1 Sickness Insurance and Work Absence

Sweden has a compulsory public sickness insurance system, which replaces forgone earnings due to temporary illnesses.

One of the largest social welfare programs in Sweden. In 2002, expenditure amounted to SEK 48.3 billion, or about 2 percent of GDP. Between 1998 and 2002, the work-absence, and hence the costs of the insurance system, increased by almost 75 percent.

Johansson and Palme (1996) "Do Economic Incentives Affect Work Absence? Empirical Evidence using Swedish Micro Data"

Theoretical model

Departs from a labor supply model, but instead of the hourly wage rate we use the share not replaced by the sickness insurance and the sum of daily non-labor income (including the replacement from the sickness insurance) as income measure.

Utility function

$$u = U(x, L; s)$$

Budget constraint: $x = w(h^* - (1-\delta)t^a) + R$,
h^*- contracted number of hours of work.
$x-$ consumption.
$w-$ daily earnings.
$\delta-$ replacement rate in the sickness insurance.
t^a- time absent.
$R-$ non-labor income.

$h^* = h + t^a$, so h is the actual number of hours of work.
$T = h + t^a + t^l$.

We use a Hausman style utility function, i.e.,

$$U\left(x, t^a\right) = \exp\left\{-\left(\frac{\beta\left(x+\overline{s}\right)}{b - T + \left(t^l + t^a\right)}\right)\right\}\left(\frac{T - \left(t^l + t^a\right) - b}{\beta}\right),$$

which allows us to derive a linear demand function for time absent, i.e.,

$$t^a = h^* - \alpha w\left(1-\delta\right) - \beta\left(R + h^* w\delta\right) - s = h^* + \alpha w^* + \beta y - s.$$

The Slutsky conditions requires that $\alpha < 0$ and that $\beta \geqslant 0$.

Empirical specification

$$t_{ih}^a = h_i^* + \mathbf{x}_i\omega + \varepsilon_{ih} + \zeta_{ih}$$

ζ_{ih} is *iid* random disturbance.

$\varepsilon_{ih} = \theta_i + \varepsilon_h$, where θ_i is an individual effect and ε_h measures the error due to time aggregation of the explanatory variables.

$$I_{ih} = \left\{ \begin{array}{ll} 1, & \text{if } t_{ih}^a > k_{ih} \\ 0 & \text{if } t_{ih}^a \leq k_{ih} \end{array} \right. ,$$

where k_{ih} is a threshold value.

The probability of working a particular day h is

$$\Pr\left(I_{ih} = 0\right) = \Pr\left(t_{ih}^a - k_i \leq 0\right) = 1/\left[1 + \exp\left(h_i^* + \mathbf{x}_i\omega + \varepsilon_{ih} + k_i\right)\right]$$

Estimation

Although we use individual data, they are aggregated in the time dimension. So we have to aggregate the day-to-day over one year, i.e.,

$$V_i = \sum_{h=1}^{N} I_{ih} \sim Bin\left(N, \pi_i\right),$$

where $\pi_i = \Pr(I_{ih} = 1)$ and N is the number of days observed.

The parameters in the model can be estimated using Maximum Likelihood.

Two problems with heterogeneity.

1. Over dispersion. The observed variance is larger than the theoretical one due to heterogeneity. This can be corrected using separate estimates of the variance.

2. Heterogeneity being correlated with the independent variables.

Several reasons for correlation for that this is the case with the cost variable, which is dependent on earnings.

1. Workers who have preferences for being absent may have lower wages. (i) health effect on earning; (ii) less on-the-job training.

2. Compensating wage differentials. Workers who take more risks on the job - have worse work environment and are more absent.

3. Efficiency wage hypothesis. Since work absence is costly for the employer - the could pay some employees above the market wage in order not to be absent.

We use over-dispersed binomial model to account for the first problem. We use $Var(V_i) = \sigma^2 N\pi_i(1-\pi_i)$.

To deal with the second problem, we use a mixing distribution model. A semi-parametric estimator. We add successive constants to the model and maximize the Akaike information criterion (AIC).

Data

1981 Swedish Level of Living Survey. Cross-section data. Dependent variable number of days of work absence during one year (1981). Register data from the National Social Insurance Board on recorded actual transactions. Use a sample of 1,967 blue-collar workers.

We use subjective measures for health status and work environment obtained from the surveys. We use factor analysis to reduce the number of parameters to be estimated in the model.

Results

Mixture distribution model resulted in two groups.

Economic incentives variables have the expected signs. Economic incentives do matter!

Johansson and Palme (2002)

Depart from the same labor supply framework. However, now we use panel data where we have recorded each day of work absence during the two years 1990 and 1991.

In March 1, 1991 there was a reform of the Swedish sickness insurance, where the replacement level was decreased from 90 percent from the first day to 65 % for day 1-3; 80 % for day 4-89 and maintained at 90 % for spells longer than 90 days. We also use the big tax and benefit reform in 1991 to create variation in costs.

Empirical specification

We assume a single index specification and a logistic functional form for the hazards. The hazard to leave work and work absence after l time periods are then equal to

$$\pi_{it}^{01}(l) = \frac{1}{1+\exp(-(a_i + \mathbf{x}_{it}\boldsymbol{\beta}_w + \lambda_w^l))} \text{ and } \pi_{it}^{10}(l) = \frac{1}{1+\exp(a_i + \mathbf{x}_{it}\boldsymbol{\beta}_{wa} + \lambda_{wa}^l)},$$

All the parameters except for the unobserved heterogeneity term, a_i, are allowed to depend on the initial state.

$\lambda_{wa}^l = \lambda_{wa} Dur_{it}^{WA}, \quad l = 8, ..., T$, where $Dur_{it}^{WA} = \sum_{s=1}^{t-1} \prod_{k=1}^{s} d_{i,t-k}$, and $d_{it} = 1$ if individual i is absent day t. For the work spells we simply assume a linear specification i.e. $\lambda_w^l = \lambda Dur_{it}^{WA}$ where $Dur_{it}^{W} = \sum_{s=1}^{t-1} \prod_{k=1}^{s} (1 - d_{i,t-k}), \quad l = 1, ..., T$.

The main difference is that we are able to control for unobserved heterogeneity using fixed effects.

Johansson and Palme (2005) "Sickness Insurance and Moral Hazard"

"Moral Hazard" - do people change their behavior depending on the insurance contract?

Analyzes also the 1991 reform.

A for practical purposes attractive feature of the reform is that these cost changes are independent of the income tax rate. This means that the costs defined above are not affected by the 1991 income tax reform. The reform has the following implications for the direct cost and cost of returning to work:

i. There is an unambiguous *direct* cost increase in beginning a work absence period due to the reduction in the compensation level for the first 90 days of the period.[1]

ii. For work absence spells of less than 91 days, there is an ambiguous effect of the reform. First, there is the increased *direct* cost of continuing a spell. Second, there is an increased cost of *returning* to work.

iii. For absence periods longer than 90 days, there is no change in the direct cost of absence. There is, however, a 25 percent point increase in the cost of returning to work. The reform, thus, implied an unambiguous *decrease* in the relative cost to remain absent in such spells.

Since the unambiguous implications (i) and (iii) work in opposite directions the a priori effect of the reform on the prevalence of work absence is ambiguous.

Two different estimation strategies:

Differences-in-differences estimator, where the difference between January/February and March/April 1990 is used as a "control". Formally, the estimator is defined as

$$DD = (\overline{m}_{jf}^{91} - \overline{m}_{ma}^{91}) - (\overline{m}_{jf}^{90} - \overline{m}_{ma}^{90}), \tag{1}$$

[1] One can note that the reform also, not only gave an unambiguous relative cost increase, gave an unambiguous cost increase absolutely.

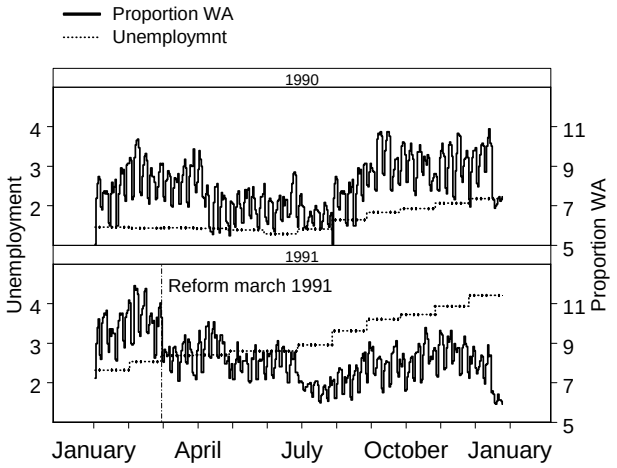

Figure 1: Prevalence of work absence in the sample and the monthly average unemployment rate in 1990 and 1991.

where $\overline{m}_{ma}^{91}$ and $\overline{m}_{ma}^{90}$ are the mean prevalence in March and April 1991 and 1990, respectively and $\overline{m}_{jf}^{91}$ and $\overline{m}_{jf}^{90}$ are the mean prevalence in January and February 1991 and 1990, respectively.

Figure 2 shows the samples day-by-day work absence rate for the period January to May 1990 (left panel) and the corresponding period in 1991 (right panel) including the date of the reform in March 1. It is apparent that there is a discontinuous shift in the work absence rate around the date of the reform, which did not take place around that date the previous year. The overall level of work absence is, however, somewhat higher in 1991.

Table 1 shows the estimated components and the overall result from the differences-in-differences estimator specified in equation (1) both for the males and females and separately for the genders. The results show a significant decrease in the work absence rate in the sample as a result of the reform. The effect is much stronger for the males than for the females.

To estimate the effect of the testable implication under (i) and (iii) above we use (discrete time) Cox regression models in the empirical analysis. For

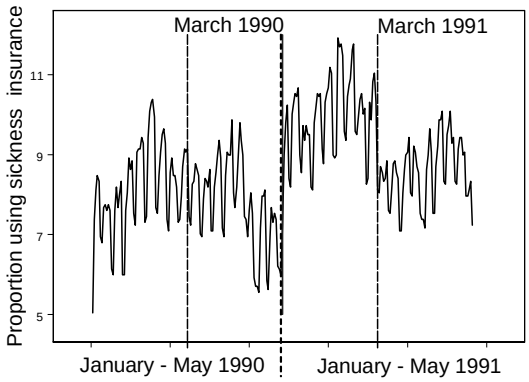

Figure 2: Daily work absence rate January 1 to April 30, 1990 and January 1 to April 30, 1991. Reform date in March 1, 1991 is marked. Males and Females.

Table 1: Estimated means and standard errors (s.e.) for prevalence of work absence for different sub-periods.

Jan/Feb 1990	Mar/Apr 1990	Jan/Feb 1991	Mar/Apr 1991	DD
		Males and Females		
8.13	7.75	9.94	8.62	0.93
(0.15)	(0.14)	(0.14)	(0.11)	(0.27)
		Males		
7.31	7.16	8.40	6.69	1.56
(0.16)	(0.13)	(0.14)	(0.11)	(0.27)
		Females		
9.07	8.41	11.68	10.81	0.21
(0.16)	(0.18)	(0.18)	(0.15)	(0.33)

Table 2: Discrete-time Cox proportional hazard model estimates ($Est.$) and standard errors ($s.e.$) of the effect of the sickness insurance reform on incidence of work absence.

	Males				Females			
	$Est.$	$s.e.$	$Est.$	$s.e.$	$Est.$	$s.e.$	$Est.$	$s.e.$
I^R	-0.316	0.005	-0.310	0.007	-0.211	0.005	-0.240	0.007
Unemployment	-		-0.114	0.010	-		-0.077	0.003
Unemployment^2	-		0.022	0.002	-		0.018	0.001
County factor	No		Yes		No		Yes	
Log likelihood	-9668.6		-9646.6		-9929.2		-9892.9	
$\chi^2(25)$; $p-value$	44.0; 0.01				72.6; <0.001			

Note: χ^2 statistics and $p-value$ for likelihood ratio test of joint significance of local labor market unemployment rate and county factor.

the incidence we use the following specification:

$$\lambda_1(t) = \lambda_0(t)e^{\delta I^R}, \tag{2}$$

where $\lambda_0(t)$ is the baseline hazard (i.e. the hazard before the reform) and I^R is a step function, taking the value 1 after the reform and 0 before.

For the absence spells we use the following specification:

$$\lambda_1(t) = \lambda_0(t)e^{I^R(1-3)\beta_1 + I^R(4-7)\beta_2 + I^R(8-90)\beta_3 + I^R(91-)\beta_4}, \tag{3}$$

where $I^R(j-k)$ are impulse functions, such that $I^R(j-k) = I^R\mathbf{I}(j \leq t \leq k)$ where $\mathbf{I}(.)$ takes the value one if the argument within the parenthesis is true, and $I^R(91-) = I^R\mathbf{I}(91 \leq t)$ is a step function. We also include $\mathbf{I}(1-3)$, $\mathbf{I}(4-7)$ and $\mathbf{I}(8-90)$ in the specification. The interpretation of β_1, β_2, β_3 and β_4 is thus the change in hazard rate caused by the reform.

Table 3: Discrete-time Cox proportional hazard regression estimates ($Est.$) and standard errors ($s.e.$) of the effect of the reform on the duration in work hazard (hazard of ending a work absence spell).

	Males				Females			
	$Est.$	$s.e.$	$Est.$	$s.e.$	$Est.$	$s.e.$	$Est.$	$s.e.$
$I^R(1-3)$	0.062	0.029	0.099	0.033	0.068	0.028	0.077	0.032
$I^R(4-7)$	0.015	0.300	0.040	0.045	0.050	0.031	0.074	0.046
$I^R(8-90)$	-0.022	0.021	0.006	0.029	-0.101	0.020	-0.095	0.027
$I^R(91-)$	-0.127	0.030	-0.100	0.036	-0.639	0.027	-0.591	0.035
Unemployment			0.150	0.028			-0.013	0.027
Unemployment^2			-0.032	0.004			0.017	0.042
County factor	No		Yes		No		Yes	
Log likelihood	-4362.5		-4350.8		-4256.7		-4241.9	
$\chi^2(25)$; $p-value$	23.32; 0.57				29.6; 0.27			

Note: The baseline hazard is specified as piecewise constant. Indicators for 1-3 days, 4-7 days and 8-90 days in a spell are also included in the specification. χ^2 statistics and $p-value$ for likelihood ratio test for joint significance of local labor market unemployment rate and county factors.

Table 4: Estimates of elasticity with respect to the cost of being absent on the incidence and duration of work absence.

	Males		Females	
	$\widehat{\varepsilon}$	$s.e.$	$\widehat{\varepsilon}$	$s.e.$
Incidence	-0.93	0.02	-0.72	0.03
Duration, 1-3 day spells	0.23	0.11	0.25	0.11
Duration, 4-7 day spells	0.14	2.77	0.47	0.30
Duration, 8-90 day spells	-0.20	0.19	-0.87	0.16

Note: Standard errors (s.e.) are obtained using Gauss approximations.

Henrekson and Persson (2004) in JOLE.

Aggregate annual time series data for the period 1955 to 1999.

Panel data 1983 to 1991 county level quarterly data.

$$S_t = \alpha_0 + \sum a_i X_{t-i} + \beta * REF_t + \sum \gamma_i S_{t-i} + \varepsilon_t$$

Pettersson Lidbom and Skogman Thoursie (2006). Analyzes 1987 reform of sickpay insurance.

Skogman Thoursie (2004)

Is the rate of sickness insurance utilization higher when there is a televised sporting event?

Peter Skogman Thoursie investigates this for the Obersdorf cross-country skiing competition and the winter Olympic Games in Calgary. He uses differences in preferences between males and females to follow sporting events on tv to identify the effect.

Differences-in-differences model

$$\delta = E\left[Y_{ip}^M - Y_{ib}^M \mid d^s = 1\right] - E\left[Y_{ip}^F - Y_{ib}^F \mid d^s = 1\right]$$

Effects of screening and control:

Hesselius P., P. Johansson & L. Larsson (2005) "Monitoring sickness insurance claimants: evidence from a social experiment"

Estimates the effects of an experiment in the Swedish counties Gothenburg and Jämtland of postponing the first formal point of monitoring during a sickness absence spell, a requirement for a doctor's certificate, from day eight to day fifteen. All insured born on an even date were asked to show a doctor's certificate after two weeks, whereas insured born on an uneven date had to show one already after one week.

Work absence and employment protection.

Ichino, A. and R. Riphahn (2005) "The Effect of Employment Protection on Worker Effort. A Comparison of Absenteeism During and After Probation"

Uses data from a large Italian bank and investigate to what extent work absence increase after the probation period, when the workers are covered by a stricter employment security legislation.

Assar Lindbeck, Mårten Palme and Mats Persson (2006) "Job Security and Work Absence Behavioral and Compositional Changes in a Natural Experiment"

Four reasons to why employment security legislation may affect work absence:

Behavioral effect:

- If workers are not protected by employment security legislation, they may avoid to be absent since they may fear to loose their employment.

Compositional effects:

- Lower firing costs. Easier for employers to get rid of workers with a high work absence rate with less rigorous employment protection.
- Lower hiring costs. It is less costly for an employer to hire a worker with a high work absence. Leads to a compositional effect towards higher work absence from less rigorous employment protection.
- Employment protection makes it more costly for the worker to give up an employment. "Seniority Capital". Employers with a bad match on the labor market, which in itself may lead to a high work absence rate, may keep their employment because the cost to give it up is too high.

This paper:

- Analyzes the effects of a reform of the job security legislation in Sweden implemented in January 1, 2001.
- Before the reform: Strict turning rules. Principle of last-in-first-out.

- After reform: Two employees excluded from the from the turning rules in firms with up to 10 employees.
- Natural experiment with an experimental group and a control group.

Three DD estimators:

- OLS - overall effect.
- Fixed effects - conditional on the individuals in the sample. Isolates the behavioral effect.
- IV. We use whether or not the individual worked in 1999 in a firm subsequently assigned to the reform as instrumental variable. "Intention to treat parameter".

	Individuals working pre-reform (in1999) in firms with1 or 11-25 employees	Individuals working pre-reform (in1999) in firms with 2-10 employees
Individuals working post-reform infirms with 1 or 11-25 employees	Group 1 "Stayers"	Group 2 "Leavers"
Individuals working post-reform in firms with 2-10 employees	Group 3 "Entrants"	Group 4 "Stayers"

Results:

- Significant overall effect on about 0.25 days reduction in work absence (about 4 percent).
- Behavioral effect of about similar magnitude as the overall effect.
- Counteracting effects: "Entrants" in firms affected by the reform had a higher propensity to be absent. This effect counteracted by the effect from leavers that also had a higher propensity to be absent.