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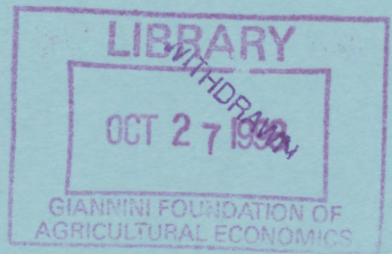
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# Is Job Stability Declining in Germany? Evidence From Count Data Models\*

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## Abstract

The macro evidence of increased adjustment pressure since the early seventies suggests that job mobility should have increased. Hence, retrospective and spell data from the German Socio-Economic Panel are combined in order to test the hypothesis that job stability for German workers declined between 1974 and 1994. Using count data regression models in which we control for labour market experience, various demographic factors, and occupation, we find that job stability did not decrease, but if anything increase, between 1974 and 1994. Our finding suggests that labour market inflexibility is an important factor in explaining the European unemployment problem.

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## 1 Introduction

There is significant interest in the U.S. and elsewhere on whether jobs have become less stable since the beginning of the 1970's. This period has seen large increases in unemployment in most OECD countries. Also, globalisation and technological change have increased the pressure for structural adjustments in the economy. But structural adjustments require a re-allocation of workers, and hence one would expect to see a decline in job stability over the period.

Job stability is generally seen as desirable from a worker's point of view. It shelters the worker from the risk of unemployment and allows for the formation of firm-specific human capital. On the other hand, job stability is harmful from the national point of view if it impedes or even prevents the necessary labour market adjustments. Indeed, it has been argued that the lack of labour market flexibility has been causal for the European unemployment problem.

Whether or not job stability has failed to decrease, and thereby slowed down the necessary labour market adjustments to take place in the face of aggregate trends, is an empirical issue. The previous evidence is mixed. The results depend both on the countries that have been studied and on the data definitions used. Swinnerton and Wail (1995) find a secular decline in U.S. job stability during the 1980s but Farber (1995) reports no such evidence. Similarly, recent papers by Booth, Francesconi and Garcia-Serrano (1997) and Blanchflower and Burgess (1996) document declining job stability for the U.K, while a OECD cross-country comparison (OECD 1997) detects no uniform trends in job stability among developed economies, the U.S. and the U.K. included.

Part of the conflicting evidence can be explained by the use of different datasets, and, closely related, by differences in the way "job stability" is measured. On one side of the spectrum are studies of job and labour turnover that use firm data, starting with the work by Davis and Haltiwanger (1992). Job turnover is the sum of over the year changes in employment levels across all establishments or firms, netting out hiring and firing within the establishment. Labour turnover, by contrast, measures gross hiring and firing. Most studies using firm data

are based on job turnover, since data on labour turnover are much harder to come by.<sup>1</sup>

On the other side of the spectrum are worker based studies. Originating in the seminal paper by Hall (1982), this literature has focussed on the distribution of job tenure and its changes over time.<sup>2</sup> The main econometric technique employed by this literature is duration analysis. We follow this second line of research and consider job stability from the worker's point of view. The novelty of this paper is that it uses job *counts* rather than job durations<sup>3</sup>.

While it is known that the two, counts and durations, are closely related concepts (see Winkelmann, 1995), it is somewhat surprising that a direct analysis of trends in job stability based on the number of job changes of a worker during a given period of time has not been attempted before. This is despite the fact that Hall's (1982) count of 10 lifetime job changes for the average male worker is one of the most frequently cited statistics in the literature on labour mobility. Moreover, data on job counts may be available, where more detailed data on job durations are not. An example for such a dataset is the *German Socio-Economic Panel* (GSOEP) where job counts are available for the period 1974 to 1994.

The objective of this paper is to study the German experience using count data models, and to test whether job stability has declined between 1974 and 1994. For this purpose, we compare the number of job changes during the 1974-84 period to the number of job changes between 1984-94. We investigate differences in job mobility between male and female workers and for workers in different occupations, and how these differences have evolved over time. Secondly, we discuss an econometric methodology that is well suited for analysing such counts of job changes. Count data regression models allow us to control for factors affecting mobility, such as previous labour market experience, marital status and union status.

Section 2 discusses the theoretical background, and section 3 presents the dataset used. Section 4 explains the applied method and section 5 summarizes the major econometric

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<sup>1</sup>There is some evidence that job turnover is about 30 percent of labour turnover (OECD 1997).

<sup>2</sup>Establishment and worker perspectives are not unrelated. For instance, the share of workers in a firm with less than one year of tenure represents the proportion of jobs in which at least one new hire has been made during the past year.

<sup>3</sup>Yet another type of studies looks at the nature of the employment contracts, namely the incidence of temporary or short-term jobs (see Marcotte, 1995)

findings. A summary and conclusions are contained in section 6.

## 2 Background

Job stability and job (or labour-) mobility are two sides of the same coin. High job stability means low mobility, and vice versa. We start with a brief consideration of the two dominant theories of labour mobility, the specific human capital model and the job search/matching model. Both theories imply that the probability that an individual changes jobs declines with job tenure, although the two theories offer different explanations. In the specific capital model the negative correlation between tenure and the probability of leaving arises because firm-specific human capital increases with tenure (Mincer, 1962), while in search models a correlation occurs because the uncertainty over the match quality is reduced over time and only satisfactory job matches "survive" (Jovanovic, 1979). In both cases, there are additional factors that make workers more or less mobile (such as moving costs, marital status, and union membership). A complication arises if differences in individual propensities to move between jobs are due to unobserved factors. For then, those who have longer tenure have a lower probability of leaving not because they have acquired specific capital or found a good match but because they are inherently less mobile.

Individual job spells can be aggregated over the life cycle of a worker into a measure of "long-term" mobility, namely the number of jobs an individual has had over parts of his/her career. We analyse how long-term mobility varies over the life-cycle, across socio-economic sub-groups of the population, and amongst all, over time, i.e. between 1974 to 1984 and 1984 to 1994.

Our main objective is to test the hypothesis that the expected number of job changes increased over the 1980s, or, equivalently, that job stability decreased in Germany. In order to ensure that any observed trend is not caused by changes in other factors (for instance, given the observed experience-mobility profiles, an increased average age of the population would, *ceteris paribus*, lead to an increase of average job stability), we use regression techniques in which we control for selected pre-determined individual characteristics such as years of

previous labour market experience, education, occupational status (in the first job), marital status and union membership.

The importance of previous labour market experience as a determinant of labour mobility is evident both in the theories of mobility and in the previous empirical literature. Hall (1982) estimates that two thirds of the ten lifetime job changes experienced by male US workers occur in the first ten years. Similar evidence exists for the UK and for Germany. Booth, Francesconi and Garcia-Serrano (1997) report that half of the average five UK job changes occur in the first ten years, while Winkelmann (1997) finds that almost half of the average four German job changes fall in this period.

### 3 Data

In order to understand the structure of our dataset, some remarks on the German Socio-Economic Panel (GSOEP) are in order.<sup>4</sup> The panel survey was initiated in 1984. In that year the questionnaire contained a retrospective question on the number of employers during the last ten years. In subsequent annual re-interviews (which were conducted throughout the calendar year), participants were asked whether, and in what month, they had taken a job with a new employer (i) during the *previous* year, or (ii) during the *current* year up to the interview date. Note that:

1. The question was asked for the first time in 1985, where it covered 1984 and part of 1985. The last year in our sample is 1994, covering 1993 and part of 1994. Hence, the effective time period exceeds ten years by a fraction of a year that varies from individual to individual.
2. At most one event could be recorded in each category (i.e. previous year or current year); this imposes a theoretical upper bound of  $2 \times 10 = 20$  (since there are ten years) total changes. For practical purposes, a job change is a much too rare event to be seriously concerned about this upper limit.

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<sup>4</sup>For a general description of this dataset, see Wagner, Burkhauser and Behringer (1993).



3. A job change could have been recorded twice, namely in the current year if it occurred before the interview data and in the next year under the category "changes in the previous year". In constructing our data set we made sure that no double counting occurred.

For both sub-periods, we select German workers who were aged between 16 and 50 and in the labour force in 1974 and 1984, respectively. Note also that the two samples are overlapping in the sense that individuals who qualify for inclusion in the first period may qualify for inclusion in the second period as well. The average number of job changes for men and women in the two periods is given in Table 1.

The average number of job changes during the two ten year periods is relatively low for all groups. This reflects the fact that the modal outcome in these data is zero, or "no job change". Under a Poisson distribution with mean 0.5, for instance, the probability of no change is 61 percent. With a mean of 1, this probability decreases to 37 percent. A comparison over the groups of workers shows that men and women had very similar rates of job mobility in both periods. Moreover, workers at the beginning of their career (i.e., with less than 5 years of experience) had significantly higher mobility rates than all workers. The mobility rates were up to twice as high for recent labour market entrants. This important "life-cycle" effect corroborates the previous findings in the literature.

The main question is whether labour mobility, measured by the number of job changes, did change between the two periods. The above numbers suggest that there was a substantial decline in labour mobility between 1974 and 1994. For male workers, for instance, the average number of job changes over a ten year period decreased from 0.7 in the first period to 0.4 in the second, a decrease that is statistically significant and large. This decline suggests an *increase* rather than a decrease in job stability<sup>5</sup>.

There are several caveats though for such an interpretation. Firstly, we have not controlled at this stage for changes in the average characteristics of workers over time. For instance, the average age in the panel population increased from 32.7 years in the 1974-84 sample to 35.3

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<sup>5</sup>The latter rate of 0.4 is roughly compatible with evidence reported in Zimmermann (1998), based on the same dataset, that the average annual probability of a job change during the period was 3.2 percent.

years in the 1984-94 sample. We expect that this increase caused a decrease in the number of job changes, *ceteris paribus*, since older workers tend to have lower mobility rates. To isolate the specific period effect, we resort to a regression analysis, the results of which are presented below.

Secondly, one should keep in mind that the two mobility measures are based on different types of information. The first mobility measure is based on retrospective information for a ten year interval, while the second measure is constructed from year-to-year information over the ten years of the panel. It is possible, though not necessarily plausible, that people exaggerate their actual number of job changes in the retrospective question, possibly by including changes that did occur before the period proper. In order to shed some light on this issue, Table 2 presents similar statistics, but for two different (shorter) subperiods, namely the five year periods 1984-1989 and 1989-1994. For these two periods, direct spell information is available and no retrospective information is used.

The numbers in Table 2 are roughly comparable with those reported previously for the aggregate 10 year period 1984-94. The sum of the average number of job changes in the two sub-periods exceeds the aggregate rate slightly, which is due to the fact that the 1989-94 figures in Table 2 include workers at the beginning of their career (i.e., those who started working between 1984 and 1989) who were excluded in Table 1 and naturally have higher mobility rates. Table 2 provides no systematic evidence for changes in labour mobility within the 1984-1994 period. While there was a decline in mobility for men, women's mobility was higher in the second sub-period. Both changes are statistically insignificant.

Hence, there is some evidence that the retrospective information for the 1974-84 period might overstate the true number of job changes in the population. However, based on both types of comparisons, it is safe to conclude that there is no evidence for *increased* mobility, or *decreased* job stability. If anything, job mobility in Germany has decreased between 1974 and 1994, for male workers in particular.

## 4 Regression Models for Job Counts

The first choice for modeling count data is the *Poisson regression model* (Winkelmann and Zimmermann, 1995, and Winkelmann, 1997). It is based on two assumptions for the underlying stochastic process. Firstly, the probability of an occurrence during a short interval of time has to be proportional to the length of the interval and constant over time (*stationarity*). Secondly, the probability of an event occurrence has to be independent of previous occurrences (*memorylessness*). Under these assumptions it can be shown that the distribution of the number of events during the period is a Poisson distribution with probability function

$$f(y_i) = \frac{\lambda_i^{y_i} e^{-\lambda_i}}{y_i!} \quad \lambda_i \in \mathbb{R}^+, y_i = 0, 1, 2, \dots \quad (1)$$

As in a loglinear model, the regression component is modeled as

$$E(Y_i|x_i) = \exp(x_i\beta) \quad i = 1, \dots, n \quad (2)$$

where  $x_i'$  is a  $(k \times 1)$  vector of explanatory variables and  $\beta$  a conformable vector of coefficients. The observed values  $y_i$  are realisations from a Poisson distribution with parameter  $\lambda_i = E(Y_i|x_i)$ . Given a sample of independent observations, estimation with maximum likelihood is straightforward. The log-likelihood is given by

$$\ln L = \sum_{i=1}^n [y_i(x_i\beta) - \exp(x_i\beta) - \ln(y_i!)] \quad (3)$$

The first order conditions  $\sum_{i=1}^n [x_i(\exp(x_i\beta) - y_i)] = 0$  state as in the linear model that the residuals  $u_i = y_i - E(y_i|x_i)$  are orthogonal to the explanatory variables. The Hessian matrix  $H = -\sum_{i=1}^n x_i' x_i \exp(x_i\beta)$  is negative-definite except for pathological cases and thus standard numerical algorithms will converge rapidly to a unique maximum of the log-likelihood function.

A particular property of the Poisson distribution is that the variance, conditional on covariates, equals the conditional mean (2). If, beside the  $x$  variables, there are additional

unobserved factors that influence the expected number of job changes, the conditional variance exceeds the conditional mean.<sup>6</sup> Such unobserved heterogeneity can be represented by a random effect in the regression:

$$\bar{\lambda}_i = \exp(x_i\beta + \epsilon_i)$$

The joint distribution of  $y_i$  and  $\bar{\lambda}_i$  is then given by

$$f(y_i, \bar{\lambda}_i | x_i, \beta, \sigma^2, k) = f(y_i | \bar{\lambda}_i) g(\bar{\lambda}_i | x_i, \beta, \sigma^2, k) \quad (4)$$

while the marginal distribution of  $y_i | x_i$  is obtained by integrating (4) over  $\epsilon_i$ . To carry out the integration, one has to assume a specific parametric distribution  $g$ . The most convenient mixing distribution is the *gamma* distribution. Let  $\epsilon_i$  be gamma distributed with  $\Gamma(\alpha_i, \alpha_i)$ . Then  $E(\epsilon_i) = 1$  and  $\text{Var}(\epsilon_i) = \sigma_\epsilon^2 = \alpha^{-1}$ . Further, it can be shown that  $\bar{\lambda}_i = \lambda_i \epsilon_i$  is gamma distributed with mean  $\lambda_i$  and variance  $\alpha^{-1} \lambda_i^2$ . The integration then leads to the negative binomial distribution for  $y_i$ :

$$f(y_i | \alpha, \lambda_i) = \frac{\Gamma(\alpha + y_i)}{\Gamma(\alpha) \Gamma(y_i + 1)} \left( \frac{\alpha}{\lambda_i + \alpha} \right)^\alpha \left( \frac{\lambda_i}{\lambda_i + \alpha} \right)^{y_i} \quad (5)$$

with

$$E(Y_i | \alpha, \lambda_i) = \lambda_i, \quad \text{Var}(Y_i | \alpha, \lambda_i) = \lambda_i + \alpha^{-1} \lambda_i^2 = \lambda_i + \sigma_\epsilon^2 \lambda_i^2.$$

The regression model is completed setting, as above,  $\lambda_i = \exp(x_i\beta)$ .

Given an independent sample, estimation with maximum likelihood is again straightforward. The negative binomial probability function converges to the Poisson probability function for  $\sigma^2 \rightarrow 0$ . This means that the negative binomial and the Poisson model are nested and that the validity of the equidispersion restriction can be tested either by an asymptotic *t*-test for  $\sigma_\epsilon^2$ , or by a likelihood ratio test.

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<sup>6</sup>The conditioning is on observed *x*-es only and not on unobserved components.

## 5 Estimation results

Our regression strategy is as follows: firstly, we pool the observations for the two sub-periods and include a dummy variable "1984-1994" which is one if the observation comes from the second period and zero if it comes from the first period. This is Model 1 which might be labelled a "short" regression. The "long" regression in Model 2 allows for a differential effect by occupational status. By including a full set of interactive terms (four classifications of occupation  $\times$  "1984-1994") we no longer impose that the change in job stability between the two periods is the same for the four occupational groups. In both the long and the short regressions, a positive estimated coefficient on the period variable (or the interacted period variable) indicates that the expected number of job changes went up, i.e. that job stability declined. However, the exact magnitude of this effect cannot be directly derived from the coefficients, since the regression is non-linear and the  $\beta$ 's do not constitute marginal effects.

One possible interpretation is as follows: If all regressors except for one are kept constant while the remaining regressor  $x_k$  is increased by one unit, we obtain a ratio of expected values (or the so called "incidence rate ratio") which is independent of  $x$ :

$$\frac{\exp[\ln t + \beta_0 + \beta_1 x_1 + \dots + \beta_k (x_k + 1) + \dots]}{\exp[\ln t + \beta_0 + \beta_1 x_1 + \dots + \beta_k x_k + \dots]} = \exp(\beta_k)$$

Hence, the exponential of the coefficient for the period dummy tells us by what factor the incidence of job changes in 1984-1990 increased or decreased relative to 1974-1984. An incidence rate ratio of one (corresponding to an estimated coefficient of zero) would indicate no change.

The following regressors are included: previous labour market experience at the beginning of the period (linear and squared); education, which is split into years of schooling (up to secondary school) and post secondary degrees: Apprenticeship, vocational school, and university; marital status; union status; and occupational status in the first job (ordinary and qualified blue collar worker and ordinary and qualified white collar worker, respectively).

The regression results are given in Tables 3 and 4 for men and women, respectively. Models 1 and 2 are estimated in both the Poisson and the negative binomial (Negbin) versions, making

a total of 4 regressions in each table. We start with some general comments on the fit of the models. In all instances a likelihood ratio test leads to a rejection of the Poisson model against the negative binomial model. The chi-squared(1) statistics range from 216.0 to 383.4. The estimated implicit variance term of the multiplicative heterogeneity term is about 1.1 for both men and women. As commonly observed in the presence of overdispersion, the estimated standard errors of the Poisson model are below their negative binomial counterparts although the differences do generally not exceed 30 percent. The parameter estimates are quite robust with respect to the stochastic specification, and most statistically significant effects in the Poisson model remain significant in the negative binomial model.

A comparison of Models 1 and 2 can be based again on a likelihood ratio test which, under the null-hypothesis of no differential effects, is chi-squared(3) distributed with 10-percent critical value of 6.25. We find that we cannot reject Model 1 without occupational interactive effects for any of the models. This is not to say that occupational effects are unimportant. In fact, they are ubiquitous and large. It says, however, that there is no evidence for a *change* in the relative occupational mobility pattern between the two periods. We also conducted likelihood ratio tests of overall significance that uniformly rejected the models with constant only.

We start with a discussion of the male results. The following factors lead to a significant decrease in labour mobility for men: years of labour market experience, being married and member in a union, and being a white collar worker. The importance of labour market experience and occupational status for job mobility is also emphasized in Zimmermann(1998) who uses tabulations from official labour market statistics. As in Zimmermann (1998), education does not matter as much. The educational variables are mostly insignificant. The point estimates are positive for apprenticeship training (i.e., apprentices have higher mobility rates and lower job stability) and negative for non-apprenticeship post-secondary vocational training. This contrasts with results reported in Winkelmann (1996) who finds a negative effect of apprenticeship training. However, those regressions did not control for occupational status while the present regressions do. For women, we find again significant effects of experience and marital status (see Table 4). However, the occupational pattern is weaker. As for men, qualified white collar workers have the lowest mobility among all groups, but the differences

are not always statistically significant.

We now turn to the main question of this investigation: Has job stability of German workers changed between 1974-84 and 1984-94? The answer is yes: the estimated period effect is ubiquitous and large. However, the change was an increase rather than a decrease. For instance, based on Negbin coefficients for Model 1, the incidence of job changes in 1984-1994 fell short of the incidence of job changes in 1974-1984 by an estimated 34 percent for males and 40 percent for females. Recall that, on average, the number of job changes during the ten year periods decreased from 0.7 to 0.4, i.e. by 41 percent, for men. Hence, changes in other factors have contributed somewhat to the decline in mobility, but most of the average change, 34 percent out of 41 percent (or "83 percent"), is unaccounted for by the model and hence a pure period effect. For women, 91 percent of the observed decline in mobility is unaccounted for by the model.

We briefly consider model 2 with occupational interactions. While these interactions are insignificant in the statistical sense, they nevertheless show an interesting pattern for male workers, where mobility rates fell least for qualified white collar workers. In other words, the mobility rates of qualified white collar workers increased *relative* to those of other occupations. Since these workers had the lowest mobility rates in the base period, it follows that occupational mobility rates actually converged somewhat over time.

## 6 Discussion

The original question was whether job stability declined in Germany between 1974 and 1994. Based on micro data on individual level labour market histories from the German Socio-Economic Panel, the answer is negative. We, in fact, find evidence for a reduction in job mobility, and hence an increase in job stability.

An increasing proportion of smaller firms, and an increasing speed of technological change have not had the same effects in Germany as in other countries. Moreover, in contrast to the experience of other OECD countries, Germany has not experienced a substantial decline in unionism prior to the mid-1990's. As a consequence, there was no pressure on job stability

from this factor.

There are other, in particular demographic, trends that might have contributed to the increase in job stability, an ageing labour force in particular. However, a disappearing of traditional family structures (i.e. a decrease in the population fraction of married workers), according to our estimates, contributed to decreased job stability. We find an overall increase in job stability after controlling for the effects of these variables.

The interpretation of our findings is conditional on the assumption that recall information and current information on the incidence of job changes are comparable. Further research is needed to definitely settle whether our estimates reflect a genuine increase in job stability or rather overcounting in recall data. However, we provided additional evidence from spell data only that is not subject to the same caveat. Based on both types of evidence we reject the hypothesis that job stability has decreased over the last two decades.

For those who study the flexibility of German labour markets in the face of accelerating adjustment needs, this finding should be of particular concern, even more so since it reflects the same labour market rigidities that have been documented by Abraham and Houseman (1994) for the German earnings distribution during the 1980's when, against the general trend in the developed world, the overall dispersion of earnings in Germany did actually narrow, mainly due to a compression of the bottom half of the income distribution.



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**Table 1: Number of Job Changes, 1974-1994**

	1974 - 1984			1984 - 1994		
	Mean	S.E.	N	Mean	S.E.	N
1. All Workers						
<i>Men</i>	0.700	0.033	1501	0.410	0.022	1507
<i>Women</i>	0.694	0.046	808	0.388	0.028	967
2. Workers with less than 5 years of experience						
<i>Men</i>	1.282	0.099	280	0.982	0.091	172
<i>Women</i>	1.288	0.117	180	0.622	0.068	204

*Notes:*

1. Individuals are in the labour force and aged 16-50 at the beginning of the period.
2. S.E. are the estimated standard errors.

**Table 2: Number of Job Changes, 1984-1994**

	1984-1989			1989 - 1994		
	Mean	S.E.	N	Mean	S.E.	N
1. All Workers						
<i>Men</i>	0.263	0.015	1507	0.245	0.018	1008
<i>Women</i>	0.240	0.018	967	0.289	0.024	716
2. Workers with less than 5 years of experience						
<i>Men</i>	0.633	0.066	172	0.440	0.067	143
<i>Women</i>	0.401	0.050	204	0.509	0.073	110

Table 3. Regression results for number of job changes: men

Variable	Model 1		Model 2	
	Poisson	Negbin	Poisson	Negbin
Constant	.7849 (.346)	.9403 (.473)	.7248 (.349)	.8564 (.475)
Experience	-.0609* (.009)	-.0661* (.013)	-.0616* (.009)	-.0672* (.013)
Experience squared	-.0001 (.000)	-.0001 (.001)	-.0001 (.000)	-.0000 (.000)
Vocational School	-.1520 (.096)	-.1743 (.127)	-.1475 (.096)	-.1726 (.127)
Apprenticeship	.1367 (.084)	.1224 (.114)	.1437* (.084)	.1270 (.114)
University degree	.0261 (.135)	.0215 (.176)	.0229 (.135)	.0195 (.175)
Years of schooling	-.0282 (.032)	-.0343 (.042)	-.0318 (.032)	-.0382 (.042)
Married	-.2097* (.069)	-.2414* (.098)	-.2062* (.069)	-.2311* (.098)
Union member	-.3430* (.052)	-.3753* (.069)	-.3426* (.052)	-.3756* (.069)
Ordinary blue collar	.3751* (.103)	.3536* (.139)	.4496* (.123)	.4626* (.170)
Qualified blue collar	.2539* (.077)	.2486* (.102)	.3728* (.093)	.4014* (.127)
Ordinary white collar	.0996 (.112)	.0892 (.149)	.1837 (.141)	.2016 (.193)
1984-1994	-.4016* (.051)	-.4161* (.067)		
Ordinary blue collar × 1984-1994			-.3596* (.131)	-.4003* (.176)
Qualified blue collar × 1984-1994			-.4895* (.067)	-.5115* (.089)
Ordinary white collar × 1984-1994			-.3890* (.185)	-.4084* (.242)
Qualified white collar × 1984-1994			-.1699 (.115)	-.1536 (.150)
Log-likelihood	-2982.8	-2791.1	-2979.8	-2788.9
$\sigma^2$		1.122		1.118
Chi-squared (1)		383.4		381.8
Pseudo R <sup>2</sup>	0.174	0.175	0.175	0.176

Notes:

Number of observations is 3008. Standard errors in parentheses. Significant coefficients (at the 0.10 level of significance) are marked with a \*. The chi-squared (1) statistic is for a test of the Poisson against the negative binomial model ( $H_0: \sigma^2 = 0$ ). The Pseudo-R<sup>2</sup>s are based on the deviance measure as proposed by Merkle and Zimmermann (1992) and Cameron and Windmeijer (1996).

**Table 4. Regression results for number of job changes: women**

Variable	Model 1		Model 2	
	Poisson	Negbin	Poisson	Negbin
Constant	.1260 (.425)	.1148 (.576)	.0763 (.427)	.0903 (.576)
Experience	-.0653* (.012)	-.0575* (.016)	-.0661* (.012)	-.0591* (.016)
Experience squared	.0004 (.001)	.0002 (.001)	.0004 (.001)	.0002 (.001)
Vocational School	-.1469 (.113)	-.1944 (.144)	-.1397 (.113)	-.1837 (.144)
Apprenticeship	.0165 (.096)	.0180 (.123)	.0202 (.096)	.0244 (.123)
University degree	-.4074* (.229)	-.4996* (.288)	-.4237* (.230)	-.5019* (.288)
Years of schooling	.0399 (.040)	.0404 (.054)	.0372 (.040)	.0328 (.054)
Married	-.2964* (.074)	-.3015* (.099)	-.2964* (.074)	-.2989* (.099)
Union member	-.1347 (.084)	-.1469 (.109)	-.1398* (.084)	-.1448 (.109)
Ordinary blue collar	.2687* (.116)	.1798 (.150)	.3074* (.138)	.2255 (.184)
Qualified blue collar	.0668 (.130)	.0425 (.167)	.2162 (.170)	.2236 (.227)
Ordinary white collar	.1071 (.083)	.0988 (.109)	.2471* (.107)	.2882 (.146)
1984-1994	-.5196* (.067)	-.5067* (.087)		
Ordinary blue collar × 1984-1994			-.4116* (.144)	-.3690* (.185)
Qualified blue collar × 1984-1994			-.6908* (.225)	-.6880* (.285)
Ordinary white collar × 1984-1994			-.6930* (.113)	-.7159* (.143)
Qualified white collar × 1984-1994			-.3512* (.117)	-.3088* (.152)
Log-likelihood	-1760.6	-1652.2	-1757.8	-1649.8
$\sigma^2$		1.107		1.099
Chi-squared (1)		216.8		216.0
Pseudo R <sup>2</sup>	0.120	0.117	0.122	0.120

*Notes:*

Number of observations is 1775. Also see Table 3.

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\* Copies of these Discussion Papers may be obtained for \$4 (including postage, price changes occasionally) each by writing to the Secretary, Department of Economics, University of Canterbury, Christchurch, New Zealand. A list of the Discussion Papers prior to 1993 is available on request.

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