUFACTURING	0.1351	1.7456	-0.1377	1.4347	0.0328	0.5491
CONSTRUCTION	-0.1732	1.7411	0.4196	5.0684	0.1760	2.8195
TOURISM	0.3416	3.6518	0.5616	6.1085	0.4515	6.8758
OTHER	base		base		base	
PREVIOUS EMPLOYMENT DURATION						
- 1 MONTH	0.0574	0.5752	-0.3437	2.2904	-0.0654	0.8004
1 - 3 MONTHS	0.1116	1.1827	-0.0176	0.1449	0.0600	0.8100
3 - 6 MONTHS	-0.0243	0.2068	0.0995	0.7470	0.0233	0.2660
6 - 12 MONTHS	-0.3125	3.3845	0.5080	5.4245	0.1342	2.1328
OVER 12 MONTHS	base		base		base	
PREVIOUS EARNINGS						
SCHILLING PER DAY	0.0004	1.9522	0.0003	1.5165	0.0003	2.2617
PREVIOUS UNEMPLOYMENT						
YES	-0.2547	3.3042	0.2273	2.4975	-0.0473	0.8286
PREVIOUS UNEMPLOYMENT DURATION						
- 3 MONTHS	base		base		base	
3 - 6 MONTHS	-0.1357	1.5219	-0.3003	4.1566	-0.2214	3.8874
OVER 6 MONTHS	-0.2274	2.0244	-0.9374	6.5770	-0.5727	6.6507
LN ALPHA	0.1846	6.9824	0.2605	9.3435	0.2153	11.2195
ALPHA	1.20		1.30		1.24	

	MAX LOGLIKELIHOOD NULL	MAX LOGLIKELIHOOD MODEL	PSEUDO R-SQUARE	CHI-SQUARE	DF
NEW JOB	-4817.5621	-4710.3849	0.0222	214.35	17
RECALL	-4443.1926	-4277.2214	0.0374	331.94	17
SINGLE RISK	-8257.5353	-8131.0948	0.0153	252.88	17

5. Concluding Remarks

The results of this study demonstrate the importance of explicitly accounting for the possibility of recalls in the analysis of unemployment composition and the determinants of unemployment spell durations in Austria. In our sample covering unemployment spells in 1985 we find that recalls accounted for nearly one half of the employment to unemployment to re–employment transitions with the probability of recall being mainly dependent on industry and job characteristics related to seasonal work. In particular, the estimation results from a simple logistic model suggest that temporary layoffs may constitute a regular pattern in the work life of individuals affected by that type of unemployment. We then analyze unemplyoment spell durations in a competing risks framework and, indeed, find significantly different hazards for the two types of risks, new jobs and recalls. While both exit rates exhibit positive duration dependence according to our estimates, the new job hazard is considerably flatter than the recall hazard. The estimated effects of the covariates are also quite different across new jobs and recalls.

Thus, failure to distinguish between different types of layoffs may lead to a serious misperception of unemployment dynamics in Austria. This assertion conforms well with previous findings for the U.S and the Danish labour market and suggests a reinspection of re–employment patterns in other European labour markets as well.

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Appendix

Table A1

COMPETING RISKS ESTIMATION OF RE-EMPLOYMENT PROBABILITIES

Male Population

	NEW JOB		RECALL		SINGLE RISK	
	B	T-VALUE	B	T-VALUE	B	T-VALUE
CONSTANT	-4.5858	24.5127	-5.3235	28.1502	-4.4055	33.4641
AGE	-0.0166	3.6157	-0.0088	2.4694	-0.0116	4.0515
MARRIED						
YES	0.0341	0.3319	0.1708	1.8501	0.1189	1.7128
HARD TO PLACE						
YES	-0.2783	2.4531	-0.9030	6.0791	-0.5467	6.2161
CHILDREN						
YES	-0.1755	1.8063	-0.3046	3.6060	-0.2392	3.6880
INDUSTRIES						
MANUFACTURING	0.1025	1.0215	-0.1625	1.3724	0.0177	0.2335
CONSTRUCTION	-0.2261	2.0875	0.4053	4.5139	0.1543	2.2715
TOURISM	0.2904	1.9368	0.6395	4.7347	0.4765	4.7263
OTHER	base		base		base	
PREVIOUS EMPLOYMENT DURATION						
- 1 MONTH	0.1277	0.9584	-0.3316	1.8990	-0.0231	0.2241
1 - 3 MONTHS	0.2322	1.8735	-0.1678	1.1529	0.0785	0.8451
3 - 6 MONTHS	-0.1183	0.7503	-0.1830	1.0584	-0.1571	1.3490
6 - 12 MONTHS	-0.2483	2.0003	0.4546	4.1529	0.1894	2.3684
OVER 12 MONTHS	base		base		base	
PREVIOUS EARNINGS						
SCHILLING PER DAY	0.0005	2.0791	0.0003	1.0462	0.0004	1.9275
PREVIOUS UNEMPLOYMENT						
YES	-0.3658	3.7117	0.1899	1.7632	-0.1016	1.4379
PREVIOUS UNEMPLOYMENT DURATION						
- 3 MONTHS	base		base		base	
3 - 6 MONTHS	-0.1029	0.9086	-0.3334	4.0988	-0.2480	3.6590
OVER 6 MONTHS	-0.0594	0.4009	-1.0092	5.2655	-0.5289	4.7040
LN ALPHA	0.1796	5.2955	0.3033	8.8781	0.2325	9.6594
ALPHA	1.20		1.35		1.26	

	MAX LOGLIKELIHOOD NULL	MAX LOGLIKELIHOOD MODEL	PSEUDO R-SQUARE	CHI-SQUARE	DF
NEW JOB	-2963.7282	-2901.2888	0.0211	124.88	16
RECALL	-2977.9951	-2842.5170	0.0455	270.96	16
SINGLE RISK	-5296.0807	-5210.5883	0.0161	170.98	16

Table A2

COMPETING RISKS ESTIMATION OF RE-EMPLOYMENT PROBABILITIES

Female Population

	NEW JOB		RECALL		SINGLE RISK	
	B	T-VALUE	B	T-VALUE	B	T-VALUE
CONSTANT	-4.1249	20.6732	-6.1182	24.5116	-4.4035	28.8688
AGE	-0.0266	4.5189	-0.0078	1.4181	-0.0175	4.3868
MARRIED						
YES	0.0873	0.7112	0.4584	3.9196	0.2659	3.1680
HARD TO PLACE						
YES	-0.2738	2.2329	-0.0299	0.2390	-0.1581	1.7942
CHILDREN						
YES	-0.3228	2.6536	-0.0233	0.2067	-0.1892	2.2957
INDUSTRIES						
MANUFACTURING	0.1405	1.1223	-0.1322	0.8675	0.0154	0.1596
CONSTRUCTION	-0.4216	0.5157	0.7131	1.8447	0.2717	0.7534
TOURISM	0.2816	2.2616	0.3127	2.5100	0.3054	3.4314
OTHER	base		base		base	
PREVIOUS EMPLOYMENT DURATION						
- 1 MONTH	-0.0787	0.5028	-0.3903	1.4627	-0.1635	1.2228
1 - 3 MONTHS	-0.0153	0.1015	0.3321	1.6810	0.1040	0.8721
3 - 6 MONTHS	0.2061	1.1518	0.7758	3.7836	0.4342	3.2705
6 - 12 MONTHS	-0.3574	2.5445	0.6051	3.8180	0.0859	0.8614
OVER 12 MONTHS	base		base		base	
PREVIOUS EARNINGS						
SCHILLING PER DAY	-0.0001	0.3106	0.0006	1.5902	0.0002	0.7197
PREVIOUS UNEMPLOYMENT						
YES	-0.0505	0.4025	0.3351	2.1729	0.0791	0.8280
PREVIOUS UNEMPLOYMENT DURATION						
- 1 MONTH	base		base		base	
3 - 6 MONTHS	-0.1894	1.2777	-0.3029	2.1978	-0.2166	2.1325
OVER 6 MONTHS	-0.4586	2.6328	-0.8709	4.3139	-0.6362	4.8333
LN ALPHA	0.2148	4.9179	0.3042	6.0974	0.2438	7.4204
ALPHA	1.24		1.36		1.28	

	MAX LOGLIKELIHOOD NULL	MODEL	PSEUDO R-SQUARE	CHI-SQUARE	DF
NEW JOB	-1852.9142	-1799.1228	0.0290	107.58	16
RECALL	-1462.3080	-1400.5007	0.0423	123.61	16
SINGLE RISK	-2961.2327	-2899.7340	0.0208	123.00	16

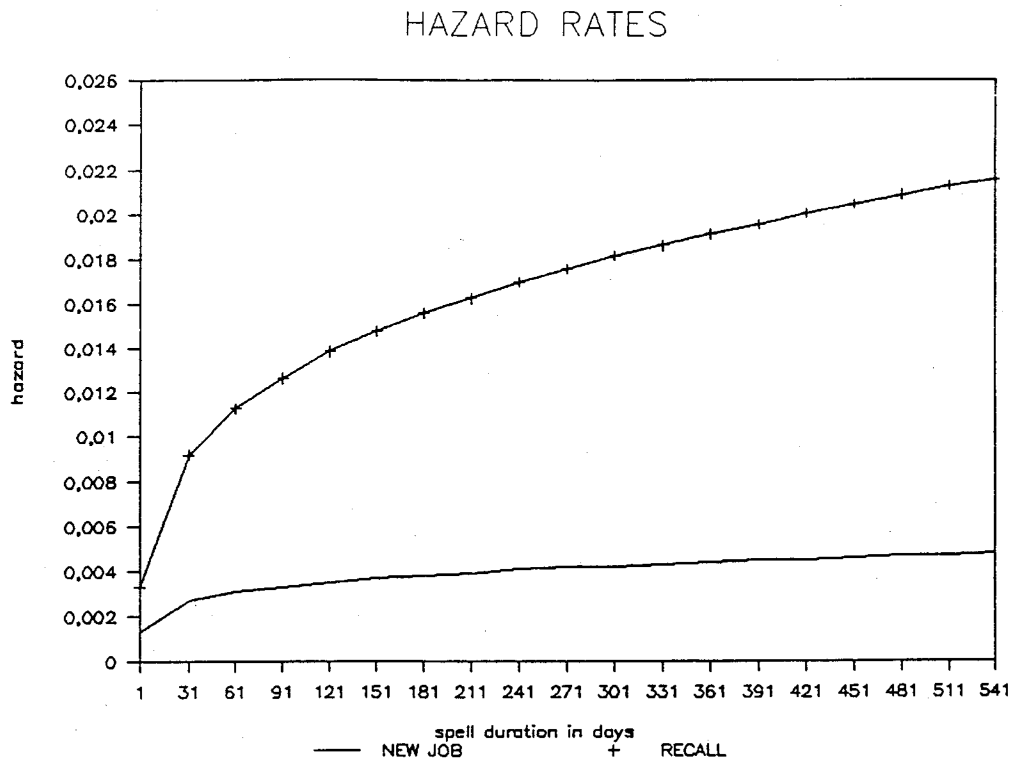

Chart A1

In Chart A1 the following configuration of covariates has been used: male; average age; unmarried; no placement restrictions; no children; construction industry; previous employment duration 6–12 months; average previous earnings; a previous spell of unemployment with a duration of less than 3 months.

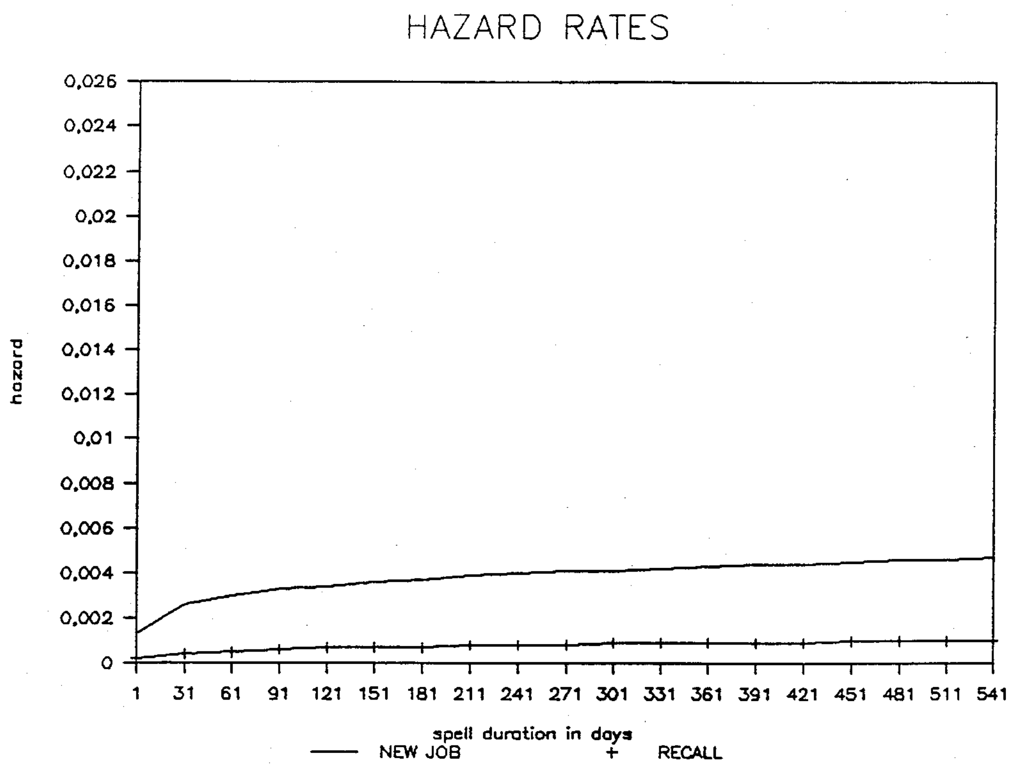

Chart A2

In Chart A2 the following configuration of covariates has been used: male; average age; unmarried; hard to place; no children; previous job other industry than manufacturing, construction or tourism; previous employment duration more than one year; average previous earnings; with previous long term unemployment.