Assessing the Effect of Public Policy on Worker Absenteeism*

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Abstract

The effect of economic incentives on worker absenteism is analyzed using panel data on work absence behavior for each day during 1990 and 1991 (i.e. 730 observations in the time dimension) for a representative sample of 1,396 Swedish blue collar workers. During the observed time period, a major reform of the sickness insurance as well as a tax reform were implemented, both of which affected the worker's cost of being absent from work. We differentiate between the dynamic dependence conditional on whether the worker is in the work absence state versus in the work presence state. We also control for unobserved heterogeneity. The results show that the cost of being absent has a significant effect on work absence behavior. They also show the importance of considering unobserved heterogeneity when modelling individual work absence behavior.

1. Introduction

The effect of tax and welfare policies on the contracted number of hours of work in the economy has been extensively analyzed in the empirical labor supply literature (see e.g. Blundell & MaCurdy, 1999, for an overview). However, as is well known, there are several additional aspects of the relationship between public policy and work effort in general. One such aspects is the effect on work absence. As shown in Barmby et al. (1999), unplanned work absence accounts for a sizable share of the aggregate number of hours of work in industrialized countries, although there is much heterogeneity between different countries in this respect. Income taxes and sickness insurances, which in most countries are regulated by the state,[1] do, of course, affect individual economic incentives for being absent from work.

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[1]See Kangas (1991) for an overview.

In this study, we empirically examine how economic incentives in general affect individual work absence behavior. To do this, we estimate an econometric model using panel data on individual work absence for each day during the period January 1, 1990, to December 31, 1991, i.e. 730 observations for each individual included in the panel. Both the Swedish sickness insurance and the income tax systems were extensively reformed during the time covered by the data. In short, the reform of the sickness insurance, implemented in March 1 in 1991, implied a quite dramatic decrease of the replacement level in the sickness insurance, and the tax reform, implemented in January 1 in 1991, a radical cut in the marginal tax rates. Both these reforms created variation in the individual cost of being absent from work and facilitates thereby estimation.

The average number of days per insured individual compensated by the sickness insurance decreased from 24 in 1990 to 22.5 in 1991.[2] This decrease in rate of work absence, which can also be seen in our sample, supports the hypothesis that the cost of being absent matters. However, in 1991, the Swedish economy entered a recession, with the unemployment rate increased from 2.1 percent in December 1990 to 4.1 percent in December 1991. That is, the composition of the labor force changed and the recession in itself may have had an effect on work absence.

The construction of the Swedish sickness insurance made it possible to develop a unique data-set for studying work absence behavior. The sickness insurance in Sweden is both compulsory and national, i.e. the insurance is administrated by the National Social Insurance Board (NSIB), whose rules apply throughout the country. This means that we are able to get reliable register information on the individuals daily utilization of the sickness insurance. These registers have been matched with a large micro data set, The Swedish Level of Living Survey (SLLS), which consists of a random sample of the Swedish population between ages 16 and 74. Consequently, we are able to obtain detailed information on individual economic conditions, such as income, as well as the cost of being absent from work, together with information on work absence during each day in the two years 1990 and 1991, for a representative sample of Swedish workers. The final sample consists of 1,396 Swedish blue collar workers (738 men and 658 women).

By modelling individual work absence behavior, we may separate how different changes over the observed period of time affected individual work absence behavior. Using a reduced form modelling strategy, allowing for state dependent parameters (where the parameters measuring the effects of economic incentives are not restricted to be the same in work and work absence), we may separate how different changes affected the individual work absence behavior during the observed period. The model used also allows for different duration dependence in the work and work absence state, respectively.[3] The advantage with this model

[2] *National Social Insurance Board, Facts on the Swedish Social Insurance 1992.*

[3] The observed dynamic dependence are thought to be due to differences in health and not due to individuals planing their work absence behavior more than one day in advance. We

compared to a conventional hazard regression models is that we allow for the same (fixed) individual effects in both states (work and work absence) when controlling for unobserved heterogeneity. As we will argue below, individual preference heterogeneity is both likely to be present and affect the inference in this application. Using the fact that fixed effects estimator are consistent in time (T), the panel data-set, with the relatively large number of observations in the time dimension ($T = 750$), allows for credible estimation.

Several interesting results emerged from this study. The estimates from the econometric model confirms that differences in the cost of being absent indeed have an effect on work absence behavior. When the model is used to simulate the reform of the sickness insurance, it is found that both the incidence and the duration of work absence spells decrease when the cost of being is increased. The results also show that unobserved preference heterogeneity is empirically important: not considering preference heterogeneity by using fixed effects in the estimation is shown to affect the inference from the model.

The remainder of the paper is organized into five main sections; the next section describes the Swedish sickness insurance and the income tax system; Section 3 presents the data and gives descriptive statistics on individual work absence behavior; Section 4 describes the econometric modelling and the estimation; Section 5 discusses the results from the estimation and the results of a simulation of the 1991 reform of the sickness insurance; and, finally, Section 6 gives some concluding remarks.

2. National Sickness Insurance and Income taxes[4]

2.1. Sickness Insurance

The national sickness insurance (SI) is financed through payroll taxes levied on wages. All employees whose employers pay payroll taxes are insured by the sickness insurance. The SI is regulated in a separate law[5] and replaces forgone earnings due to temporary illnesses up to the social security ceiling of 7.5 "basic amounts" (BA). Most social insurances in Sweden are connected to the BA. Although the BA is politically determined it has historically followed the Consumer Price Index (CPI) very closely (see Palme and Svensson, 1998). In 1995, about 7 percent of all insured workers had an income exceeding the social security ceiling.[6]

A worker who has lost her ability to do her regular work due to temporary health deficiencies, is entitled to compensation from the SI. To judge whether the individual's health status meets the requirements is entirely left to the insured

acknowledge the fact, however, that some regressors in the model may be predetermined and endogenous, i.e. depending on both previous work absence behavior and unobservables.

[4] A extensive description of Sweden's tax and social insurance systems is provided in Aronsson and Walker (1997).

[5] "Lagen om allmän försäkring".

[6] See National Social Insurance Board (1997).

worker for the first six days in the sickness insurance spell. However, for continued compensation, the worker needs a certificate from a physician from the seventh day of the work absence spell. This certificate needs to be renewed on the 29th day, and after that each month. The Local Social Insurance Office is responsible for controlling abuse of the insurance. However, the control for abuse on the first seven days of a sickness spell is very light. This is partly due to the fact that it is very hard, even for a experienced physician, to decide whether or not a worker is entitled to compensation from the insurance.

The replacement level, the share of the worker's earnings that is compensated from the SI, has been changed on several occasions in recent time. Let us briefly review the changes relevant for this study.

In 1987, the SI went through a major reform. In this reform, the replacement from the insurance was changed to be more closely related to the insured worker's regular work hours, i.e. sickness benefits can only be claimed for these work hours. This rule prevails during the entire time period considered in this study. At the same time, the replacement level was substantially increased to 90 percent of foregone regular earnings (below the social security ceiling) from the first day in each sickness spell.

The insured worker is also entitled to compensation from the SI during vacation. However, in this case, a certificate from a physician is needed already from the first day of the sickness spell. The worker is then able to save the number of vacancy days corresponding to the days compensated by the SI.

On March 1, 1991, the replacement level was decreased to 65 percent of the insured worker's income for the first three days of the sickness spell and to 80 percent from day 4 through 89. From day 90 and on the replacement level remained at 90 percent.[7] This reform, as mentioned in the introduction, occurred during the time period covered by our panel data set and constitutes the main source of variation in the cost of being absent.

In 1992 the responsibility for the SI, during the first ten days in each spell, was taken over by the employer. As short term sickness spells after this reform are no longer administrated by the NSIB, it is not possible to obtain as detailed data as we use in this study, after 1991.

Several groups in the Swedish labor market have negotiated sickness insurances in addition to the compulsory national insurance. Furthermore, most white-collar workers have local agreements with their employers for sickness payment. These agreements can generally not be recovered from our data-set. Therefore, we have restricted our analysis to blue-collar workers. The blue-collar worker's sickness insurance plan ("AGS"), which is negotiated by the blue-collar worker's

[7] This reform was one of several means taken by the government in order to cut the budget deficit. There were several motives for the design of the reform. For instance, the social democratic government was restricted by a pledge given in the campaign preceding the 1988 election that the SI should continue to compensate forgone earnings from the first day of all sickness spells. Furthermore, the motive for letting the compensation level remain at 90 percent, for the period in the spell exceeding 90 days was one of income distribution.

union ("LO"), does not only cover actual members of the trade union, but all workers whose workplace is covered by collective agreements between the trade union and the employers confederation ("SAF"), corresponding to about 98 percent of all employed blue collar workers.

AGS can only be claimed for sickness spells longer than 7 days, i.e. sickness spells that requires a certificate from a physician. On the other hand, the compensation is payed, from day one to seven retroactively, when AGS is admitted. The AGS plan was substantially changed in the 1991 reform of the National SI. Before the 1991 reform, the replacement was 8.5 percent of the worker's benefit from the National SI from day 1 to day 14; from day 15 to day 90 it was changed to be between 6 and 21 SEK per day, depending on the size of the benefit from the National SI; after day 90 it is changed to 6 SEK per day irrespective of the worker's sickness benefit. After the reform, the AGS was changed to 15.4 percent of the worker's sickness benefit for the first three days in a spell; to 12.5 percent between day 4 and 90; from day 91 AGS was abolished. The AGS benefits were also made taxable in 1991.

2.2. Income Taxes and Benefits

Sweden has an integrated income tax system. Taxes are payed both to the national and to the local governments. The national government determines the tax base for both these taxes. The local government tax is proportional and is determined by each of Sweden's 288 local governments, although some income redistribution does take place between high and low income municipalities. In 1991, the local government tax rate varied between 26.87 to 33.48 percent while the mean tax rate was 30.3 percent.

The income tax system was radically changed in the tax reform of January 1991. The highest marginal tax rate was reduced from the local government tax rate plus 42 percent national tax rate (with a maximum set to 75 percent in combined marginal tax rate) to a 20 percent national tax rate in addition to the local government tax. Most full time wage earners, in the income interval between 70 000 and 170 000 SEK, received a substantial (on average about 20 percent) reductions in their marginal tax rates.

Another important element of the tax reform was a tax base change. Being unified in the pre-reform tax system, it was after the reform divided into *earned* and *capital* income. The tax on all capital income was changed to a flat rate on 30 percent. The tax base was also made substantially broader. Several previously non-taxed fringe benefits were made taxable after the reform. Also, several services were made VAT-taxable. Finally, child and housing allowances were substantially increased. The child allowance, which is the same amount per child, irrespective of the parents income, was increased by about one third. The magnitude of the average increase of housing allowances, which is means tested, was also about one third.

3. Data and Descriptive Statistics

3.1. Data and Measurement of Included Variables

We use data from the 1991 SLLS. In this survey, about 6 000 Swedish individuals, aged between 16 and 74, were interviewed about economic resources and living conditions in a broad sense (see Fritzell & Lundberg, 1994, for a detailed description of this survey). The interviews were carried out in the spring of 1991. Information from three different registers: (i) the NSIB register; (ii) the tax register; and from (iii) a register of unemployment benefits provided by the Labor Market Board have been matched with the SLLS survey and the information from the registers has been added to the data base.

The measurement of day-to-day work absence behavior for each individual in the sample, i.e. the dependent variable in the reduced form, is obtained from the NSIB register. This register contains information on the exact dates corresponding to each payment from the SI.[8] Data to construct the cost and income variables are obtained from tax registers and the SLLS survey.[9]

In order to define the cost and virtual income variables, let us first define the worker's daily budget set. L_t represents leisure time and consists of two components: contracted leisure time, t_t^l and time in work absence t_t^a (i.e. $L_t = t_t^l + t_t^a$). We assume that the contracted leisure time is fixed over the time period studied (two years), i.e., $t_t^l \equiv t^l$. The sub-index, t, is an index for each day covered by our data-set, i.e. we allow the preferences to be different between days. The daily budget constraint can then be defined as

$$x_t + (1 - \delta_t)\, w_t t_t^a = w_t h^* + R_t$$

where h^* is the contracted number of daily working hours, R_t is income from sources other than labor, w_t is net hourly wage and $(1 - \delta_t)$ is the share of the income the worker receives when absent. Assuming that the worker maximize a utility function which includes weakly separable consumption of goods, services and leisure time, it is straight forward to obtain the following general demand function for time absent

$$\tau_t^a = f\left(h^*, w_t(1 - \delta_t), R_t + h^* w \delta_t, \mathbf{s}_t, \varepsilon_t\right) = f\left(h^*, c_t, \mu_t, \mathbf{s}_t, \varepsilon_t\right),$$

where $c_t = w_t(1-\delta_t)$ and $\mu = R_t + h^* w_t \delta_t$ are the cost and virtual income of being absent, respectively; $\mathbf{s}_t$ is observable and ε_t unobservable personal characteristics.

The data-set enables us to measure the cost and income variables for the individuals in the sample. The hourly wage rate is obtained by dividing the

[8] Since actual payments are linked to the data we use, the accuracy of the our measure is likely to be very high. According to SAF (1986) the fraction of unplanned work absence, not compensated by the SI is only 2.9 percent.

[9] The SI in Sweden is financed through payroll taxes. The level of the payroll tax is, hence, determined by the utilization of the SI. In this study we (reasonable) assume that the individual acts as if there are no costs of work absence through the payroll tax.

potential annual labor income by the number of work hours (h). To calculate the potential annual income from labor, we have added the share of income from labor not covered by the SI, δ_t, for each day recorded in which the individual has been compensated by the SI.[10]

The cost of being absent is also influenced by the AGS insurance. We assume that the individual makes her decision on work absence on a day-to-day basis. Therefore, we disregard the AGS benefits for the first seven days and only consider it for the cost of being absent from the eighth day in an absence spell. However, after the eighth day, we use the rules for the AGS insurance and calculate the individual benefit from this insurance.

Finally, the cost of being absent is influenced by income taxes. We use the marginal tax rate for each individual corresponding to their taxable income if they had not been absent at all during the year. There is a potential endogeneity problem with this procedure. As labor income is not fully compensated for by the SI, the marginal tax rate may depend on the number of days the individual is absent from work.

To get a measure of R_t, we take the daily average of each individual's annual income; i.e. divide the annual income with 220 (i.e. 365 minus the average number of work free days during a year).[11] If the individual is married or cohabiting, total family income, including observed labor income from the spouse, is summed and divided by two

An inevitable problem, when measuring income using administrative data is how to measure income from real wealth, such as properties like owner occupied homes and durable consumption goods, as we do not have access to reliable estimates on the worth of these items. We will follow the strategy of Statistics Sweden and assume that negative income from capital corresponds to a equal worth income from real wealth.

To measure the number of hours of work, we use the information obtained in the SLLS interviews. We assume that each individual in the sample has the same number of regular work hours as in the spring of 1991 during the entire time period considered. To make the cost and income variables comparable between 1990 and 1991, we deflate the prices and costs in 1991 by the CPI so that all figures are in 1990 prices.

In the econometric model we include the unemployment rate on each worker's

[10] As the sickness insurance only compensate for earnings below the social insurance ceiling, this is only true for labor earnings below this ceiling. The social insurance ceiling corresponded to an annual labor income of 222,750 SEK in 1990 and 241,500 SEK in 1991. However, as all individuals in our sample, like most blue collar workers, do not have income above the ceiling this is not an issue in our sample.

[11] This procedure creates two problems. First, as we take the daily average of annual income, the measure of income is contingent on the work absence behavior during the other days of the year. The second problem is that we have to make the unrealistic assumption that the worker does not have access to credit markets, or is able to make intertemporal consumption decisions, in order to smooth out consumption during the time-period studied.

local labor market in the vector of socio-economic variables, $\mathbf{s}_t$. To measure this unemployment rate we use the monthly unemployment rate in each of Sweden's 24 counties provided by the *Labor Market Board*. One problem with using the monthly unemployment rate is that there are seasonal variations in the unemployment rate. Such variation are unlikely to affect the worker's work absence behavior. To account for these seasonal variations we use multiplicative seasonal components provided by the *National Institute of Economic Research.*[12]

3.2. Sample Selection

As described in the previous subsection, the entire data set consists of about 6 000 individuals. However, in order to obtain the sample used in the estimations, several selections were made. Table 3.1 shows the number of individuals remaining after each of the steps in this selection process.

In the first step, non-responses (about 15 percent of the original sample) are removed, and as the normal age of retirement is 65 years in Sweden, we have restricted the sample to individuals between 20 and 65 years of age in 1991. In the second step, information about the respondents activities the week before the survey interview was conducted were used to exclude unemployed as well as individuals not in the labor force, i.e. students and pensioners (individuals who receive disability or old-age pension). Self-employed, military personnel, and white collar workers were excluded in the third step. That is, the study is restricted to employed blue collar workers. The reason for excluding the groups of individuals, other than blue collar workers, is to limit the heterogeneity arising from differences in the SI systems.

The fourth step in the selection process was to use the information in SLLS on activities during 1990 to exclude 55 individuals who were not employed during more than 4 weeks in 1990. Finally, 197 individuals with "inconsistent" observations on their labor earnings during either 1990 or 1991 were removed.[13]

3.3. Descriptive Statistics

The final sample consists of 1,396 individuals, 738 men and 658 women. From this sample ten individuals (5 men and 5 women) were absent from work the whole period under study, and 158 (97 men and 61 women) were not absent at all. The observations on individuals who did not change state at all can not be used in a fixed effect regression model (see Section 4). Furthermore, four individuals (2

[12] These are obtained from an ARIMA model estimated on monthly unemployment data for the entire nation between 1988 and 1998.

[13] These respondents claimed that they worked more than 16 hours each week and earned less then one BA (29 700 SEK in 1990 and 32 200 SEK in 1991). Possible explanations for these observations can be measurement errors in at least one of the relevant variables or that the person works in the "informal" sector of the economy and is therefore not covered by the SI.

Table 3.1: Number of observations in the sample remaining after each step in the sample selection.

	Number of observations		
	Men	Women	Total
Respondents, age 20-65	2,191	2,106	4,297
Employed and in 1991	1,140	1,256	2,396
Blue collar workers in 1991	833	815	1,648
Workers not laid-off or unemployed more than 4 weeks during 1990	812	781	1,593
Taxed annual earnings that are consistent with the stated regular number of hours of work	738	658	1,396

men and 2 women) was deleted since they do not change state after their first, left censored work absence spell is removed. We now have 1,224 individuals who have variation in the work absence behavior over the studied time period.

Table 3.2 compares the work absence rates for the 1,396 workers with the final sample of 1,224 individuals divided in four different age groups. It also shows in which subsamples the individuals excluded in the final step were located. It is evident from Table 3.2 that the work absence rate is somewhat higher for women in all age groups except in the oldest group, aged between 50 and 65. Also, the rate of work absence remains almost constant for both gender groups until the age group 50 to 65, at which it increases substantially, with the largest change for men. Finally, the effect of the reform of the sickness insurance can be studied in Table 3.2: the work absence rate in the entire sample decreases from 8.4 to 8.0 percent. It is noticeable that the work absence rate actually increases in the oldest age group, for both men and women, in the oldest age group.

Table 3.3 reports descriptive statistics for the variable measuring the daily costs of being absent from work and the virtual income variable. Comparing the average cost of being absent in 1990 and 1991, it can be seen that there is marked increase in 1991 in all subsamples. This effect is expected due to the reform of the sickness insurance. It is also evident that the virtual income on average is somewhat lower in 1991. This difference is due to the higher compensation levels in the SI as well as marginal tax rates in 1990, which make the slope of the budget set steeper. Since the men on average have higher labor earnings the fact that the cost of being absent is higher for men than for women is understandable. Women have on average higher virtual income. At first sight this may seem non-intuitive, however, most workers in our sample are married or cohabiting; i.e. their spouses earnings are included in their virtual income but only the share of their own earnings covered by the SI. Finally, Table 3.4 reports some socio-economic characteristics of the sample.

Table 3.2: Average rates of work absence and number of observations (n) in different groups by age, gender, before and after the SI reform in 1991. Figures in parentheses applies to the sample used in the estimation, i.e. the sample after removing the individuals who did not change state (work or work absence) during the entire two years time period covered by the panel.

	Average share absent from work		
	Men	Women	Total
Age Group		20-29	
Before Reform	.062 (.067)	.090 (.097)	.073 (.079)
After Reform	.047 (.051)	.063 (.069)	.054 (.058)
Total	.056 (.061)	.079 (.085)	.065 (.071)
n	197 (182)	134 (124)	331 (306)
Age Group		30-39	
Before Reform	.067 (.077)	.085 (.090)	.075 (.083)
After Reform	.057 (.065)	.073 (.077)	.064 (.071)
Total	.063 (.072)	.080 (.085)	.071 (.078)
n	177 (155)	147 (139)	324 (294)
Age group		40-49	
Before Reform	.073 (.075)	.093 (.090)	.083 (.083)
After Reform	.059 (.059)	.091 (.088)	.076 (.074)
	.067 (.068)	.093 (.089)	.080 (.079)
n	190 (157)	195 (179)	385 (336)
Age group		50-65	
Before Reform	.104 (.107)	.101 (.104)	.103 (.105)
After Reform	.131 (.140)	.118 (.124)	.124 (.132)
Total	.115 (.121)	.108 (.112)	.112 (.116)
n	174 (140)	182 (148)	356 (288)
Age group		All	
Before Reform	.076 (.080)	.093 (.095)	.084 (.087)
After Reform	.073 (.076)	.089 (.090)	.080 (.083)
Total	.075 (.079)	.091 (.093)	.082 (.086)
n	738 (634)	658 (590)	1,396 (1 224)

Table 3.3: Daily virtual income (μ) and cost (c) of being absent in different age and gender groups. Final sample of 1,396 blue collar workers. (s is the standard deviation, Q_1 and Q_3 are the first and third quartiles, respectively.)

	Males				Females			
	Mean	s	Q_1	Q_3	Mean	s	Q_1	Q_3
				1990				
				Age 20-29				
c	3.60	0.97	3.15	4.11	3.40	0.92	2.86	3.82
μ	476.1	112.1	399.5	546.9	524.9	160.1	431.7	641.4
				Age 30-39				
c	4.01	0.94	3.63	4.43	3.40	0.78	2.88	3.84
μ	566.2	167.1	464.6	668.3	629.5	140.8	548.5	710.7
				Age 40-49				
c	4.04	0.82	3.56	4.54	3.65	0.87	3.07	4.00
μ	605.9	168.5	522.2	704.2	644.1	173.2	548.0	735.5
				Age 50-65				
c	4.18	1.11	3.57	4.43	3.47	0.94	2.92	3.81
μ	613.6	137.0	506.8	695.6	616.1	172.0	510.9	694.7
				All				
c	3.94	0.98	3.46	4.36	3.49	0.89	2.94	3.91
μ	560.6	157.2	452.0	660.8	608.6	168.6	502.4	700.1
				1991				
				Age 20-29				
c	15.54	3.70	13.85	17.70	13.80	4.37	11.59	15.20
μ	352.3	113.4	274.6	424.9	384.4	119.2	287.5	470.8
				Age 30-39				
c	15.86	3.53	14.60	17.91	14.80	3.64	13.29	16.71
μ	425.5	109.7	331.3	502.7	504.3	150.2	429.2	558.9
				Age 40-49				
c	16.27	3.77	15.09	18.29	15.27	4.26	13.23	17.70
μ	463.7	116.3	364.1	540.5	505.3	171.5	407.2	565.1
				Age 50-65				
c	16.34	5.789	14.42	18.22	15.01	5.015	13.20	17.29
μ	439.4	107.8	359.2	503.9	449.4	162.0	356.9	516.0
				All				
c	15.97	4.235	14.49	18.10	14.79	4.38	12.98	16.86
μ	417.0	119.8	321.5	503.9	465.6	161.2	377.7	531.8

Note: The quartiles are computed from the individual means for each individual.

Table 3.4: Socio-economic characteristics of the sample.

	Men	Women
Share working in industry sector	0.61	0.11
Share working in service sector	0.39	0.89
Share working "full time" in 1991	0.95	0.53
Share married or cohabiting	0.67	0.76
Share divorced	0.05	0.09
Share with one dependent child	0.19	0.23
Share with two dependent children	0.17	0.21
Share with three or more dependent children	0.07	0.09
Average age	38.5	40.1

3.3.1. Duration in Work and Work Absence Spells

Figure 3.1 shows the Kaplan-Meier (KM) estimates of survival in the work state before and after the reform of the SI.[14] The corresponding estimates for survival in the work absence state is reported in Figure 3.2. Figure 3.1 shows a marked increase in the duration in work due to the reform of the SI. This can be seen in all age and gender groups. The largest effect is for men in the oldest age-group. However, the effect of the reform on the duration in work absence is in general not measurable according to the KM estimates reported in Figure 3.2. An exception is the two youngest female age-groups, where a small decrease in the duration of the spell after the reform is detected.

Figures 3.1 and 3.2 shows that the higher rate of work absence in the female subsample is due to more frequent, rather than longer, work absence spells. However, in the age group 50-65, the average work absence rate is somewhat higher among men. In this age group, which is evident from Figure 3.2, men return to work considerably slower than women.

Finally, by comparing Figure 3.1 with Figure 3.2 it can be seen that the work absence spells are very different from the work spells. As health deficiencies lies behind most work absence spells, the explanation for the observed difference is, of course, that there are different biological processes underlying the transition from work to work absence compared to from work absence back to work. This may be a trivial point, but it is important to have in mind when work absence is modelled econometrically.

3.3.2. Weekday Effects

Since workers can only claim benefits from the SI for lost earnings from regular work hours and most workers do not work on Saturdays and Sundays, there are no incentives for a worker to begin a sickness spell on a work-free day. Therefore, the rate of work absence will differ between different week days. Figure 3.3 shows

[14] In the KM estimation we neglect left censored spells in 1990, i.e., the initial sample can be considered as a flow sample.

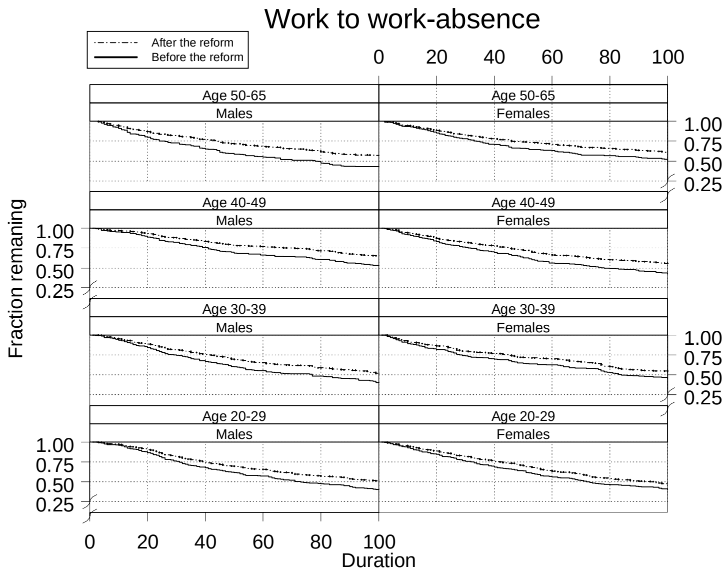

Figure 3.1: Kaplan-Meier non-parametric estimates of survival rates in work spells.

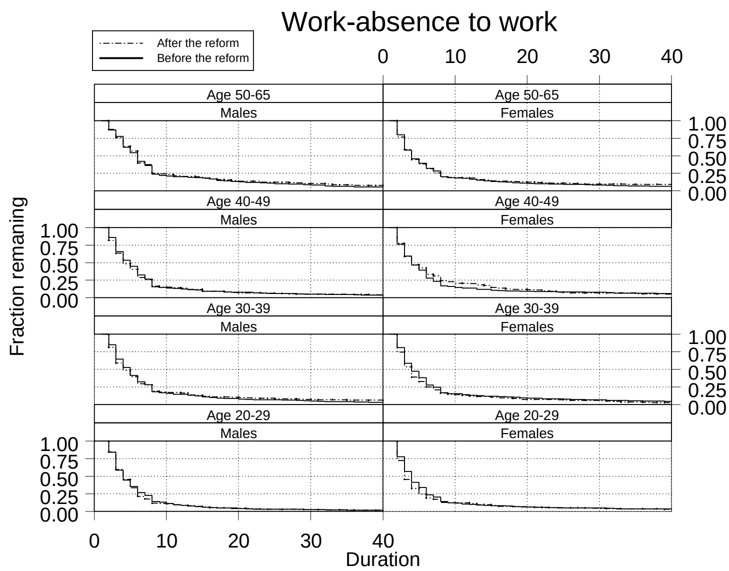

Figure 3.2: Kaplan-Meier non-parametric estimates of survival rates in work absence spells.

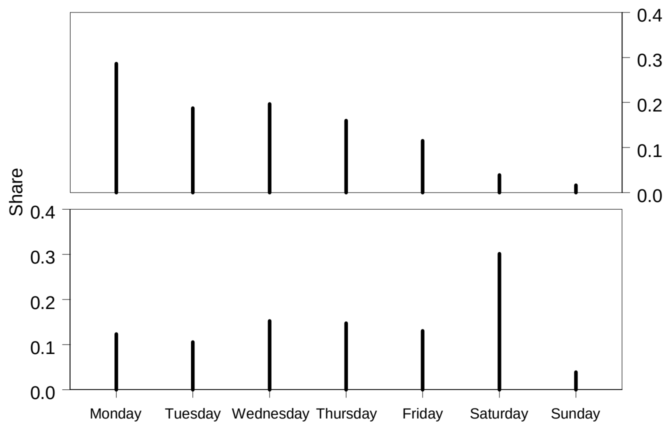

Figure 3.3: Share of sickness spells starting (first panel) and ending (second panel) on different days of the week.

the weekday distribution of starts and endings of work absence spells. The figure shows what can be expected given the rules for the SI, i.e. the relatively large proportion of sickness spells that begin on Mondays reflect the fact that some health problems may accumulate during the weekends. The large proportion of sickness spells that end on a Saturday has a similar explanation, as recoveries from temporary illnesses also accumulate during the weekends.

3.3.3. Seasonal Variation in Work Absence

Since the daylight and outdoor temperature varies considerably over the year in Sweden, a seasonal variation in work absence rates may be suspected. Figure 3.4 shows the average number of reported SI spells between the years 1987 and 1991 by month obtained from aggregate data provided by the NSIB. Two properties of the seasonal differences should be noted. First, work absence varies considerably between different seasons. Second, there is a sharp decline in the reported sickness spells in July. This is because July is the month of summer vacation for workers in Sweden. As describes in the previous Section 2.1, although the insured worker is entitled to compensation from the SI during the vacation, a certificate from a physician is needed from the first day of the sickness spell.

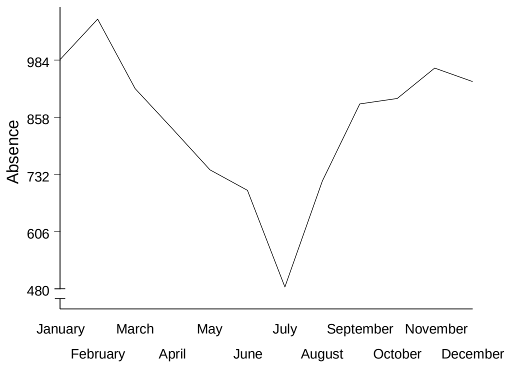

Figure 3.4: Average number (in thousands) of reported sickness spells in different months between 1987 and 1991. Source: National Social Insurance Board, Stockholm.

4. Modelling and Estimation

Several determinants of the daily decision of work attendance have been proposed in the literature. Different disciplines - such as economics, management science, social medicine and applied psychology - have stressed different explanations to observed differences. The economic model in Section 3.1 showed that a general formulation of individual demand for work absence, derived from constrained individual utility maximization, gave contracted hours of work, cost of work absence and virtual income as determinants of the work absence decision. However, the general formulation did also allow for several other determinants.

Although we have a very detailed data set on individual characteristics, it is obvious that there are several determinants of work absence that we do not have access to. Most important is probably general individual heterogeneity in preferences for being absent from work, and - since it is highly plausible that the utility of leisure is considerably higher if an individual suffers from a temporary illness - changes in individual health status over the time-period covered by our data set. We will, therefore, confine the analysis to a very general specification which allow for preference heterogeneity as well as for state dependence. The hazard of leaving the state of work (W) and work absence (WA) conditional on being in the state, l periods is respectively

$$\Pr(d_{it} = 1|W, l) = \pi_i^{01}(c_{it}, \mu_{it}, \mathbf{z}_{it}|\boldsymbol{\theta}_w) = \pi_{it}^{01}(l)$$

and

$$\Pr(d_{it} = 0|WA, l) = \pi_i^{10}(c_{it}, \mu_{it}, \mathbf{z}_{it}|\boldsymbol{\theta}_{wa}) = \pi_{it}^{10}(l),$$

where $\mathbf{z}_{it}$ is a vector of variables likely to affect work absence. $\boldsymbol{\theta}_w$ and $\boldsymbol{\theta}_{wa}$ are state dependence parameters. 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 in these states is then equal to

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

respectively, where $\mathbf{x}_{it}\boldsymbol{\beta}_w = \alpha_{wa} + \gamma_w c_{it} + \psi_w \mu_{it} + \mathbf{z}_{it}\boldsymbol{\gamma}_w$, $\mathbf{x}_{it}\boldsymbol{\beta}_{wa} = \alpha_{wa} + \gamma_{wa} c_{it} + \psi_{wa}\mu_{it} + \mathbf{z}_{it}\boldsymbol{\gamma}_{wa}$, λ_w^l and $\lambda_{wa}^l, l = 1, ..., T$, are parameters that describes the duration dependence.[15] All the parameters except for the unobserved heterogeneity term, a_i, are allowed to depend on the initial state.

The following variables are included in $\mathbf{z}_{it}$: indicator variables for weekdays (Tuesdays is taken as a numerator); indicators for public holidays (HO)[16] and for days between public holidays and Sundays or Saturdays (BH); indicator variables for each month of the year (January is taken as a numerator). In addition to these variables, a variable measuring the local unemployment rate ($UNEM$) is included in $\mathbf{z}_{it}$. As is described in the introduction, the Swedish economy entered a recession in the spring of 1991 and the unemployment rate increased rapidly. Several hypothesizes about the relationship between the unemployment rate and work absence has been suggested (see Bäckman, 1998, for an overview). Most of these hypothesizes use arguments from the efficiency wage literature. Following for example the model of Shapiro & Stiglitz (1984), a high unemployment rate will act as a "worker discipline device". If a worker is caught shirking and dismissed, the worker's cost of finding a new job, and consequently her cost of getting dismissed, will be higher during times of high unemployment. Some form of work absence can be seen as shirking.

The specification of state dependency is a important aspect of this model. We believe that temporary changes in the worker's health status is one of the main determinants of work absence. The day-to-day health status can of course not be observed in the data and is also likely to change slowly over time. The non-parametric analysis (see Figures 3.1 and 3.2) suggest non monotonic duration dependence for the work absence spells, at least so for the first seven days, whereas it seems reasonable to assume that $\lambda_w^l \equiv 0$ for all l. It is hence reasonable to specify the first seven days in a work absence spell unrestricted and, in order to save degrees of freedom, we restrict the following duration dependence to be linear, i.e. $\lambda_{wa}^l = \lambda_{wa} Dur_{it}^{WA}, \quad l = 8, ..., T$, 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 $Dur_{it}^{W} = \sum_{s=1}^{t-1}\prod_{k=1}^{s}(1 - d_{i,t-k}), \quad l = 1, ..., T$.

Another important aspect of the model is that it allows for preference heterogeneity through the individual specific term a_i. As is evident from section 2 the net wage, together with the SI , determines the individual's cost of being absent.

[15] Meghir & Whitehouse (1997) use a model that allow for different parameters depending on state when modelling labor market transitions for older men.

[16] There are about seven public holidays each year.

There are at least four reasons why unobserved preference heterogeneity may be correlated with the individual's wage rate. First, a worker with strong preferences for work absence may choose a job that allows her to be absent from work to the cost of, ceteris paribus, a lower wage rate (see e.g. Allen, 1981). Second, since jobs differ in the cost of absenteeism for the employer it may, thus, be profitable for the employer to pay some employees more, i.e. an efficiency wage, in order to elicit them to have low work absence rate (see e.g., Shapiro and Stiglitz (1984); Wiess (1985)). Third, when there is economic returns to on-the-job training then individuals with strong preferences for work absence will, everything else equal, earn less. And last, it is a well known empirical fact that workers with bad health on average have a higher work absence rate than workers without health problems (see e.g. Broström et al., 1998). For some jobs, it is reasonable to assume that workers with bad health are less productive and therefore earn less than workers with good health status.[17]

If at least one of these arguments applies empirically, not controlling for unobserved heterogeneity will create a spurious correlation between c and μ and work absence, as workers with higher wage rates, on average, have higher costs of being absent from work. Since the main objective of this study is to study the policy implications of changing the compensation level in the SI it is of great importance to control for effects of unobserved heterogeneity using a fixed, rather than random, effects approach.[18]

The models are estimated using maximum likelihood (ML). The log-likelihood are equal to

$$\begin{aligned} \ln L &= \sum_{i=1}^{N}\sum_{t=2}^{T} d_{it-1}\left\{d_{it}ln(\pi_{it}^{11}(l)) + (1-d_{it})ln(\pi_{it}^{10}(l))\right\} \\ &\quad +(1-d_{it-1})\left\{d_{it}ln(\pi_{it}^{01}(l)) + (1-d_{it})ln(\pi_{it}^{00}(l))\right\}, \end{aligned} \tag{4.1}$$

where $\pi_{it}^{11}(l) = 1 - \pi_{it}^{10}(l)$ and $\pi_{it}^{00}(l) = 1 - \pi_{it}^{01}(l)$. Since we have state dependent intercepts, a_w and α_{wa}, the parameters are estimated, under the restriction $\sum_{i=1}^{n} a_i = 0$.[19]

5. Results

This section reports the results. The first subsection presents and discusses the parameter estimates. The second subsection reports the result when the model is used for simulating a policy changes.

[17] The relation between wages and absenteeism has been studied in several theoretical as well as empirical studies (see Brown and Sessions, 1996, for an overview).

[18] Note that the fixed effects model estimator is consistent (if $T \to \infty$) in a model with state and duration dependence and unknown initial state (see e.g. Heckman, 1981).

[19] Standard errors are calculated from the negative of the inverse Hessian.

5.1. Parameter Estimates and Sample Heterogeneity

Tables 5.1 and 5.2 shows the estimates for the male and female subsamples, respectively. In order to save space, the parameter estimates for the indicator variables for weekday and month of the year are excluded from the tables. However, the parameter estimates for these variables are very much in line with what can be expected from the descriptive statistics presented in Section 3.3.

If duration dependence is present for either work or work absence spells, the sample needs to be restricted to non left censored observations. This means that the first step in the estimation of the model is to test for duration dependence in each subsample. To do this, we first estimate the model without left censored work and work absence spells. For work absence spells, duration dependence is present in all subgroups and, consequently, the left censored work absence spells were left out. For the work spells, the situation is less obvious. The hypothesis of no duration dependence for the work spells ($\lambda = 0$), is tested with an ordinary Wald test. A small negative duration dependence in the work state is found in three sub-groups: males aged 20-29 ($t = -2.36$) and females aged 20-29 ($t = -3.94$) and 30-39 ($t = -3.49$). For the other subgroups the model is reestimated, including also the left-censored work spells.[20]

The parameter estimates of the duration dependence in the work absence state are very similar in the different subgroups. In general, the parameter estimates of the seven indicator variables are significantly negative with the largest parameter in absolute value for the seventh day, while the parameter estimates of the linear component are positive, although not significantly so for all subgroups. The interpretation of this pattern is that the workers are more likely to return to work given that they have been in the work absence spell less than seven days. After that, they are less likely to return to work, with an increasing duration in work absence as a consequence.

Section 4 stressed the importance of considering preference heterogeneity when modeling work absence behavior. When studying the estimation results, two questions is of interest. First of all, can preference heterogeneity be observed? If so, does it affect the parameter estimates of the other included variables, i.e. is the unobserved heterogeneity correlated with the independent variables?

Tables A1 and A2 in the Appendix contains estimates corresponding to those of Tables 5.1 and 5.2 when fixed effects are not included in the model. Performing likelihood ratio tests the restricted models are forcefully rejected for all subgroups. Figure 5.1 shows kernel estimates of the distribution of the fixed effects for the eight different sub-samples. The estimated heterogeneity is relatively large, with a heavy left tail, for the youngest age groups for both females and males as well as for the age group 30-39 for the females. The distribution of the fixed effects are close to be symmetric in the other five sub groups. A possible explanation to

[20]For males 30-39, 40-49 and 50-65 we get $t = -.74, .89$ and $-.82$, respectively. For females 40-49 and 50-65, $t = -.62$ and $-.61$, respectively.

Table 5.1: Parameter estimates for the male sub-samples ($N = 634$).

Age	20-29		30-39		40-49		50-65	
	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$
				Work				
α_w	-4.67	-9.51	-4.93	-16.07	-4.12	-11.33	-4.10	-9.32
c/γ_w ($\times 10^2$)	-0.02	-2.34	-0.05	-5.34	-0.04	-4.17	-0.04	-3.95
μ/ψ_w	0.05	0.94	0.02	0.64	-0.09	-1.80	-0.09	-1.56
$UNEM$	0.03	0.37	0.32	4.40	0.11	1.48	0.23	2.77
HO	-1.82	-5.64	-1.58	-4.90	-1.79	-4.65	-1.16	-3.73
BH	-1.67	-2.85	-1.38	-2.36	-0.30	-0.80	-1.02	-1.73
λ ($\times 10^2$)	-0.91	-2.37						
				Work absence				
α_{wa}	0.70	1.37	1.51	3.95	2.25	5.40	1.54	3.05
c/γ_{wa}	-0.04	-3.88	-0.08	-6.94	-0.08	-5.55	-0.06	-5.14
μ/ψ_{wa} ($\times 10^2$)	0.11	1.98	0.00	0.10	-0.07	-1.38	0.08	1.36
$UNEM$	0.19	3.03	0.30	4.43	0.23	3.18	0.39	5.20
HO	-0.25	-1.11	-0.39	-1.55	-0.53	-1.99	-0.25	-0.93
BH	0.00	0.00	0.00	0.00	1.33	1.24	0.12	0.19
λ_{wa}^1	0.12	0.74	-0.19	-1.07	-0.20	-1.04	-0.32	-1.66
λ_{wa}^2	-0.79	-5.02	-0.87	-4.99	-0.79	-4.17	-0.22	-1.11
λ_{wa}^3	-0.52	-3.20	-0.53	-2.93	-0.79	-4.11	-0.93	-5.12
λ_{wa}^4	-0.66	-3.81	-0.90	-4.92	-0.72	-3.48	-0.49	-2.37
λ_{wa}^5	-0.75	-4.25	-0.74	-3.87	-1.22	-6.20	-1.31	-7.12
λ_{wa}^6	-0.56	-2.63	-0.55	-2.25	-0.99	-4.17	-0.83	-3.50
λ_{wa}^7	-1.77	-9.46	-1.93	-9.93	-2.36	-11.16	-2.26	-11.78
λ_{wa} ($\times 10^2$)	1.53	4.00	1.00	3.09	0.79	2.74	1.13	4.88

Note: c is the cost of being absent; μ is the virtual income; HO is the parameter for the public holiday; BH is the parameter for days between public holidays and Sundays or Saturdays; $UNEM$ is the parameter for the local unemployment rate; λ is the parameter for linear duration; λ_{wa}^j j=1...7 are the parameters for the duration in work absence for day 1 - 7; λ_{wa} is the linear duration dependence in work absence for days 8 an onwards; also a weekday factor and a monthly factor are included in the estimation, but excluded from the table.

Table 5.2: Parameter estimates for the female sub-samples ($N = 590$).

Age	20-29		30-39		40-49		50-65	
	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$	$\widehat{\beta}$	$\widehat{\beta}/s_{\widehat{\beta}}$
				Work				
α_w	-4.28	-10.43	-4.55	-2.62	-5.21	-18.41	-4.44	-16.85
c/γ_w	-0.02	-2.18	-0.07	-5.95	-0.03	-3.96	-0.02	-2.87
μ/ψ_w ($\times 10^2$)	0.09	1.67	-0.05	-1.22	0.08	2.43	-0.01	-0.65
$UNEM$	-0.00	-0.09	0.37	4.19	0.25	3.60	0.13	1.67
HO	-2.23	-4.93	-1.15	-3.72	-2.36	-5.23	-1.55	-4.03
BH	-0.78	-1.70	-0.60	-1.30	-0.17	-0.55	-1.01	-1.71
λ ($\times 10^2$)	-1.42	-3.94	-1.39	-3.49				
				Work absence				
α_{wa}	0.97	2.13	1.26	0.72	1.27	3.89	2.83	8.04
c/γ_{wa}	-0.04	-3.54	-0.06	-4.60	-0.05	-4.70	-0.06	-5.56
μ/ψ_{wa} ($\times 10^2$)	0.19	3.27	0.12	2.44	0.08	2.40	-0.03	-1.18
$UNEM$	0.28	3.77	0.25	3.26	0.32	5.15	0.36	4.80
HO	-0.20	-0.90	-0.56	-2.33	-0.50	-2.36	-0.07	-0.25
BH	-0.39	-0.72	0.25	0.40	-0.08	-0.17	-0.77	-1.82
λ_{wa}^1	-1.02	-5.97	-0.86	-5.03	-1.06	-7.06	-1.23	-6.93
λ_{wa}^2	-1.41	-8.18	-1.33	-7.91	-1.12	-7.23	-1.38	-7.79
λ_{wa}^3	-1.35	-7.59	-1.20	-6.77	-0.97	-6.03	-1.39	-7.47
λ_{wa}^4	-1.06	-5.38	-1.12	-5.80	-0.81	-4.46	-1.03	-4.83
λ_{wa}^5	-1.54	-7.84	-1.52	-8.11	-1.31	-7.84	-1.17	-5.48
λ_{wa}^6	-0.72	-2.75	-1.06	-4.39	-1.02	-4.78	-1.31	-5.42
λ_{wa}^7	-1.98	-9.15	-1.93	-9.30	-1.72	-9.22	-1.97	-9.21
$\lambda_{wa}(\times 10^2)$	0.85	3.10	0.49	2.38	0.91	4.07	0.65	3.36

Note: c is the cost of being absent; μ is the virtual income; HO is the parameter for the public holiday; BH is the parameter for days between public holidays and Sundays or Saturdays; $UNEM$ is the parameter for the local unemployment rate; λ is the parameter for linear duration; λ_{wa}^j j=1...7 are the parameters for the duration in work absence for day 1 - 7; λ_{wa} is the linear duration dependence in work absence for days 8 an onwards; also a weekday factor and a monthly factor are included in the estimation, but excluded from the table.

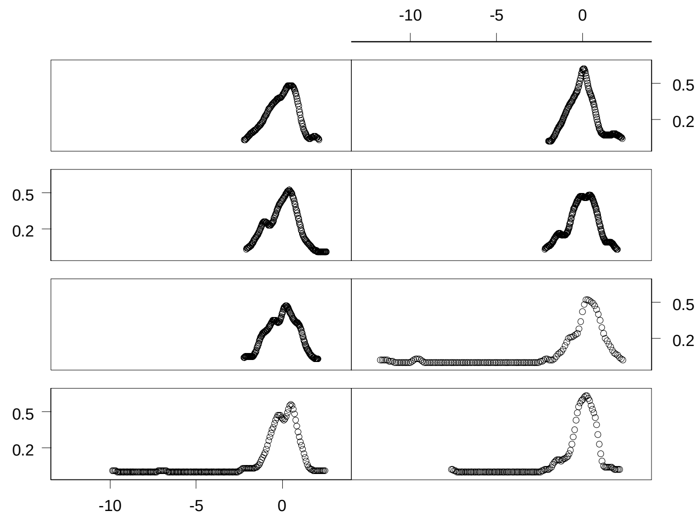

Figure 5.1: Kernel density (Gaussian with bandwith 0.5) estimates of the distribution of the fixed effects in the male (first column) and female (second column) sub-samples by different age-groups. First row: age group 50-65; second row: age group 40-49; third row: age group 30-39; fourth row: aged 20-29.

the lower heterogeneity found in the older age groups is that workers with bad health (and/or high preferences for being absent) may have left the labor force before reaching the oldest age group.

The parameter estimate for the cost variable is, as expected, significantly negative in all subgroups. This is true, conditional on both states. These results are consistent with theory, but contradictory to what was seen using KM estimator in Section 3, where, we found no effect of the SI reform on the duration in work absence. An explanation for this seemingly contradictory result can be found in the change of the composition of work absence spells after the reform. The individuals with relatively low (high) incidence to be absent from work are the same as those with relatively short (long) work absence spells. This is an implicit feature of the significant individual effects. After the reform, the incidence of work absence spells decreased. This decrease implied that individuals with low incidence of being absent, i.e. with few absence spells decreased their number of spells after the reform. Furthermore, the individuals with relatively high incidence and long work absence spells shortened their time in work absence. Consequentially the hazard rate in work absence seemed to be unaffected. However, the econometric model is able to distinguish these two effects.

The parameter estimates of ψ_w (virtual income in the work state) are generally insignificantly different from zero (one exception is for the females aged 40-49). In the work absence state, the estimates of ψ_{wa} are, as expected, positive and significantly different from zero in four of the subsamples (the three youngest female age-groups and the youngest male group) and insignificantly different from zero in the other subsamples. The frequent insignificant parameter estimates reflect the fact that these parameter are estimated with relatively low precision. An explanation may, as is discussed in Section 3, be the difficulties in measuring income.

Are the parameter estimates of the cost and income variables affected by unobserved heterogeneity? If we do not correct for unobserved heterogeneity then we would expect the parameter estimates for the cost variable ($\widehat{\gamma}_w$ and $\widehat{\gamma}_{wa}$) to be larger in absolute value. Comparing the magnitude of the parameter estimates with and without consideration of unobserved heterogeneity in discrete choice regression models may not be useful however, since the larger the unexplained variance in a structural model, ceteris paribus, will decrease the estimated effect.[21] For ψ_w and ψ_{wa} it is likely to be two effects working in different directions. One affect arising from dependence with the non-labor income, R, and the other from dependence with the wage rate, w. Following the discussion in Section 4, we expect the correlations to be positive and negative, respectively. Hence, not controlling for heterogeneity might change the sign of ψ_w and ψ_{wa}, due to the negative correlation between the heterogeneity and wage. Comparing the parameter estimates shown in Tables 5.1 and 5.2 with those in Tables A1 and A2 some differences between the two sets of parameter estimates ($\widehat{\psi}_w$ and $\widehat{\psi}_{wa}$) can be found. As noted above, the parameter estimates for the virtual income variable from the fixed effect model are often positive and never significantly negative. In the model without fixed effect the estimates are most often negative as well as significant on several occasions. In one sub sample, men aged 20-29, $\psi_{wa} = 0.05$ and $\psi_w = 0.11$ ($t = 1.98$) whereas when we do not control for heterogeneity, $\psi_{wa} = -0.03$ and $\psi_w = -0.08$ ($t = -3.55$), i.e. the parameter estimates changed signs significantly between the two specifications.

Another way of studying the potential effect of unobserved heterogeneity is through the sample correlation between the included variables and the estimated fixed effects. Table 5.3 shows that the point estimates of the correlation between the wage rate in 1989[22] and the estimated fixed effects are, as expected, negative, although not significantly different from zero, in all sub samples except females aged 50-65.[23] The highest point estimate of the correlation, in absolute value,

[21] Assume the following state specific utility specification:

$$u_{it}(l) = a_i + \mathbf{x}_{it}\boldsymbol{\beta}_k + \lambda_k^l + \varepsilon_{it},\ k = wa, w,$$

where $Var(\varepsilon_{it}) = \sigma^2$. We can only estimate $\boldsymbol{\beta}_k$ up to scale i.e., $\boldsymbol{\beta}_k/\sigma$. Removing the fix effect, a_i, will increase σ^2 and hence reduce $\boldsymbol{\beta}_k/\sigma$.

[22] This year's wage rates are used since the 1990 and 1991 wage rates are used in the estimation.

[23] The test statistic is distributed as Student's t if the wage rate and the fix effects are bivariate normal. We use the (natural) logarithm of the wage rate and the individual fixed effects in the

Table 5.3: Estmated correaltions between the (natural) logarithm of the wage rate (in 1989) and the estimated fix effects.

Age	Males		Females	
	$\widehat{\rho}$	t	$\widehat{\rho}$	t
$20-29$	$-.10$	-1.30	$-.04$	$-.48$
$30-39$	$-.09$	-1.13	$-.01$	$-.17$
$40-49$	$-.07$	$-.94$	$-.00$	$-.04$
$50-65$	$-.06$	$-.67$	$.05$	$.64$

Note: $t = \widehat{\rho}\left((1-\widehat{\rho}^2)/(n-2)\right)^{-1/2}$

is found in the subsample of men aged 20-29. In this group, as noted above, it was also found that controlling for unobserved heterogeneity actually changed the estimated signs of the parameters for virtual income $(\psi_w$ and $\psi_{wa})$.

The parameter estimates of the local unemployment rate variable are in general significantly positive in all sub-groups and in the work as well as in the work absence state. The effect seems to be somewhat stronger in the work absence state. These results are contrary to what is expected from the efficiency wage hypothesis (see e.g. Shapiro and Stiglitz, 1984) that predicts a negative correlation between absenteeism and unemployment rate.[24] It is also different from what has been obtained in most empirical studies (see Drago & Wooden, 1992; Johansson & Palme, 1996; Lantto & Lindblom, 1987), although there are some exceptions (see Bäckman, 1998).

If these results are compared to those obtained from the model without fixed effects (see Tables A1 and A2 in the Appendix), it can be seen that the result changes radically: the parameter estimates are in general insignificantly different from zero. In some subgroups, particularly in the work state for the female subgroups, the parameter estimates actually change sign to be significantly negative. As most empirical studies do not control for preference heterogeneity, these results can give some guidance to why we get different results compared to previous empirical studies.

There are reasons as to why we may observe a spurious relation between work absence and unemployment. It is a well known fact that work absence is lower in small communities in general and in firms with few employees in particular, due to a more extensive social control (SAF, 1986). The unemployment rate in Sweden, and in most other countries, is in general higher in small communities. These two unrelated properties may have made up the previously observed negative correlation between work absence and unemployment.

There are at least two explanations to why we may observe a positive pa-

calculation. The results are rubust to different transformations of the included variables.

[24] A high unemployment rate will work as a worker discipline device. If the worker is caught shirking and dismissed during times of high unemployment rate, the worker's cost of finding a new job, and thereby of getting dismissed, will be higher than it otherwise would have been.

rameter estimate for the local unemployment rate variable. First, an increased unemployment rate may create economic stress that may be damaging for the health (see Vahtera, Kivimäki and Pentti, 1997). This, in turn, may lead to a higher rate of work absence. Second, when aggregate demand decreases at the beginning of a recession, firms will try to layoff their workers rather than permanently dismiss them if they expect the demand to increase again later on. It is, of course, cheaper for the firm if the laid off workers claim sickness benefits compared to getting paid from the firm. There is anecdotal evidence that employers may in fact urge their employees to do so. If this is the case, we may also observe a positive correlation between unemployment and work absence in the beginning of a recession, when the unemployment rate is increasing.

Finally, the parameter estimates for the public holiday indicator variable, HO, as well as the indicator for day between public holiday and weekend, BH, give no support to the anecdotal evidence on an increased use of the SI on days between public holidays and weekends. The parameter estimates for BH is always negative when a individual is in the work state, however only significantly so for the males in age group 20-29 and 30-39.

5.2. Policy Simulation

In order to evaluate the estimated model, we will use the model to simulate the effects on work absence of the reform of SI in March 1991. The probability to stay in work and work absence for an individual who have been in either of these states for l period is

$$\pi_{it}^{00}(l) = \frac{1}{1+\exp(a_i + \mathbf{x}_{it}\boldsymbol{\beta}_w + \lambda_w^l)} \quad \text{and} \quad \pi_{it}^{11}(l) = \frac{1}{1+\exp(-a_i - \mathbf{x}_{it}\boldsymbol{\beta}_{wa} - \lambda_{wa}^l)},$$

respectively. The survival function for a work and work absence spell up to time period τ is then

$$S_{i\tau}^{00} = \prod_{l=1}^{\tau} \pi_{it}^{00}(l) \quad \text{and} \quad S_{i\tau}^{11} = \prod_{l=1}^{\tau} \pi_{it}^{11}(l).$$

Since the survival functions depends on $\mathbf{x}_{it}$, we assume that the spell starts on January 1, 1991 and use the wage and non-labor income in 1991 for the individuals. The survival functions are calculated with cost, $c_i = (1-\delta_l)w_i$, and virtual income, $\mu_i = R_i + \delta_l h_i^* w_i$, where $\delta_l = 0.35, 0.20$ and 0.10 depending on the spell length in work absence after the reform and $\delta_l = 0.1$ for all l before the reform. The survival function for a work or work absence spell up to time period τ is calculated at the average of the individuals survival functions, hence $S_\tau^{00} = \frac{1}{n}\sum_{i=1}^{n} S_{i\tau}^{00}$ and $S_\tau^{11} = \frac{1}{n}\sum_{i=1}^{n} S_{i\tau}^{11}, \tau = 1, ..., T$.

Figures 5.2 and 5.3 show the results of the simulations on the work and work absence spells, respectively. The most noticeable with Figure 5.2 is the resemblance with the KM estimates shown in Figure 3.1. The effect of the reform change

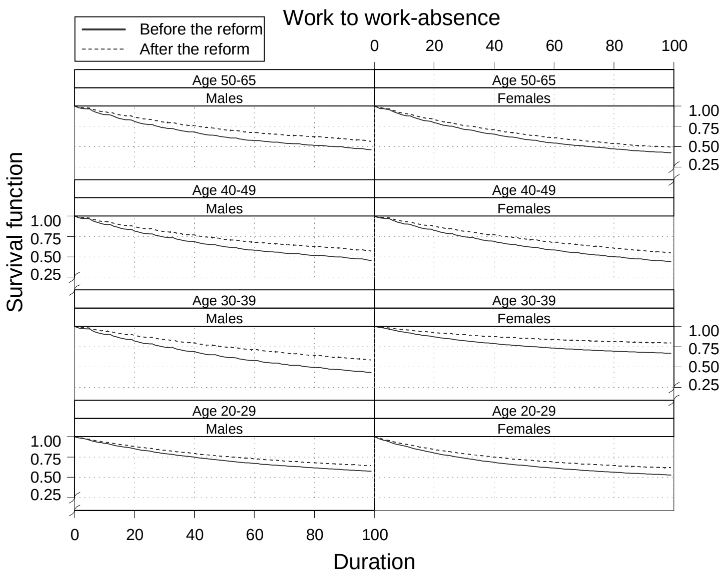

Figure 5.2: The mean survival function in work for the individuals in the eight sub-groups. (The spell starts in January 1, 1991)

is very similar to the KM estimates in all age groups except for the age-group 30-39, (both genders) where the change in the reform gives a larger effect than in Figure 3.1. The reason for this may be the relatively large positive parameter estimate found for the unemployment rate for these two groups.

Turning to the predictions of the work absence spells shown in Figure 5.3 a quite large difference as compared with the KM estimates in Figure 3.2 can be seen. First, the overall predictions for the work absence spells are inferior compared to the predictions of the work spells. The effect of the SI reform is larger than that seen in Figure 3.2. This difference can be attributed to the negative parameter estimate for the cost parameter discussed above.

Comparing the results in the different subsamples, it can be seen from Figure 5.3 that the model predicts the general pattern of an decreasing hazard with age very well. The simulation shows that the effect of the reform on work absence spells is largest in the oldest age group. At least a part of the explanation for this difference is due to economic incentives through lower cost of being absent from work (due to lower wage rates) and lower non-labor income pertaining to the younger age groups.

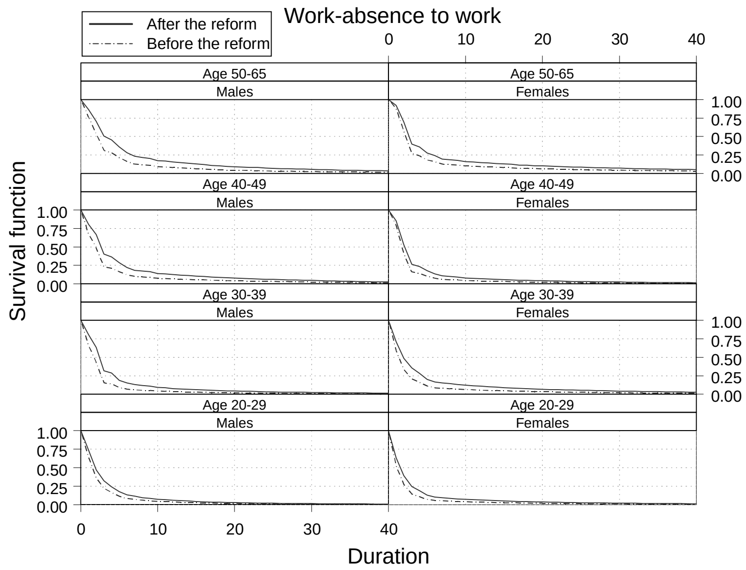

Figure 5.3: The mean survival function in work absence for the individuals in the eight sub-groups. (The spell starts in January 1, 1991)

6. Conclusions

To conclude, let us discuss two issues which, as we see it, are of central importance in this study. First, how do economic incentives affect work absence? In general, the results obtained from the estimates of the model support the view that the cost of being absent affect work absence behavior. The model simulation of the effect of the reform of the sickness insurance shows that the increased cost, rather than the higher unemployment rate, caused the decrease in the work absence rate. Contrary to what appeared from the non-parametric Kaplan-Meier estimates of the reform, the estimates from the econometric model shows that the cost of being absent affects both the incidence of work absence spells as well as the expected length of the spells. Significant impact of virtual income on incidence of work absence spells and the expected length of the spells were fond in some subgroups of the sample.

Second, is preference heterogeneity empirically important? We have chosen to model work absence behavior using state dependent parameters and different duration dependence depending on state. If unobserved heterogeneity is empirically unimportant, or uncorrelated with the independent variables, conventional hazard regression models could have been used as well. To the extent that unobserved heterogeneity are present it could have been dealt with using a random effects specifications. However, the estimates obtained here shows that unobserved heterogeneity is present, and more important, that it affects the inference of the parameters.

The main limitation of this study is that we do not allow the individual to plan their work absence behavior. The individual has at least one motive to do so. By being absent from work when affected by a temporary illness the worker may maintain a better long term health status. From the results obtained in this study it is hard to have any a priori views on how the investment in health motive affect work absence behavior. In the 1991 reform of the SI, the cost of making such investments increased. This may, of course, discourage the worker from doing such investments. On the other hand, the payoff from doing such investments increased as well, as the cost of long sickness spells also increased after the reform.

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