SHORT-TERM ABSENTEEISM DUE TO SICKNESS: THE SWEDISH EXPERIENCE, 1986 - 1991*

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Abstract

The goal of this paper is to analyze short term-absences from work (i.e., periods of seven days or less) in Sweden during a period with two different reforms. As a theoretical model we use a utility-maximization framework with two restrictions (time and budget constraints). Using multiple spell data, short-term absenteeism is analyzed for a period with three regimes, and it is found that the 1991 reform (which lowered the replacement rate) had a stronger effect on the hazard of ending short-term absenteeism than did the 1987 reform (which eliminated the previous unpaid "waiting day", while restricting the remuneration to only those days when people were scheduled to work). Even though economic incentives mattered, people with *poorer* health did not "shorten" their absences in the same extent as those with *better* health.

Key words: short-term absenteeism sickness spells, repeated events, unobserved heterogeneity.

JEL classification: I18; J22; J32; J33.

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1 Introduction

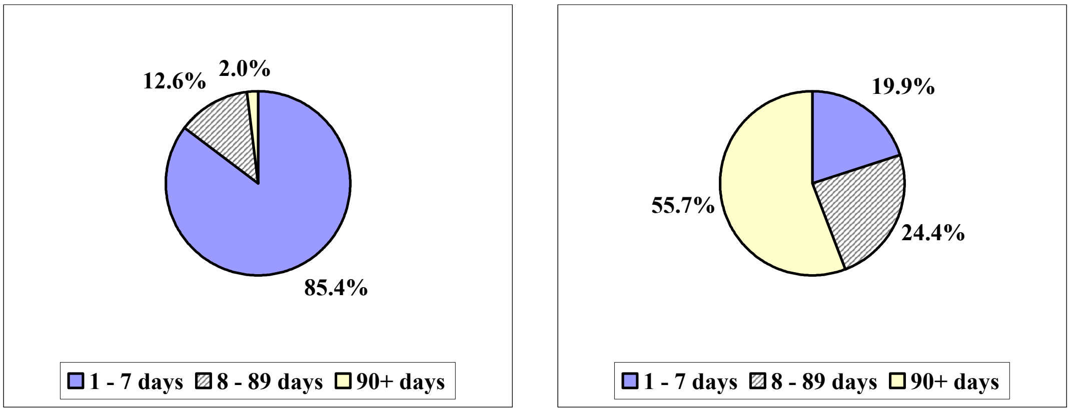
Employee absenteeism has long been an important subject of psychological, sociological, and economic research. Labor absence can be thought of as any time spent away from the workplace that is not anticipated or scheduled by the employer. The causes of work absenteeism are debated from the firm level to the macro level. People may be absent from their job because of either their own or another family member's sickness, because of death in the family, or for other strictly personal reasons. But there are also working-environment factors that determine absence from work, such as job involvement and satisfaction, a culture with strict attendance norms, etc. Persons with a high level of job satisfaction, or whose work-culture includes strict attendance norms, may seldom be away due to poor health, whereas low satisfaction, and/or lax norms, may lead to greater absenteeism. The purpose of this study is to analyze short-term absenteeism (i.e. spells of 1-7 days) due to sickness.[1] A medical certificate is required from the eighth day of sickness, so 1-7 days is a natural (short-term sickness) category. We will study the duration of these spells with regard to individual and labor market characteristics, but also with regard to characteristics related to the spells: diagnosis, the season when they occurred, and the weekday when they started.

Figures 1 a and b, and Table 1 (which motivated the interest for this study) show that about 85% of all sickness spells which ended in 1991 had a duration of 7 days or less, and accounted for about 20% of all days with a sickness benefit.[2] On the other hand, sickness spells of 90 days or more accounted for only 2% of cases, but for over 55% of compensated says. The percentage of 1-7 day cases had been substantially lower in 1986 and 1987 (Table 1). The jump in 1988 (and thereafter) appears to relate to the

[1] Short-term sickness is an absence of 1-7 day due to sickness. It could be considered "voluntary" because a medical certificate is not required until the eighth day.

[2] In Sweden, "sickpay" is sometimes provided by the employer, and sometimes by the social insurance system. When the distinction matters, the term "sickness cash benefit" will be used for the later, and the term "sickpay" reserved for the former. During the study period, excepting one waiting day before December 1987, social insurance covered all days of employees' sickness.

fact that, from December 1987, the previous unpaid "waiting day" was eliminated, although the compensation was provided for only those days when people were scheduled to work.



a) Number of sickness spells ending in 1991 distributed by duration

b) Total number of days of sickness by spell duration

Figure 1 Duration-shares for sickness spells and for total number of benefit days,[3] by spell-duration, in 1991

Table 1 Sickness spells and the total number of sickness cash benefit days, by spell-duration, 1986-91 (%)

Year ended	Cases of sickness concluded			Days of sickness cash benefit		
	1-7* days	8-89 days	90+ days	1-7* days	8-89 days	90+ days
1986*	76.9	21.1	2.0	16.8	38.1	45.1
1987*	77.4	20.6	2.0	16.8	37.0	46.2
1988	84.4	14.0	1.6	23.8	30.3	45.9
1989	84.9	13.4	1.7	23.0	29.0	48.0
1990	85.6	12.7	1.7	22.8	27.4	49.8
1991	85.4	12.6	2.0	19.9	24.4	55.7

*Before December 1, 1987, the day when the sickness was reported was not covered by social insurance, so that the number of actual sickness cash benefit days was **1-6**, **7**-89, and 90+.

[3] The Swedish National Social Insurance Board (RFV) is the *source* of data for the entire paper, except were other sources are mentioned.

The next section reviews the literature, while Section 3 describes the institutional setting. Section 4 describes the theoretical framework used, and Section 5 the data (mostly via Appendix 2). The econometric specification, and results, are presented in Sections 6 and 7. Section 8 draws conclusions and discusses further research possibilities.

2 Literature review

Douglas (1919) was perhaps the first to analyze absenteeism, which was mentioned as "another factor in the instability of labor, which has not been given the attention that it deserves". Given the fact that during that time it was often confused with labor turnover, Douglas defined absenteeism as "absence from work at the job at which one is employed", including absence for both all-day or only partially. He examined the amount of absenteeism (i.e., the number of days and hours lost) at the plant and company level, pointing out its causes, the resulting losses, and possible methods of reducing it.

In both economic and psychological research, a common assumption is that individuals rationally allocate their labor supply by making daily decisions to attend the work or non-work setting with the highest expected utility [e.g., Winkler (1980), Chelius (1981), Youngblood (1984), Lantto (1991)]. Allen (1981a) presented a mathematical form of this model, and concluded that if absence is a normal good, then absence following a wage increase can be expected to rise as an income effect (also with respect to non-labor income), but to decline as a substitution effect. In this framework, an increase in monetary penalties for absence, such as fewer available days for future paid sick leave, would reduce absence. Additionally, Allen (1981b) argued that employees trade off wage levels against expected absences when choosing employment.

Employees may also have an incentive to be absent if the contractual hours specified by the employer exceed their optimal amount of labor supply (Brown and Sessions, 1996). Barmby et al. (1994) presented a static model in which contractual considerations played a prominent role. Building on this static model, Brown (1994) addressed two major weaknesses in the previous theory of absence behavior: the lack of demand-side considerations, and dynamics over time. Kaiser (1996) presented a model in which absence behavior was jointly determined by the employer and employees through interactive processes; first, between the absence culture of a work group and the preferences/behaviors of its individual members, and second, between the larger organization and the absence behavior of the work group. Brown (1999) used the conventional labor supply model of absence behavior, extended to multi-period analysis. Her analysis suggested again that absenteeism is primarily affected by contractual characteristics, such as the wage rate and penalties for absence. Other studies [e.g., Allen (1981a, 1981b), Barmby et al. (1991, 1995), and Johansson and Palme (1996)] have also found effects of economic incentives on individual absences.

A critical feature of the idea of *short-term choice* is that it explicitly considers the utility of both work and non-work alternatives. But sometimes the non-work alternative is more necessary, as, for example, when poor health reduces capacity to work in a given environment.

When health variables have been incorporated into absence models [e.g., Allen (1981b); Paringer (1983); and Leigh (1983, 1986, 1991)], they have usually been found to be the most important determinants of absenteeism. Other studies, however, have found that some health variables were not significant. For example, French et al. (1998) using various measures of current and lifetime drug-use, and accounting for alcohol-use co-morbidity, found no significant relationship between drug-use and either wages or absenteeism, regardless of gender.

The unemployment rate often appears to be negatively correlated with absenteeism. Some studies predicted that increased unemployment would lead to less absenteeism at the *individual* level [e.g., Larson and Fukami (1985), Leigh (1985), and Drago and Wooden (1992)], at the *plant* level [e.g., Markham and McKee, 1991], at the *industry* level [e.g., Leigh, 1985], and at the *national* level [e.g., Doherty (1979), Leigh (1985), Lantto and Lindblom (1987), and Bäckman (1998)]. Lidwall and Skogman Thoursie (2000), using official statistics produced by the Swedish National Social Insurance Board found that short-term sickness absence increases at lower levels of unemployment, and decreases at higher levels of unemployment.

Absenteeism has been found to be significantly reduced by profit sharing and employee share-ownership plans [e.g., Brown et al. (1999)].

Absenteeism has also been found to be different for women and men [e.g., Allen (1981b, 1984), Leigh (1981), Dunn and Youngblood (1986), Johansson and Palme (1996), and Vistnes (1997)]. Nevertheless, Vistnes (1997), investigating the extent and determinants of gender differences in days lost from work due to illness, found that, for both men and women, health status measures, such as self-reported health status and medical events, explained work absence more consistently than did economic incentives.

Married persons have been found to have generally lower absence rates [e.g., Keller (1983), Allen (1984), and Leigh (1986)], which might be explained by the family budget constraint, especially when only one member of the household is working and earnings replacement is well below 100%. They might also just feel better.

White-collar workers have been found to be absent less than blue-collar workers [Kenyon and Dawkins (1989)]. This may be because they are less likely to be injured at work, or work more often in occupations where it is considered acceptable to work with colds, but also easier to work with acute back pain, etc. Blue-collars workers also do more shift work, which has been associated with higher rates of absence [Drago and Wooden (1992)].

Some studies have used temporal patterns of absenteeism to make inferences about short-term absence processes, focusing on *when* an employee will be absent, rather than on *how long*. Fichman (1988, 1989) demonstrated that the daily probability of absence went up as the time since the last absence increased, but went down to the extent that fulfilling non-work events occurred (e.g., a paid holiday). Harrison and Hulin (1989) found that absence on a daily basis was uniquely associated with short-term attendance-history and with temporal variables (the weekday, and the season or month).

Some studies on absenteeism in Sweden have used day-to-day data. For example, using a sample of blue-collar workers (from the Swedish Level of Living Survey) with day-to-day data from 1991, Cassel et al. (1996) found strong economic incentive effects on absenteeism, but they also found that the sharp decrease in work absences after the 1991 reform, which lowered the replacement rate, could not be attributed solely to the higher cost of missing work. Using a linear demand function and the 1981 cross section from the Swedish Level of Living Survey, Johansson and Palme (1996) estimated absenteeism as an individual day-to-day decision. Their binomial maximum likelihood estimators were consistently estimated under the assumptions of unobserved heterogeneity and serial correlation. For the male subsample, they found a negative effect of lost earnings on work absence. Using a generalized method of moments estimator, Johansson and Brännäs (1998) empirically tested a household model for the day-to-day absenteeism decision, with similar results.

Some studies have used time-series for the average number of compensated days of sickness [e.g., Lantto and Lindblom (1987); Gustafsson and Klevmarken (1993); and Bäckman (1998)], and found a negative effect of the unemployment rate on the sickness absence rate. Except for the unemployment rate, the model specification differed across these studies. For example, Gustafsson and Klevmarken used the replacement rate as an explanatory variable; whereas Lantto and Lindblom, as well as Bäckman used dummy variables for changes in social insurance rules.

3 Sickness cash benefit in Sweden, rules and statistics

The study period for this paper is January 1986 through December 1991, during which all residents of Sweden, aged 16-64 years, and whose annual income was at least 6000 Swedish crowns (i.e., about 1100 US dollars in 1991) were eligible for a sickness cash benefit if they lost income due to sickness.[4] The National Insurance Act gives no general definition of sickness, but according to the National Social Insurance Board's recommendation, *sickness* is an abnormal physical or mental condition;[5] if it reduces normal work capacity by at least 25%, the afflicted individual can qualify for a sickness cash benefit. *Normal work capacity* is defined as either the ability to perform the same task, or the ability to earn the same income, as prior to sickness.

There were two social insurance reforms during the study period, effective December 1, 1987, and March 1, 1991. The changes that affected short-term

[4] This applied not just to employees, but also to the self-employed, who had a choice of applicable coverage, however.

[5] The sickness cash benefit is actually granted by the local social insurance offices. The National Social Insurance Board cannot set binding policy for them, but can only recommend its interpretation of law.

absenteeism due to sickness are summarized in Table 2.

Table 2 Social insurance rule changes affecting short-term absenteeism during 1986-91

Changes (in **bold**)	Regime 1	Regime 2	Regime 3
	Jan 1986 – Nov 1987	Dec 1987 - Feb 1991	March 1991 – Dec 1991
Coverage	• The first day of reported sickness NOT covered. • Holidays NOT covered	• **The first day of reported sickness covered** • **Only scheduled work days are covered.**	
Replacement rate	90%	90%	**65% first 3 days** **80% day 4 - day 90** **90% day 91-**

The compulsory sickness insurance when it was implemented in 1955 stipulated a *waiting period* of three days and a limit of two years replacement in long-term sickness. In 1967 the waiting period was reduced to the day of calling in sick. In 1985 some administrative changes (for state employees) implied that also the day for calling in sick and weekends were counted as sickness absence days. In the period covered by this study, before December 1987, there was one unpaid "waiting" day before a sickness cash benefit could be claimed. For sickness spells of 7 days or less (excluding the first day), the compensation was not provided for non-working days (at most, two days).[6] Starting with December 1987, the waiting day was abolished, and a sickness cash benefit was provided from the day the sickness was reported to the social insurance office. However, a cash benefit was now only provided for scheduled workdays during the first *fourteen* days of absence. Until March 1991, the sickness cash benefit replaced 90% of lost earnings.

For most countries, including Sweden, absenteeism follows a typical pattern over the business cycle: There are more and longer absences when unemployment is low [e.g. Allen (1981a), Kenyon and Dawkins (1989), Drago and Wooden (1992), Johansson and Palme (1996)]. Figure 2 shows this inverse cyclical fluctuation of

[6] For longer spells, compensation was paid for all days, except the registration day. The self-employed could choose between waiting period of either 3 or 30 days.

absenteeism and unemployment in Sweden during the last three decades. During the economic slowdown from 1976 through 1983, the average annual number of compensated days of sickness per insured person declined from about 23 days to about 18, while unemployment reached a peak in 1983 (at the end of second OPEC recession). During the expansion of 1983 to 1989, the unemployment rate decreased, while the average number of compensated days of absenteeism due to sickness increased through 1988; and the inverse cyclical pattern then generally continued. The Swedish recession during 1991-1994 resulted in a huge increase in unemployment, from 2% to almost 10%, while absenteeism fell, reaching a low in 1997.

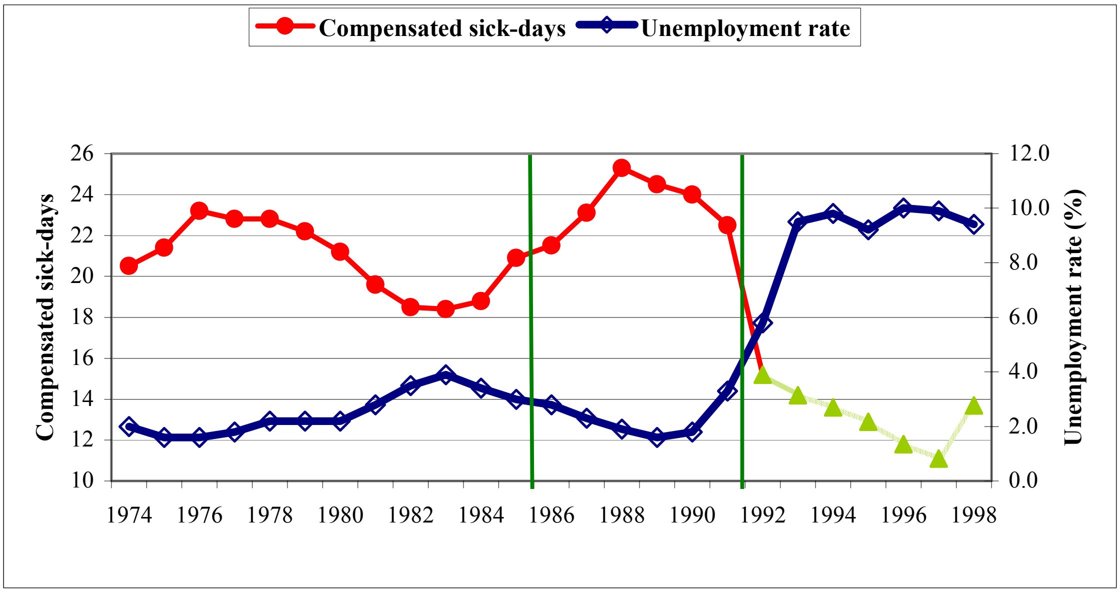


Figure 2 Average compensated sick-days per insured person[7] and the unemployment rate, 1974-1998

Economic incentives associated with the social insurance system also appear to have influenced absenteeism. After the unpaid waiting day was abolished in December 1987, there was a significant jump in the average number of compensated days of absence due to sickness, even though during the first *two* weeks, only scheduled

[7] Additionally to RFV's source, OECD Health Data 1998 is also used.

workdays were now covered. After the replacement rates were lowered (especially during the first three days) in early 1991, the absence rate fell drastically. Besides the high unemployment and lower replacement rate, the introduction of a two week "employer period" in January 1992 (represented by the "triangles" in Figure 2), contributed to a drop in average days of absence due to sickness.[8]

4 Theoretical framework

The conventional labor supply model of absenteeism focuses on contractual arrangements, assuming that the wage rate plays a central role. If markets were perfect, spot contracts would be used, and an employee who would benefit from absence on a given day would simply not go to work that day. In addition to the wage rate, however, employment contracts typically specify standard hours during which the employee is expected to work (on any given day); if these standard hours of work exceed the worker's preferred hours at the given wage, then there exists a potential utility gain from absence during the contracted hours.

In this study, short-term absenteeism is analyzed using a utility maximization framework based on Allen (1981a), Kenyon and Dawkins (1989), and Johansson and Palme (1996, 1998). It assumes that the distribution of information between employers and employees is asymmetric, in that employers must accept their employees' word regarding their actual state of health. Thus, there is an element of moral hazard in the decision to miss work.

The potential for absenteeism is determined by the employment contract. At a given wage, an employee who wants more leisure than provided therein can choose to be absent from work, possibly with sickness compensation. For the firm, this can obviously result in lost output, but it is assumed here that employees do not consider this impact directly; they rather base their decisions on their own well-being.

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[8] The "employer period" requires the employers to pay for the first weeks of sickness. Since January 1992 there has been a two-week employer period, except the time period January 1997 up to and including March 1998 when the employer covered the first four weeks.

The model uses the theory of choice in which purchased goods are one of the inputs into the production of commodities that directly enter preferences. Preferences are assumed to be a function of an ordered set of commodities Z_1, Z_2, and Z_3, the pair-wise indifference curves of which are assumed strictly convex. Z_1 is work, Z_2 is absenteeism, and Z_3 is leisure.

Let t_w measure the time spent by an employee i in work activities, t_a measure the time spent doing other activities than those specified in the work contract during the time when the employee is supposed to realize work activities specified in the contract, and t_l measure the time spent for leisure (non-market work activities, recreation, etc.). The characteristic of absenteeism due to sickness is that the employee can do any other activity than those specified in the work contract, being entitled to a compensation for the loss of earnings during the sickness period.

We can then write the model as (1)-(3)

(1) $(\max)U = U(Z_1, Z_2, Z_3)$

(2) $t_w + t_a + t_l \equiv t$

(3) $\sum_{j=1}^{3} p_j x_j = w(t_w + \rho t_a) + V$

where $Z_1 = f_1(t_w, x_1)$, $Z_2 = f_2(t_a, x_2)$, and $Z_3 = f_3(t_l, x_3)$ are three commodities produced by the employees, combining different market goods (x_j) and time (t_j); f_j are the household production functions, where f_1 and f_2 are assumed to have the same monotonic behavior. Restriction (2) represents *the time constraint*, in which t is the analyzed time period, expressed in hours, days, weeks, etc.; it can be normalized to one. Restriction (3) represents *the budget constraint*, in which the parameter ρ ($0 \le \rho \le 1$) is the replacement rate used in computing the sickness cash benefit, w ($w > 0$) is the wage rate, and V represents other income minus taxes.

The Lagrangian expression is

(4) $L = U(Z_1, Z_2, Z_3) - \lambda\left(\sum_{j=1}^{3} p_j x_j - w(t_w + \rho t_a) - V\right) - \mu(t_w + t_a + t_l - t),$

where $\lambda > 0$ represents the marginal utility of money income, which converges towards zero as income becomes high, and μ represents the marginal utility of time.

Using equilibrium conditions for the allocation of time, and substituting (see Appendix A1), we get

$$\textbf{(5)} \qquad \lambda(1-\rho)w = \frac{\partial U}{\partial t_a} - \frac{\partial U}{\partial t_w}.$$

It follows that, regardless of the value of the hourly wage (w) and marginal utility of money (λ), assumed to be strictly positive, individuals would be indifferent between work and be absent (i.e., $\frac{\partial U}{\partial t_a} = \frac{\partial U}{\partial t_w}$) only if $\rho = 1$ (i.e., 100% replacement rate).

If $\frac{\partial U}{\partial t_a} > \frac{\partial U}{\partial t_w} > 0$, individuals would choose to both work and be absent only if $0 < \rho < 1$, and $\lambda > 0$. Thus, given that the marginal utility of money is positive, and that the marginal utility of being absent is greater than the marginal utility of working, the level of voluntary absenteeism is determined in the model by the replacement rate. If it is low, people will be less likely to be absent, but if it is high, they will be more likely. In this model, then, economic incentives clearly affect labor participation.

If $\frac{\partial U}{\partial t_w} > \frac{\partial U}{\partial t_a} > 0$, individuals would choose to both work and be absent only if $0 < \rho < 1$ and $\lambda < 0$. In other words, if people enjoyed work very much, they would choose to be absent only if the marginal utility of money were negative. But a negative marginal utility of money is not allowed in the model, so it cannot happen that the marginal utility of work is higher than the marginal utility of being absent. Indeed, whether work is pleasant or not cannot be ascertained easily from observed behavior, but nevertheless, we expect that people who enjoy their work and working place would choose to be absent less frequently than people who are not very happy with or devoted to their work.

In sum, the model developed here addresses some aspects of *time allocation with respect to voluntary absenteeism*. Assuming that voluntary absenteeism is possible, the model illustrates that both the marginal utility of money and the replacement rate are important for the decision to miss work. In order to increase work discipline, it might be necessary to have a restriction on absenteeism. In this paper, however, the model is used *without* this restriction, in order to match the actual historical situation in Sweden during the study period.

5 The data

The LS database used here is a longitudinal database provided by the National Social Insurance Board of Sweden. The data encompass about 4500 individuals born on the 25th day of the month sometime during 1926-1966. The observation period was January 1, 1986 through December 31, 1991, which means all the individuals we will analyze are of working age.

There are two samples: 1) **IP**, a national sample based on stratified selection of the *entire insured population* of Sweden; and 2) **LSIP**, a national sample of the *long-term sick insured population*, selected from those who had at least one sickness spell of at least 60 days during the period January 1986 to December 1989. As there could be differences between the two populations with respect to the subject of this study, both samples were examined. Descriptive statistics of these samples and their analysis are presented in Appendix A2.

6 Econometric specification

The approach taken by the majority of researchers [e.g., Allen (1981a, 1981b), Dunn and Youngblood (1986), Chaudhury and Ng (1992)] has been to explain *the amount of observed absence* (i.e., absence rates across individuals or work places) with a set of regressors, such as wages and contractual hours, using a model derived from the income-leisure framework. Another approach analyzes the *probability of being absent on a particular day*, which is likely to be dependent on whether or not the individuals were absent the day before, given that the health status of an individual on a particular day is not independent of their health status the day before. This study uses this approach to analyze the hazard of ending the spell of short-term absence.

Let D_i measure the completed duration of absence due to sickness for employee i, and $S(t)$ measure the probability that an employee would be absent from work for at least t days, where $S(t) = \Pr(D > t)$ and $t > 0$. The corresponding distribution function of

D is then $F(t) = \Pr(D < t) = 1 - S(t)$, where $t > 0$. From information on D, we want to estimate the impact of observable (p) and unobservable (u) personal characteristics on the duration of absence.

One can estimate nonparametric, semiparametric, and parametric regression models with censored survival data. The semiparametric approach has the advantage that it does not make any assumption about the underlying distribution of completed sickness spells. Assuming that the hazard function can be factored into a function of time and a function of variables related both to spell and to individual, we can model the hazard of ending short-tem absenteeism (or the hazard of returning to work) as

(6) $$h_i(t; x_i) = h_0(t)\exp(\beta x_i),$$

where (β_1, β_2,…, β_k) is a vector of unknown parameters; x_i is the vector of k covariates for employee i, which may depend on time, or not; and $h_0(t)$ is the baseline hazard function, an unknown function of time. The expression $h_0(t)$ gives the hazard function for the standard set of conditions $x = 0$, and leaves $h_0(t)$ parametrically unspecified.

The data used here have a multi-episode design, which means that we have to check if there is a significant difference between absences across observation units, i.e., whether or not the sample is heterogeneous (neglected heterogeneity between observation units can lead to incorrect conclusions). For example, there are many techniques for analyzing duration data that are based on the assumption that the durations of distinct individuals are independent of each other, but in the case of repeatable events, this assumption is questionable, especially when same individuals have many spells.

There are basically two approaches to analyzing repeated events: 1) a separate analysis for each successive event; or 2) an analysis of all spells together, treating each spell as a distinct observation. The first approach gives a biased sample of later spells (for example, only people who have already had two spells, in the analyzed period, can have a third spell), and it could be inefficient, especially if the underlying process is unchanged from one period to the next, which would result in several redundant estimates. The second approach has the potential problem of dependence among multiple observations, which can be thought of as arising from *unobserved* *heterogeneity*, leading to declined hazard functions, and coefficients that are attenuated toward zero.

There are methods [e.g., Chamberlain (1985), Wei et al. (1989), and Allison (1996)] that correct some of these problems.[9] A fixed-effects version of a Cox regression (partial likelihood) is available for data in which (at least two) repeated events are observed for each individual [Chamberlain (1985), Yamaguchi (1986), Allison (1996)].

In the applied econometric literature on the estimation of multiple-duration models, the range of different models is actually not very large. Van den Berg (2000) provides an overview of duration analysis, with an emphasis on models for multiple durations, especially on the mixed proportional hazard (MPH) model and its multivariate extensions. For the multivariate mixed proportional hazard (MMPH) model, in which the marginal duration distributions each satisfy an MPH specification, and the durations can only be dependent by way of their unobserved determinants, he discusses the dimensionality of the heterogeneity distribution, and compares the flexibility of different parametric heterogeneity distributions.

Frequently in the analysis of survival data (e.g., how long sick employees "survive" before returning to work), survival times within the same "group" are correlated due to unobserved covariates. One way these covariates can be included in the model is as *frailties*; a frailty term represents the common covariates that are not

[9] Chamberlain (1985) introduced an approach, called Fixed-Effects Partial Likelihood, which corrects for some or all of the bias in the coefficients caused by unobserved heterogeneity; however, he expressed reservation about its use when the number of intervals varies across individuals, and when spell-duration depends on the lengths of the preceding spells. Wei et al. (1989) proposed a method for getting robust estimates that allows for dependence among multiple spells, allowing the computation of efficient pooled estimates of the coefficients and their standard errors, but it does not correct for biases in the coefficients due to unobserved heterogeneity. Using Monte Carlo simulations, Allison (1996) concluded that, except in cases where the number of previous spells is included as a covariate, there is, in practice, little or no problem regarding Chamberlain's concerns.

observed or are neglected. A *frailty model* [10] is a random effects model for time variables, where the random effect (the frailty) has a multiplicative effect on the hazard. This model can be used to describe the influence of unobserved covariates in a proportional hazard model, for example with multivariate failure times generated independently given the frailty for "groups" (both for survival times for related individuals, like twins or family members, and for repeated events for the same individual). These frailty random block effects generate dependency between the survival times of the individuals that are conditionally independent given the frailty.

Here we will assume that all individual variation in the hazard function can be characterized by a finite-dimensional vector of observed explanatory variables x and an unobserved heterogeneity term u. We can thus redefine our model (6) as

(7) $h_{ij}(t; x_{ij}, u_i) = h_0(t) u_i \exp(\beta x_{ij})$, $i = 1, 2, \ldots, G,\ j = 1, 2, \ldots, n_i$.

where h_{ij} represents the hazard rate of subject j in group i; u_i (u_i=exp(σw_i) can be interpreted as a function of unobserved explanatory variables. According to Lancanster (1990), u_i may also to some extent represent measurement errors in D and x. The u_i's are assumed independent and identically distributed from a distribution with mean 1 and some unknown variance. When u_i >1, employees in a given group tend to "fail" (in this case, return to work) faster than under an independence model (where u_i = 1). When u_i <1, employees in a given group tend to "fail" slower than under an independence model. The unobserved heterogeneity term is assumed constant from one event (absence) to the next, and has a specified distribution, independent of x_{ij}.

In this study, we expect that rule changes created substantial variation in the cost of being absent from work during the period studied, and changed the pattern of short-term absences due to sickness during the study period. The longitudinal structure of the

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[10] Clayton (1978) and Oakes (1982) were the first to consider frailty models for multivariate survival data, using gamma distribution for the frailty. Hougaard (1986) introduced the G-family of distributions, which includes the gamma distributions and inverse Gaussian distributions. He also used the positive stable distribution for the frailty, along with arbitrary and Weibull hazards. Lu and Bhattacharyya (1990) used the Weibull distribution to model the frailty parameter, while Whitmore and Lee (1991) studied a model with inverse gamma frailties.

data provides *multiple spells*, which makes it possible both to analyze these changes, but also requires that we deal with *unobserved heterogeneity*. Therefore, we used a frailty model in the empirical analysis.

7 Empirical Results

In the *fist* step, using nonparametric estimation, a preliminary analysis of the short-term absences due to sickness was produced for both samples. Tables 3 and 4 show the plots of the estimated survival and hazard functions, respectively, stratified by the weekday when the spells of absence began. In general, the closer was the beginning of the spell to the following weekend, the shorter was the spell, so that the most likely *ending* day was Friday. In both samples, in both survival and hazard plots, this effect is especially visible (in both of the following periods) after the 1987 reform, which restricted the coverage of the earnings lost only to scheduled work (which increased the probability of uncompensated weekends). For the survival functions, it is highlighted (with "balloons") for the LSIP sample in the period after the 1987 reform, but it remains equally visible after the 1991 reform as well. For the hazard functions, it is similarly highlighted for the IP sample, Generally speaking, for both samples and all three periods, absences which started on the weekend (especially Sunday) lasted the longest.

The hazard rates were much higher for the IP sample than for LSIP, meaning that people in the IP sample were more likely to return to work sooner, and the rates increased for both samples after the 1987 reform, and again after the 1991 reform, which indicates that in both regimes the spells of absenteeism due to sickness became shorter. This is expected after the 1991 reform, because the replacement rate was lowered, from 90% to 65% for the first 3 days, and to 80% from day 4.

Table 3 Survival functions (*s*) by the weekday when the short-term absences began, and by regime, and sample

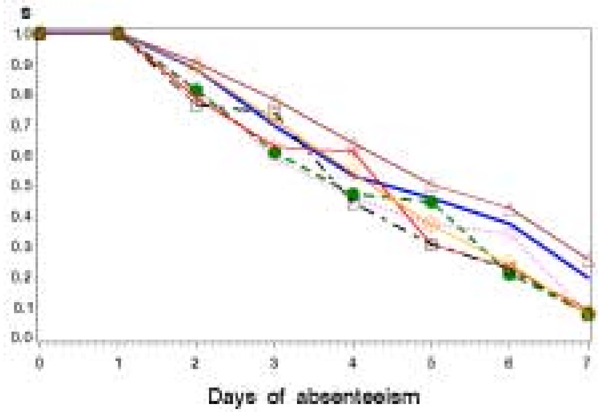
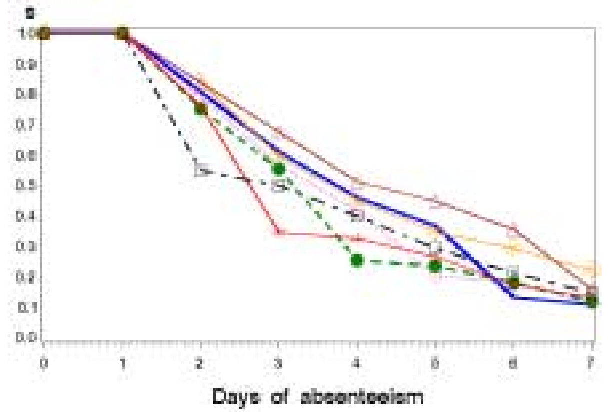
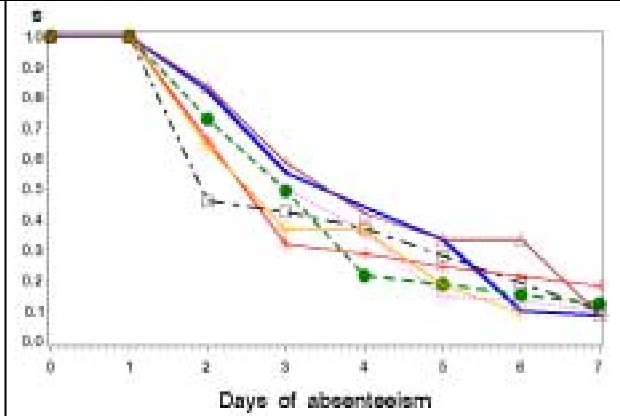
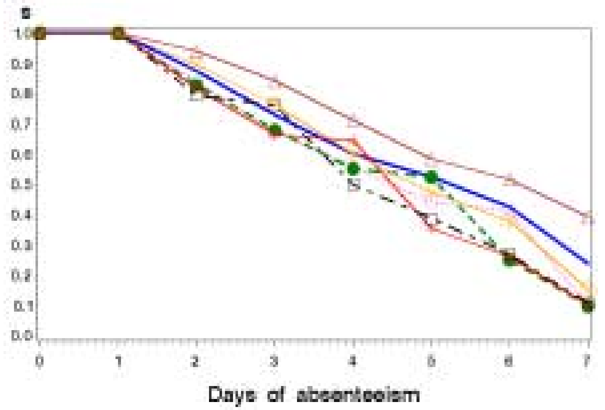
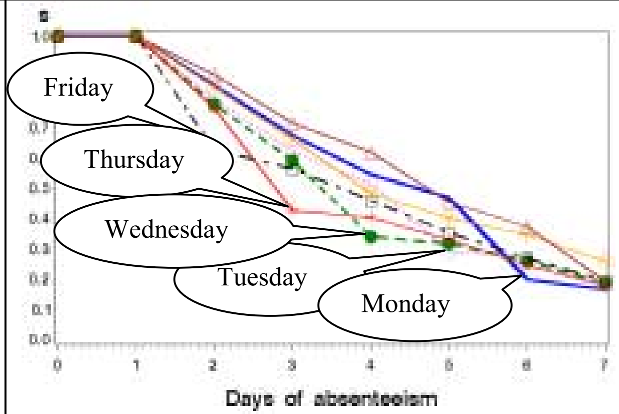
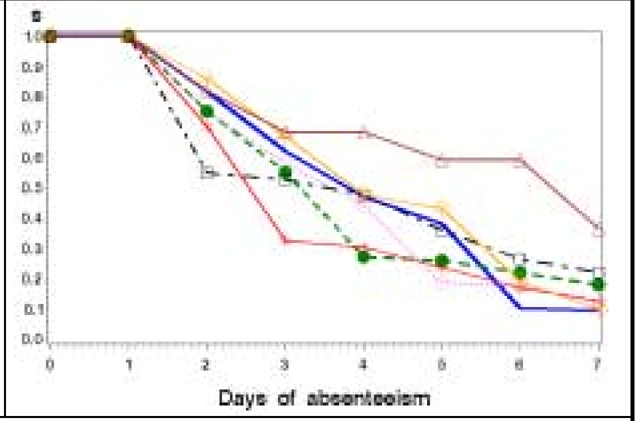
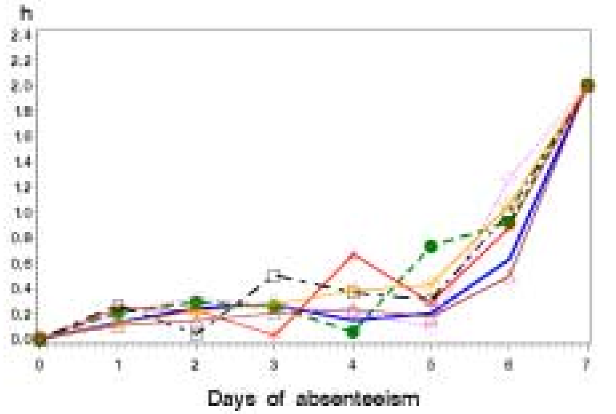
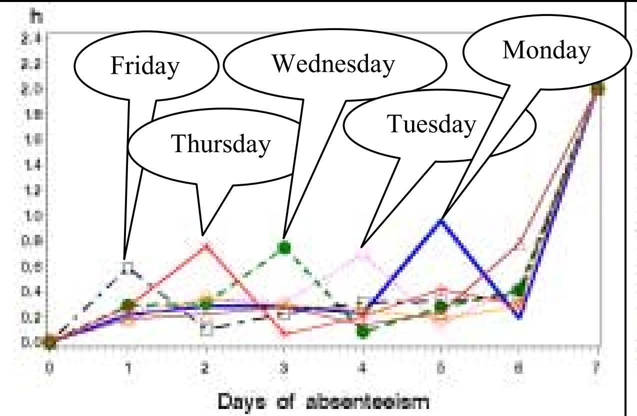
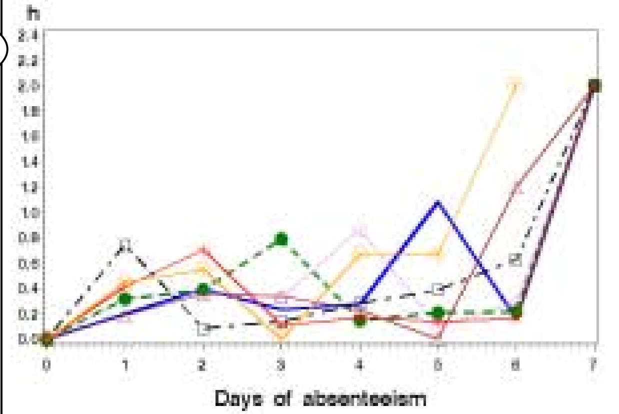
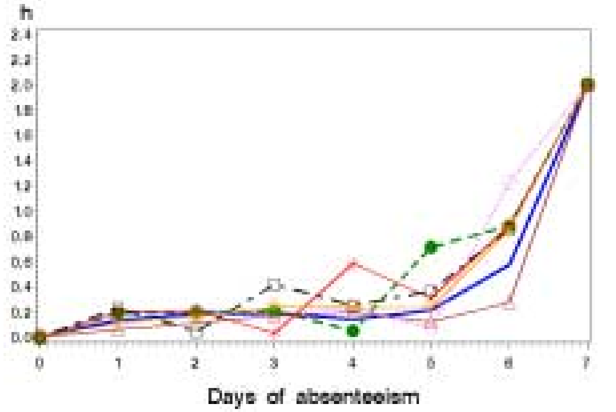
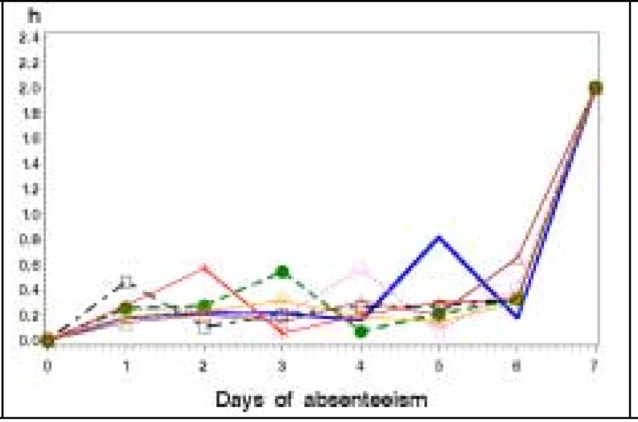
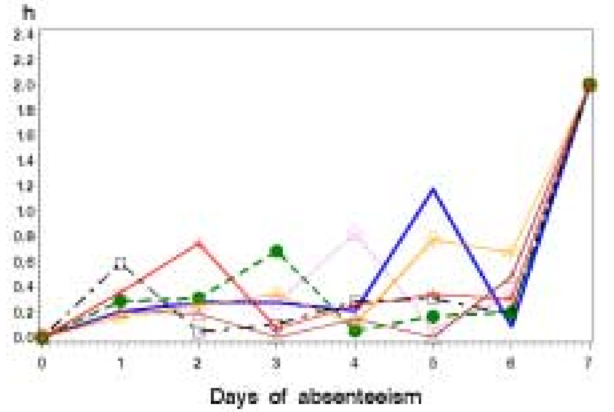
Data	Before the 1987 reform	December 1987 – February 1991	After the 1991 reform
IP			
LSIP			
Weekday when the spell began	Monday Tuesday Wednesday Thursday Friday Saturday Sunday		

Table 4 Hazard Functions (*h*) by the weekday when absences began, and by regime, and sample

<table>
<tr><th>Data</th><th>Before the 1987 reform</th><th>December 1987 – February 1991</th><th>After the 1991 reform</th></tr>
<tr><td>IP</td><td>
</td><td>
</td><td>
</td></tr>
<tr><td>LSIP</td><td>
</td><td>
</td><td>
</td></tr>
<tr><td colspan="4">Weekday when the spell began — Monday, Tuesday, Wednesday, Thursday, Friday, Saturday, Sunday</td></tr>
</table>

Thus, nonparametric analysis suggests that there were differences across regimes, and also that, during each regime, there was a significant relationship between the weekday when the absenteeism started and its duration; i.e. spells that started at the beginning of the week were longer than those that started at the end of week. The result that absences that started on a Sunday or Saturday after 1991 are the longest, can be interpreted as an effect of the lower replacement rates during the first 3 days.

In the *next* step of the analysis, the effect of various factors on short-term absence due to sickness was estimated using a semiparametric model. Table 5 shows coefficient estimates, standard errors, and hazard ratios of the gamma frailty model for both IP and LSIP samples during the entire period1986-1991, using dummies for the three policy regimes. Kendall's τ was quite small for both samples (about 0.05), which suggests very weak association within the groups, i.e., spells grouped by person. For both samples, there was thus a significant random effect related to the duration of short-term absences.

Women had a higher hazard of ending absenteeism within 7 days than did men, about 1.2 times higher for both samples. In general, the hazard of ending absenteeism was lower for older people, which means that younger people generally returned to work sooner.

For naturalized Swedes and other foreign born individuals, for both samples, the hazard of ending absenteeism was about 81-86% of that of Swedish born people, which means that Swedish born generally returned to work sooner. A poor health background, selection to specific work environments due to ethnic background, and/or cultural differences might explain this. For married people, the hazard of ending absenteeism was slightly higher than for singles; i.e., married employees returned to work sooner.

Although, as discussed in Appendix A2, the absence rate during summer months was the lowest during the year, the hazard ratios by quarter show that, for both samples, absences which began during summer lasted longest. This means that people use sick leave instead of vacation days.

Table 5 Estimation results for short-term absences during 1986-1991 (gamma frailty)

Variables	Insured Population (IP) Estimate	Standard error	Hazard ratio	LTS-Insured Population (LSIP) Estimate	Standard error	Hazard ratio
Frailty	**0.096**	0.008	1.101	**0.106**	0.007	1.111
Female (CG[a]: Male)	**0.194**	0.030	1.214	**0.181**	0.024	1.199
Age (CG: -35 years)						
36-45 years	-0.046	0.032	0.955	0.027	0.025	1.028
46-55 years	**-0.136**	0.035	0.873	-0.035	0.028	0.966
56-65 years	**-0.203**	0.044	0.816	-0.042	0.033	0.959
Citizenship (CG: Swedish born)						
Naturalized Swede	**-0.195**	0.061	0.823	**-0.187**	0.043	0.830
Other foreign born	**-0.151**	0.053	0.860	**-0.213**	0.039	0.808
Married (CG: Unmarried)	**0.090**	0.027	1.095	**0.089**	0.021	1.093
Quarter (CG: Winter)						
Spring	-0.007	0.026	0.993	-0.002	0.020	0.998
Summer	**-0.096**	0.026	0.908	**-0.079**	0.020	0.924
Autumn	-0.028	0.025	0.972	-0.015	0.020	0.985
Diagnosis (CG: Respiratory)						
Musculoskeletal	**-0.074**	0.033	0.929	0.004	0.023	1.004
Cardiovascular	-0.199	0.119	0.820	0.150	0.081	1.162
Mental	-0.180	0.111	0.835	0.027	0.060	1.028
General symptoms	**0.375**	0.030	1.454	**0.374**	0.023	1.454
Injuries and poisoning	**-0.127**	0.050	0.881	-0.017	0.038	0.983
Other	**0.376**	0.024	1.456	**0.410**	0.019	1.506
Weekday when absence started (CG: Weekend)						
Monday	**0.100**	0.044	1.105	**0.117**	0.033	1.125
Tuesday	**0.182**	0.045	1.200	**0.181**	0.034	1.198
Wednesday	**0.196**	0.046	1.216	**0.212**	0.034	1.236
Thursday	**0.193**	0.047	1.213	**0.248**	0.035	1.281
Friday	**0.178**	0.049	1.195	**0.200**	0.037	1.222
Previous cases [b]	-0.001	0.001	*-0.072*	0.001	0.001	*0.085*
Previous LTS[c] **cases**	-0.006	0.024	*-0.643*	-0.018	0.012	*-1.807*
Daily loss[d] (100 SEK)	**0.005**	0.001	*0.475*	**0.004**	0.001	*0.363*
Unemployment Rate	-0.005	0.011	*-0.500*	-0.013	0.008	*-1.258*
Regime (CG: before Dec 1987)						
Dec 1987 - Feb 1991	**0.127**	0.027	1.135	**0.090**	0.022	1.094
After Feb 1991	**0.275**	0.046	1.316	**0.253**	0.039	1.288
Kendall's τ	0.046			0.050		
	no Frailty	Frailty	Chi-Square	no Frailty	Frailty	Chi-Square
-2 Log Likelihood	54312.35	53901.91	410.4417	92021.48	91175.16	846.3198

Note: **Bolds** are significant for the IP sample at the 5% level, and for the LSIP sample at the 1% level; *Italics* for hazard ratio (*hr*) indicate that, for the continuous variables, it has been recomputed as *phr* = 100*(*hr*-1).

[a] CG indicates the comparison group.

[b] Previous cases of sickness before the analyzed spell, since January 1983, regardless of their duration.

[c] Previous cases of long-term sickness (LTS) before the analyzed spell, since January 1983, given that are at least 60 days of duration.

[d] Daily earnings loss due to absence.

Hazard ratios by diagnosis are rather different for the two samples. For the IP sample, persons with musculoskeletal, cardiovascular, and mental diagnoses, as well as injuries and poisonings, were slower to return to work than those with respiratory diagnoses. However, for the LSIP sample, except for injuries and poisonings, all diagnostic groups returned to work faster than those with a respiratory diagnosis.

For both samples, those whose absences started during the week (Monday-Friday) returned to work faster than those whose absences started on a weekend, and those whose absences started *earlier* in the week generally returned to work faster than did those whose absences that started later, although this trend was broken on *Friday* for both samples, and already on *Thursday* for IP.

Loss of earnings (due to sickness) is another factor that had a statistically significant impact on absence duration. For each 100 Swedish crowns in daily earnings loss, the hazard of ending an absenteeism spell went up by about 0.4-0.5%.

Neither of previous sickness history variables (total cases and total LT cases) was found to be a significant determinant of short-term absence duration. Regional unemployment also failed to pass the significance test.

The regime dummies are also statistically significant, and show that the 1991 reform had an especially strong impact on absenteeism. After the 1987 reform, people in both samples were more likely to return to work sooner (hazard ratios: 1.13 and 1.09), and even more so after the 1991 reform (hazard ratios: 1.32 and 1.29). Given the differences on the magnitude of regime dummies, a separate analysis was also done for each regime. Table 6 shows similar estimation results for the IP sample alone, but divided into the three regimes. The gamma frailty model was estimated for the first two regimes, but a standard Cox model for the last.[11] Table 7 shows estimation results from the gamma frailty model for the LSIP sample alone, also divided in three regimes.

[11] The data for the last regime of the IP sample did not support the gamma frailty model, possibly due to a short time horizon for this regime. The EM algorithm computes a likelihood assuming independence, i.e., $\theta = 0$, and then increases this values until it finds a likelihood which is larger than the likelihood at $\theta = 0$. From there it starts a numerical routine to find the root. If it cannot find that point, then it is considered the independence case. Therefore for the IP sample, only the first spell of short-term absence after the 1991 reform was used, i.e., 559 spells out of 967 total. For the other two regimes, and for all three regimes with the LSIP sample, there is a significant random effect related to absence duration by person.

Table 6 Estimation results for short-term absences, by regime, IP sample

Variables	Before Dec 87 (n = 3580)			Dec 87 – Feb 91 (n =8326)			After Feb 91 (n = 559)		
	Estimate	S.E.	HR	Estimate	S.E.	HR	Estimate	S.E.	HR
Frailty	**0.07**	0.02	1.07	**0.08**	0.01	1.08			
Female (CG[a]: Male)	**0.22**	0.05	1.25	**0.21**	0.03	1.24	**0.26**	0.10	1.30
Age (CG: -35 years)									
36-45 years	-0.03	0.05	0.97	-0.05	0.04	0.95	-0.17	0.12	0.85
46-55 years	**-0.19**	0.06	0.83	**-0.11**	0.04	0.90	**-0.21**	0.12	0.81
56-65 years	**-0.29**	0.08	0.75	**-0.14**	0.05	0.87	**-0.38**	0.14	0.68
Citizenship (CG: Swedish Born)									
Naturalized Swede	-0.08	0.09	0.93	**-0.20**	0.07	0.82	**-0.40**	0.22	0.67
Foreign born	-0.13	0.08	0.88	**-0.14**	0.06	0.87	-0.26	0.18	0.77
Married (CG: Unmarried)	**0.13**	0.04	1.14	**0.08**	0.03	1.08	0.02	0.09	1.02
Quarter (CG: Winter)									
Spring	0.01	0.05	1.01	0.00	0.03	1.00	-0.09	0.11	0.92
Summer	-0.06	0.05	0.94	**-0.09**	0.03	0.91	-0.25	0.13	0.78
Autumn	0.05	0.05	1.05	-0.04	0.03	0.96	0.19	0.43	1.21
Diagnosis (CG: Respiratory)									
Musculoskeletal	**-0.15**	0.06	0.86	-0.05	0.04	0.95	-0.15	0.15	0.86
Cardiovascular	**-0.41**	0.25	0.66	-0.15	0.15	0.86	-0.24	0.52	0.79
Mental	**-0.32**	0.19	0.73	-0.14	0.15	0.87	-0.45	0.43	0.64
General symptoms	**0.19**	0.06	1.21	**0.39**	0.04	1.48	**0.73**	0.13	2.07
Injuries & poisoning	**-0.18**	0.09	0.84	**-0.13**	0.07	0.88	0.13	0.26	1.13
Other	**0.24**	0.05	1.28	**0.42**	0.03	1.52	**0.47**	0.11	1.59
Weekday when absence started (CG: Weekend)									
Monday	-0.02	0.07	0.98	**0.15**	0.07	1.16	0.24	0.29	1.27
Tuesday	**0.16**	0.07	1.17	**0.21**	0.07	1.23	0.36	0.29	1.43
Wednesday	**0.23**	0.07	1.26	**0.24**	0.07	1.27	0.21	0.30	1.23
Thursday	**0.20**	0.07	1.22	**0.25**	0.07	1.28	0.21	0.31	1.24
Friday	**0.25**	0.07	1.28	**0.20**	0.07	1.22	0.29	0.30	1.33
Previous cases [b]	**-0.01**	0.01	*-1.00*	0.00	0.00	*-0.20*	-0.01	0.00	*-0.48*
Previous LTS[c] cases	-0.12	0.09	*-11.51*	-0.03	0.03	*-2.49*	0.07	0.07	*6.74*
Daily loss[d] (100 SEK)	**0.01**	0.00	*0.78*	**0.01**	0.00	*0.47*	0.00	0.00	*0.18*
Unemployment rate	-0.02	0.02	*-2.14*	0.01	0.02	*0.93*	-0.02	0.05	*-2.07*
Kendall's τ	0.03			0.04					
	No frailty	Frailty	Chi-Sq.	No frailty	Frailty	Chi-Sq.	No covariate	Cov.	Chi-Sq.
-2 Log Likelihood	15462.6	15422	39.9	34583.4	34435	148.9	6254	6174	79.7

Note: **Bolds** are significant for the IP sample at the 5% level, and for the LSIP sample at the 1% level; *Italics* for hazard ratio (hr) indicate that, for the continuous variables, it has been recomputed as phr = 100*(hr-1). [a] CG indicates the comparison group.
[b] Previous cases of sickness before the analyzed spell, since January 1983, regardless of their duration.
[c] Previous cases of long-term sickness (LTS) before the analyzed spell, since January 1983, given that are at least 60 days of duration. [d] Daily earnings loss due to absence.

Table 7 Estimation results for short-term absences, by regime, LSIP sample

	Before Dec 87 (n = 3580)			Dec 87 – Feb 91 (n =8326)			After Feb 91 (n = 559)		
Variables	Estimate	S.E.	HR	Estimate	S.E.	HR	Estimate	S.E.	HR
Frailty	**0.08**	0.01	1.08	**0.10**	0.01	1.11	**0.07**	0.03	1.08
Female (CG[a]: Male)	**0.19**	0.03	1.20	**0.19**	0.03	1.21	0.13	0.08	1.14
Age (CG: -35 years)									
36-45 years	0.02	0.04	1.02	**0.05**	0.03	1.06	0.02	0.09	1.02
46-55 years	-0.05	0.04	0.96	-0.02	0.03	0.98	-0.02	0.09	0.98
56-65 years	-0.07	0.05	0.93	-0.02	0.04	0.98	-0.02	0.12	0.98
Citizenship (CG: Swedish Born)									
Naturalized Swede	**-0.13**	0.06	0.88	**-0.22**	0.05	0.81	-0.17	0.14	0.85
Foreig born	**-0.19**	0.06	0.83	**-0.22**	0.05	0.80	-0.14	0.12	0.87
Married (CG: Unmarried)	**0.12**	0.03	1.12	**0.08**	0.03	1.09	**0.14**	0.07	1.15
Quarter (CG: Winter)									
Spring	0.01	0.04	1.01	-0.01	0.03	0.99	-0.03	0.10	0.98
Summer	**-0.02**	0.04	0.98	**-0.09**	0.03	0.92	**-0.17**	0.10	0.84
Autumn	0.05	0.05	1.05	-0.04	0.03	0.96	-0.39	0.21	0.68
Diagnosis (CG: Respiratory)									
Musculoskeletal	-0.01	0.04	0.99	0.01	0.03	1.01	-0.02	0.09	0.98
Cardiovascular	0.14	0.13	1.15	**0.27**	0.11	1.31	-0.53	0.34	0.59
Mental	-0.14	0.10	0.87	0.11	0.08	1.11	-0.30	0.32	0.74
General symptoms	**0.27**	0.04	1.31	**0.43**	0.03	1.54	0.13	0.11	1.14
Injuries & poisoning	-0.03	0.06	0.97	-0.02	0.05	0.98	-0.23	0.17	0.79
Other	**0.31**	0.03	1.37	**0.46**	0.03	1.59	**0.15**	0.09	1.16
Weekday when absence started (CG: Weekend)									
Monday	**0.11**	0.05	1.11	**0.08**	0.05	1.09	0.23	0.19	1.25
Tuesday	**0.26**	0.05	1.30	**0.14**	0.05	1.14	0.26	0.19	1.29
Wednesday	**0.35**	0.05	1.42	**0.16**	0.05	1.18	0.23	0.20	1.25
Thursday	**0.37**	0.05	1.45	**0.19**	0.05	1.21	**0.35**	0.19	1.42
Friday	**0.37**	0.05	1.44	**0.16**	0.06	1.17	-0.31	0.22	0.74
Previous cases [b]	0.00	0.00	-0.34	0.00	0.00	0.18	0.00	0.00	0.06
Previous LTS[c] cases	**-0.05**	0.03	-5.23	**-0.04**	0.01	-3.60	-0.02	0.03	-2.05
Daily loss[d] (100 SEK)	**0.01**	0.00	0.50	**0.00**	0.00	0.36	0.00	0.00	0.19
Unemployment rate	0.00	0.01	-0.14	**-0.03**	0.01	-2.50	0.02	0.04	1.86
Kendall's τ	0.04			0.05			0.04		
	No frailty	Frailty	Chi-Sq.	No frailty	Frailty	Chi-Sq.	No frailty	Frailty	Chi-Sq.
-2 Log Likelihood	30854	30754	100	54469	54077	392	19231	19110	121

Note: **Bolds** are significant for the IP sample at the 5% level, and for the LSIP sample at the 1% level; *Italics* for hazard ratio (hr) indicate that, for the continuous variables, it has been recomputed as phr = 100*(hr-1).

[a] CG indicates the comparison group.

[b] Previous cases of sickness before the analyzed spell, since January 1983, regardless of their duration.

[c] Previous cases of long-term sickness (LTS) before the analyzed spell, since January 1983, given that are at least 60 days of duration.

[d] Daily earnings loss due to absence.

Kendall's τ was quite small (about 0.04) for both samples and all regimes, which suggests very weak association within the groups. During all three regimes for both samples, women returned to work faster than men (their hazard ratios were in the range 1.14-1.30). Differences between the IP and LSIP samples might be due to the different average health status of the two samples (those from LSIP returned to work slower than did those from IP). Differences between the first two and the third regime, for IP and LSIP, might relate to changes in the replacement rate.

In general, as one might expect, younger people returned to work faster than did older people although, not all estimates were statistically significant, especially for LSIP.

Although all the results did not meet the statistical significance test, the hazard of ending absenteeism for foreign-born people (whether naturalized or not) was always lower than that for Swedish born, across all regimes, and in fact, for both samples, seemed to go down after the first reform, and at least for the IP sample, it went down further after the second.

For the IP sample, the hazard for ending absenteeism was higher for married people than for those who were unmarried, though it fell after the first reform, and still further after the second. For the LSIP sample, the fall after the first reform was reversed after the second.

For the IP sample, during all three regimes the hazard of ending absenteeism was lower for those with musculoskeletal, cardiovascular, and mental diagnoses, as well as those with injuries or poisonings, compared to those with respiratory diagnoses. During the third regime, this was also true for the LSIP sample. During the other two regimes, however, for the LSIP sample people with a diagnosis of general symptoms had a higher hazard of ending absenteeism that did people with a respiratory diagnosis.

The results again show that there was a *timing of absenteeism* with respect to the weekday when the spell began, after the 1987 reform. For both IP and LSIP samples, regardless of which weekday their absence began on, employees were less likely to return to work sooner (compared to the weekend) before the first reform, and even after this reform (except for Friday), but the trend disappeared after the second reform. On the other hand, the magnitude of the impact of the first reform was not the same for the IP and LSIP samples, which might be related to the different health status of the two samples, though exactly how or why is not obvious. Moreover, for both samples, loss of earnings had a very weak effect. For each 100 Swedish crowns increase in daily earnings loss, the hazard of ending an absenteeism spell went up by about one half percent.

The level of regional unemployment had no significant effect on absence duration before the 1987 reform, nor after the 1991 reform, but it had a significant *negative* effect for the LSIP sample during the middle regime (December1987-February 1991). Each additional percentage point of regional unemployment was then associated with about a 2.5% decrease in the hazard of ending the absenteeism spell.

8 Summary and conclusions

Both the nonparametric and the semiparametric analyses suggest that there were differences between the IP and LSIP samples (i.e., insured people, and insured people with poorer health) for the *entire* period, but also *across regimes*. Differences between the IP and LSIP samples might be due to the *different* average health status of the two samples, while the differences across regimes might be due to *different* replacement rates (i.e., different economic incentives).

The *nonparametric analysis* suggests that there was a significant relationship between the *weekday* when the absenteeism started and its duration. In general, the closer was the beginning of the spell to the following weekend, the shorter was the spell, so that the most likely *ending* day was Friday. In both samples, in both survival and hazard plots, this effect is especially visible (in both of the following periods) after the 1987 reform, which restricted the coverage of the earnings lost only to scheduled work, which increased the probability of uncompensated weekends. After both the 1987 and the 1991 reform, spells became shorter. For the 1987 reform, this indicates that the effect of limiting the compensation to only scheduled work was (much) stronger than the effect of eliminating the first waiting day. This is an expected result after the 1991 reform, because the replacement rate was lowered, from 90% to 65% for the first 3 days, and to 80% from day 4. The result that absences that started on a Sunday or Saturday after 1991 are the longest can be interpreted as an effect of the lower replacement rates during the first 3 days. Another general result across regimes was that the hazard rates were much *higher* for the IP sample than for LSIP, meaning that people in the IP sample were more likely to return to work *sooner*.

The *semiparametric analysis for the entire period* tells us that while the *direction* of impact of the *determinants* of the absence duration was the same, the *magnitude* of their impact was different. Women had a higher hazard of ending absenteeism within 7 days than did men, and in general, the hazard of ending absenteeism was lower for older people, which means that younger people generally returned to work sooner. For naturalized Swedes and other foreign born individuals, for both samples, the hazard of ending absenteeism was lower than that of Swedish born people, which means that *Swedish born persons* generally returned to work *sooner*. For married people, the hazard of ending absenteeism was slightly higher than for singles, i.e., *married* employees returned to work *sooner*. For both samples, absences started during the week (Monday-Friday) were shorter than absences started on a weekend, in line with the nonparametric analysis, and absences during the summer lasted longest. The regime dummies are also statistically significant, and show that while the direction of the impact of both 1987 and 1991 reform was the *same* for the IP and LSIP samples, the *magnitude* of the impact of both reforms was higher for the IP sample. This last result might be related to the different health status of the two samples, as both reforms had a higher positive impact on the hazard of those from the IP sample. This means that even though economic incentives mattered, people with poorer health did not “shorten” their absences in the same extent as those with better health.

The *semiparametric analysis for the three regimes* estimated separately suggests that women returned to work faster than men. For the IP sample, the hazard for ending absenteeism was higher for married people than for those who were unmarried, though it fell after the first and second reforms. For the LSIP sample, the fall after the first reform was reversed after the second. Except general symptoms, and other diagnoses, people with respiratory diagnoses (mainly, *common colds*) get well faster, and this is generally true for both IP and LSIP samples. For both IP and LSIP samples, regardless of which weekday their absence began on, people were less likely to return to work sooner (compared to the weekend) before the first reform, and even after this reform (except for Friday), but the tendency disappeared after the second reform. Moreover, for both samples, loss of earnings had a very weak effect: For each 100 Swedish crowns increase in daily earnings loss, the hazard of ending an absenteeism spell went up by about one half percent. The level of regional unemployment had no significant effect on absence duration before the 1987 reform, nor after the 1991 reform, but it had a significant *negative* effect for the LSIP sample during the middle regime (December1987-February 1991).

In sum, the 1991 reform, which reduced the replacement rate, had a stronger effect on reducing the duration of short-term absences than the 1987 reform, which restricted the payment of sickness cash benefit to only scheduled workdays. After the 1987 reform, fewer reported sickness starting on the weekend, and more on Monday. Generally, the closer to the end of the week was the beginning of the absence, the shorter was the spell. The change in the frequency of spells by the weekday when they started, before and after the 1987 reform (i.e., fewer absences started on weekend, more on Monday), may be explained by the existence of a waiting day prior December 1, 1987, while the change in the frequency of spells by the weekday when they ended (i.e., the most ended on Friday) can be explained by the restriction of the coverage only to the scheduled days of work. In conclusion, the *rules* clearly influenced people's decisions about *when* to report the beginning and ending of sickness spells.

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Appendix A1 To get equation (5)

(A1) $\frac{\partial U}{\partial t_w} = -\lambda w + \mu$

(A2) $\frac{\partial U}{\partial t_a} = -\lambda \rho w + \mu$

(A3) $\frac{\partial U}{\partial t_l} = \mu$,

where $\frac{\partial U}{\partial t_j} = \frac{\partial U}{\partial Z_j} \times \frac{\partial f_j}{\partial t_j}$.

By substitution for μ from (A1) into (A2) and (A3) we get

(A4) $\lambda w = \frac{\partial U}{\partial t_l} - \frac{\partial U}{\partial t_w}$

(A5) $\lambda \rho w = \frac{\partial U}{\partial t_l} - \frac{\partial U}{\partial t_a}$,

From (A4) and (A5) results

(A6) $\lambda(1-\rho) w = \frac{\partial U}{\partial t_a} - \frac{\partial U}{\partial t_w}$,

which is equation (5).

Appendix A2 Descriptive statistics for the IP and LSIP samples

Table A1 shows descriptive statistics of both IP and LSIP *full* samples by individual. The LSIP is slightly older on average. It also contains more women, more single persons with deceased spouse, and more persons with lower education and with lower earnings than the IP sample.

Table A1 Descriptive statistics of individuals in the IP and LSIP samples, 1991

Variable	Insured population (N=1813)		Long-term sick insured (N=2761)	
	Mean	Std Dev	Mean	Std Dev
Age	44.56	11.25	47.57	11.88
Gender (1= Female, 0=Male)	0.49	0.50	0.55	0.50
Citizenship				
Swedish born	0.88	0.33	0.85	0.36
Foreigner born	0.07	0.26	0.08	0.28
Nationalized Swede	0.05	0.22	0.07	0.26
Marital Status				
Unmarried	0.17	0.38	0.17	0.38
Married	0.74	0.44	0.68	0.47
Divorced	0.08	0.27	0.12	0.33
Widow/er	0.01	0.10	0.02	0.14
Level of education				
Low	0.49	0.50	0.63	0.48
Medium	0.35	0.48	0.28	0.45
High	0.16	0.37	0.09	0.28
Annual earnings (deflated using 1997 CPI)	183365	100998	146795	87233
Days of absenteeism (1986-1991)	18.22	21.27	20.18	22.74
Absent during 1986-1991	0.78	0.43	0.78	0.43

Note: *Italics* indicate dummy variables.

The data sets provide exact dates and defined states for the beginning and end of each compensated sickness spell, as well as diagnosis. Table A2 shows the characteristics of those individuals in both IP and LSIP samples with short-term (ST) sickness during 1986-1991. Even though those in the LSIP sample were long-term (LT) sick at least once during 1986-1991, the two samples were very similar with respect to short-term absenteeism. The average number of days absent due to sickness was only slightly higher for the LSIP, and slightly higher in each diagnosis category as well. The average number of spells of ST sickness spells was also slightly higher for the LSIP-sample, and also by diagnosis, with the single exception of the number of respiratory spells. In both samples, the average number of spells that began on a Monday was higher than that of those that began on any other weekday, and the number decreased through the week. This “non-randomness” suggests a kind of “timing” of absenteeism, or perhaps a tendency for spells to conclude by the end of the following weekend.

Table A2 Descriptive statistics of reported ST absences due to sickness by individual, IP and LSIP samples, 1986-1991

Variables	IP (N=1416)		LSIP (N=2165)	
	Mean	Std Dev	Mean	Std Dev
Total days ST absent due to sickness, 1986-91	30.85	29.75	37.26	33.19
Total days of ST sickness, 1986-91, by diagnosis				
Musculoskeletal	3.71	8.68	6.13	11.44
Cardiovascular	0.23	1.47	0.30	1.55
Respiratory	15.66	16.60	16.15	16.74
Mental	0.31	2.75	0.75	5.38
General symptoms	3.34	6.06	4.48	7.82
Injuries and poisonings	1.39	3.78	1.82	5.24
Others	6.21	9.12	7.64	11.18
Total number of ST sickness spells, 1986-91	9.11	8.19	9.88	8.80
Total number of ST spells, 1986-91, by diagnosis				
Musculoskeletal	0.93	2.12	1.44	2.59
Cardiovascular	0.06	0.35	0.08	0.42
Respiratory	4.25	4.10	3.85	3.84
Mental	0.07	0.65	0.18	1.17
General symptoms	1.20	2.02	1.42	2.38
Injuries and poisonings	0.34	0.81	0.42	1.15
Others	2.27	3.20	2.49	3.45
Number of spells, by the weekday they began				
Monday	2.53	2.73	2.82	3.06
Tuesday	1.95	2.27	2.05	2.30
Wednesday	1.67	2.05	1.73	2.04
Thursday	1.46	1.77	1.59	2.03
Friday	1.02	1.48	1.09	1.49
Weekend	0.48	1.01	0.60	1.13

As Figure A1 shows, for both samples, the frequency of one- and two-day spells increased substantially after the 1987 reform eliminated the unpaid “waiting day”, while the frequency of six-day spells decreased dramatically.[12] The increases suggest strongly that the number of ST absences due to sickness was affected by the availability of a sickness cash benefit from the first day after the reform. On the other hand, the decrease in six-day (but not seven-day) spells might be interpreted as a “timing” of absenteeism.

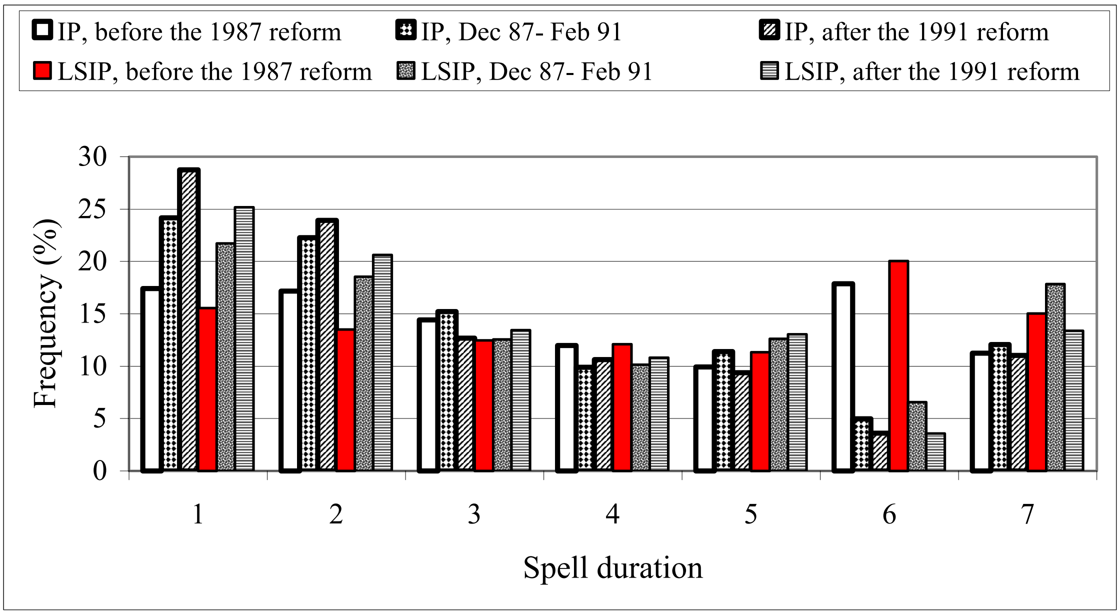


Figure A1 Distribution of ST sickness spell-durations, IP and LSIP, under three regimes

After the 1991 reform reduced the replacement rate from 90% to only 65% of lost earnings for the first three days (and then to 80% through the 90th day) people seem to have returned to work sooner; i.e., the proportion of one- and two-day absences again increased for both samples, and the proportion of 6-day (and even 7-day) absences again decreased.

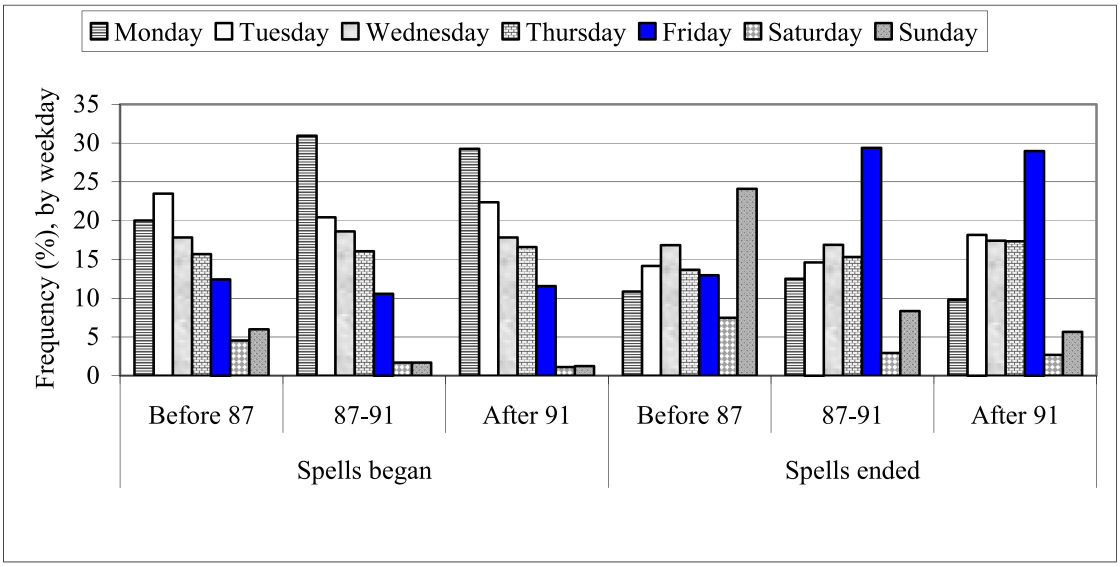
[12] Before the 1987 reform, people could in principle call in sick on any day and then be compensated for up to seven of the first seven days without a medical certificate. Social insurance did not cover the first day, but there were collective agreements that covered even this day for some occupations, such as day-care and restaurant personnel, where it was thought especially important to shield customers/clients from infectious diseases. Nevertheless, many more people seem to have reported very short illness after social insurance started paying from the first day.

Regardless of the regime, the proportion of one- and two-days spells was higher for the IP sample than for the LSIP sample, while the proportion of five-, six- and seven- day spells was higher for LSIP-sample, compared to IP-sample.

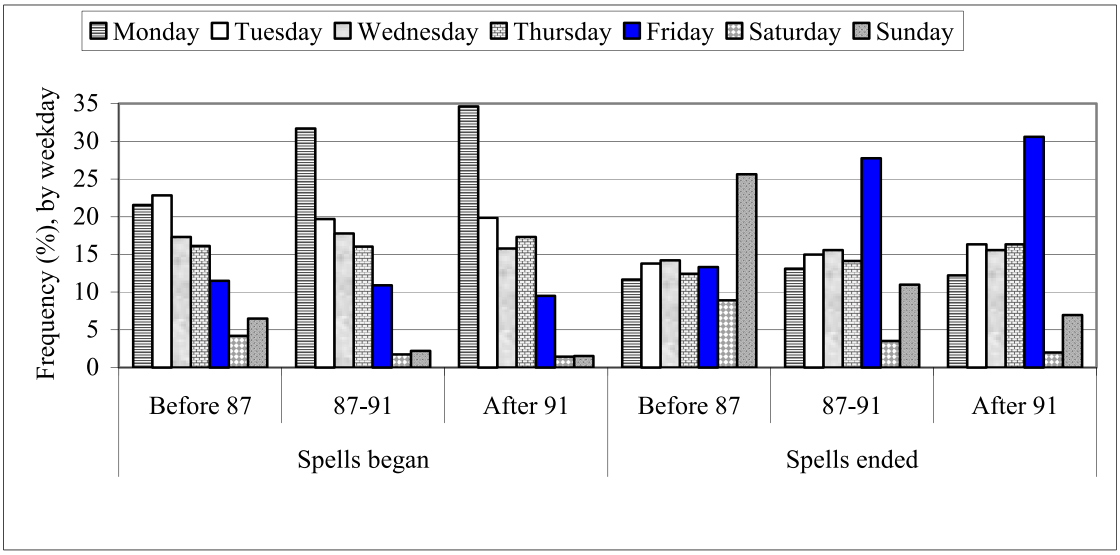
Figures A2 a) and b) show, for the two samples, how the three regimes compared with respect to the weekday when reported short-term absences began and ended, while Figures A3 a) and b) show the overall distribution for the entire period, i.e., for the three regimes pooled together.

Before the first reform, the highest proportion of reported short-term sickness, for both samples, started on Tuesdays. In principle, the database is designed to record all days of sickness, including uncompensated days (such as waiting days and regular non-working days). However, the Tuesday phenomenon could indicate that some spells recorded then during the first regime actually began on Monday (which would have been the unpaid waiting day). After the 1987 reform, this sort of “confusion” would have disappeared, and Monday clearly became the most “popular” starting day, for both samples, during the remaining two regimes. After the 1987 reform eliminated coverage on non-working days, there were also fewer spells reported starting on the weekend.

For both samples, the 1987 reform clearly had a big impact on the weekday when spells ended. Before the reform, spells ended most often on Sunday (regardless of when they started), but afterwards they ended most often on Friday. The 1991 reform made little difference in this respect.

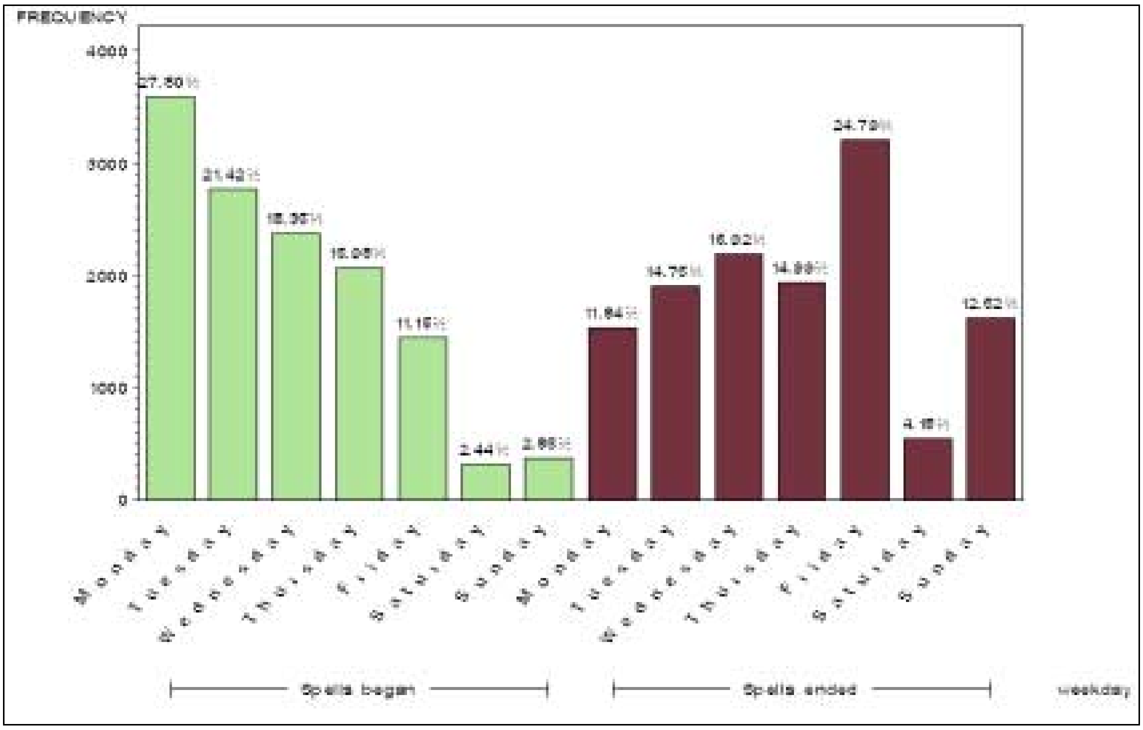


a) Insured population (IP)

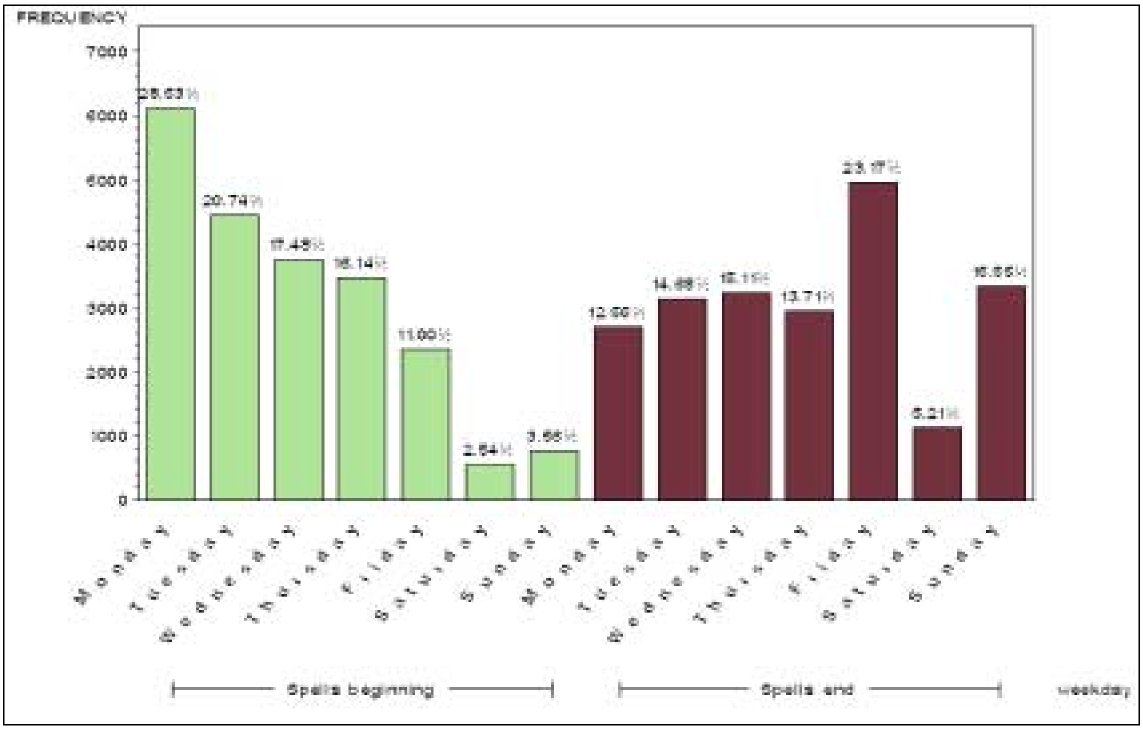


b) Long-term sick insured population (LSIP)

Figure A2 Distribution of ST sickness spells by the day they began and ended, IP and LSIP, under three regimes



a) Insured Population (IP)



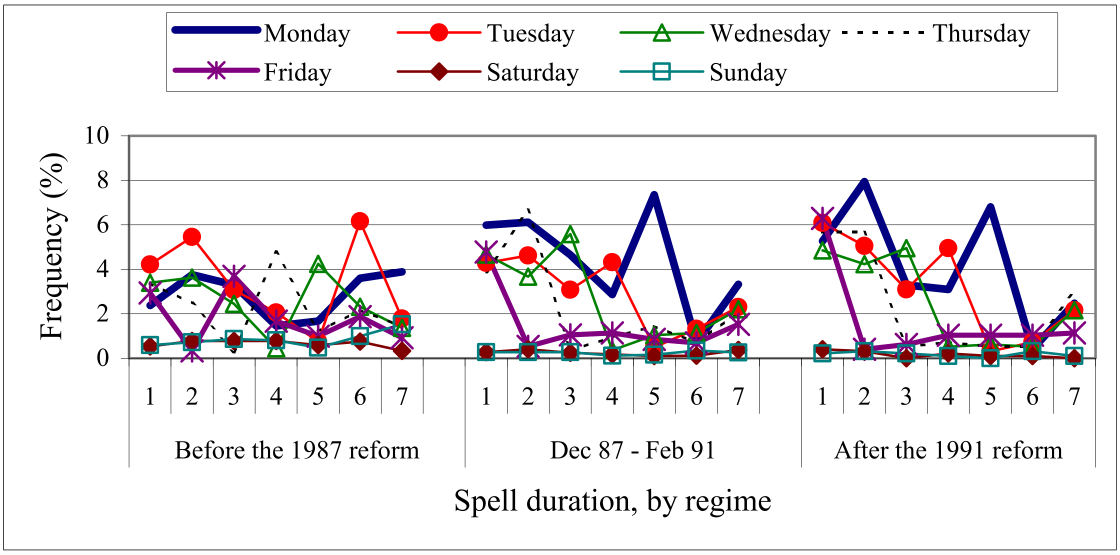
b) Long-Term Sick, Insured Population (LSIP)

Figure A3 Distribution of pooled ST sickness spells by the day they began and ended, IP and LSIP, 1986-1991

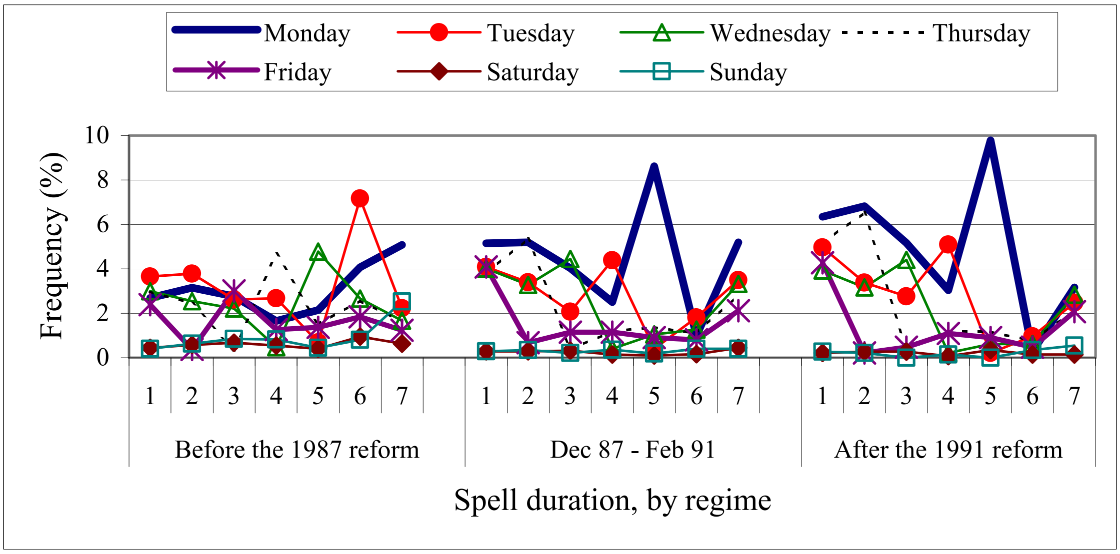
Figures A4 a) and b) illustrate, for the two samples, the distribution of ST absences by duration and the weekday when they began, during the three regimes. As we also saw in Figure A2, before the 1987 reform there were more spells starting on Tuesday, afterwards more on Monday (including after the 1991 reform). Of those starting on Monday after the 1987 reform, the percentage of one- and two-day absences was slightly higher for the IP sample than for the LSIP sample (as was also visible in Figure A1, regardless of the starting day), while the percentage of five- and seven-day absences was slightly higher for LSIP sample. For both samples, before the 1987 reform, the highest percentage of reported sickness started on Tuesday and lasted six days (i.e., ending on Sunday), whereas after the reform, the highest percentage started on Monday and lasted five days (i.e., ending on Friday).

After the 1991 reform, two-day spells starting on Monday were the most frequent for the IP sample, while five-day spells starting on Monday were most frequent for the LSIP sample.

Before the 1987 reform, it seems that, regardless of the weekday when they started, most spells ended on Sunday, whereas afterwards most spells ended on Friday (i.e., before the weekend). Nevertheless, afterwards most spells *started* on Monday.



a) Insured population (IP)



b) Long-term sick insured population (LSIP)

Figure A4 Distribution of ST sickness spells by duration and the day they began, IP and LSIP, under three regimes

In sum, regardless of the rule changes, it seems that there was no big difference between insured people in general (the IP sample) and insured people with long-term sickness (the LSIP sample) with respect to the distribution of weekdays when ST absences began and ended.

As Figure A5 shows, during at least the first two regimes,[13] there was also little difference between the IP and LSIP samples with respect to the distribution of absences by the month when they began, and the slight difference *between* the two regimes is not necessarily the result of the 1987 reform. Rather, the higher percentage of spring and fall cases during the earlier period could reflect different epidemiological conditions during that period, i.e., more virulent colds, and/or flu. There were fewer spells of ST absence due to sickness reported during summer, when most people take vacations, even though they were entitled to sickness cash benefit even then (and thus, if they were sick, could save their vacation for later). But, besides indicating "honesty" in reporting, it is also possible that summer is generally a healthier time, especially with lower depression (caused by the "darkness" of other seasons) and less stress generally, because of vacations.

[13] The third regime (after the 1991 reform) is not included because the observation period ended in December 1991, and thus would not allow analyzing a full 12 months.

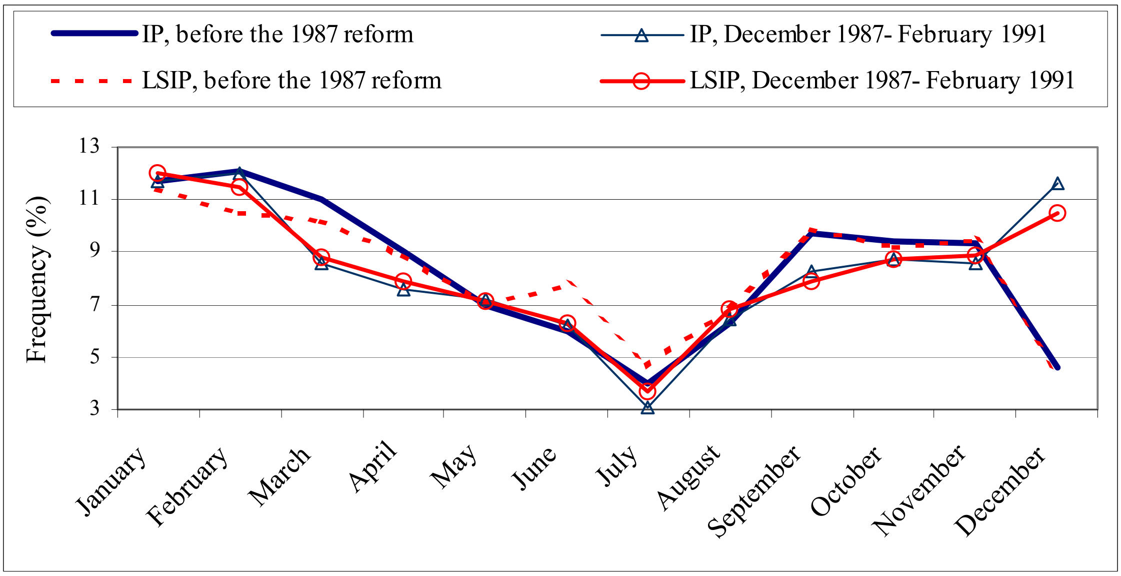
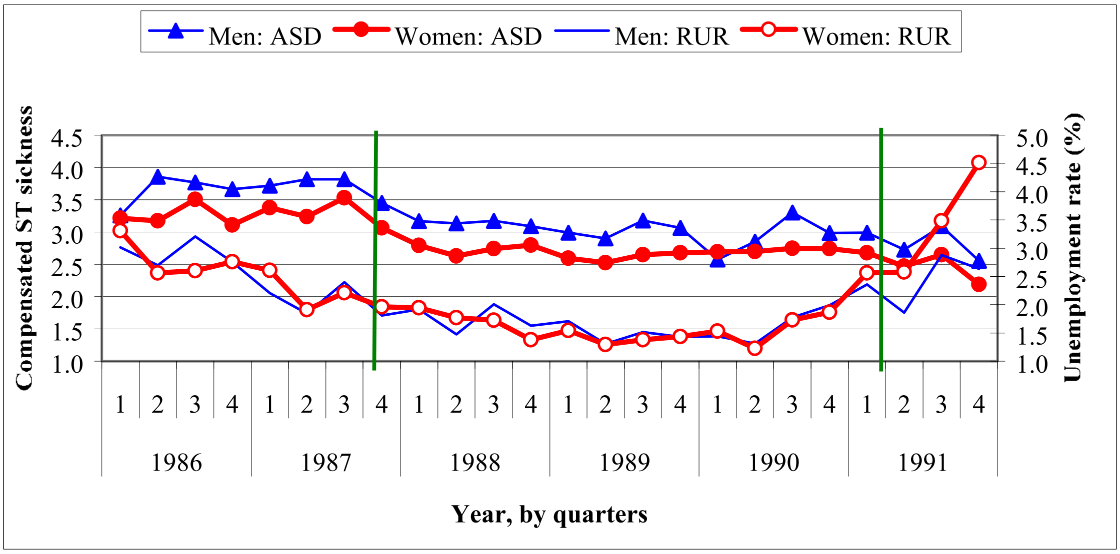
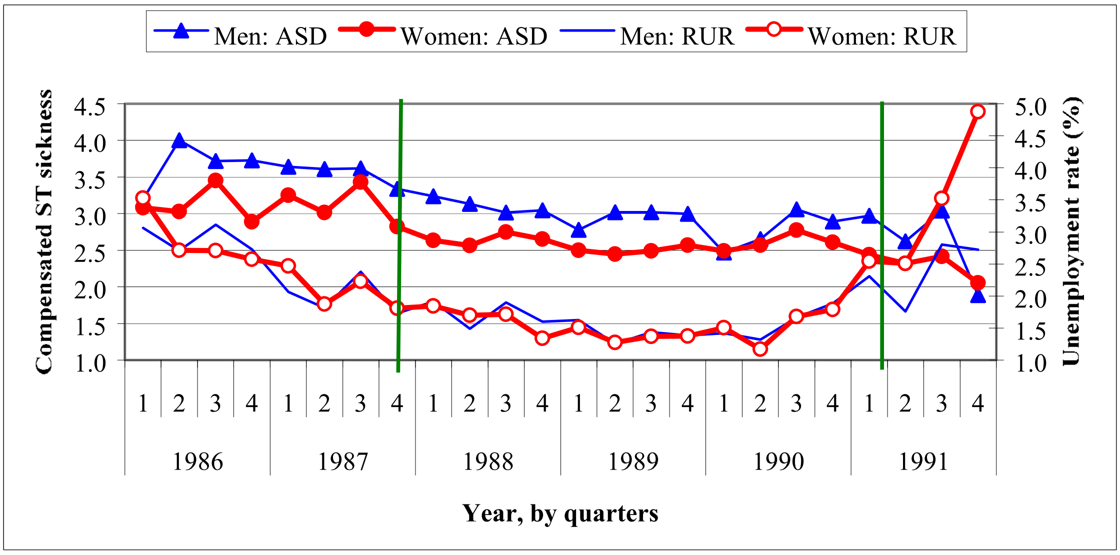


Figure A5 Distribution of ST sickness spells by month when they began, IP and LSIP, under two regimes

Figures A6 a) and b) show, for each sample, the average short-term sickness duration (ASD) and the average regional unemployment rate (RUR) during 1986-1991, by quarter and gender. Men's durations were generally longer than women's. Unemployment was generally declining until mid-1990, and average ST sickness durations also generally declined correspondingly. After that, unemployment increased spectacularly, while durations remained virtually unchanged, or fell even further. The direction of any effect of regional unemployment on the duration of short-term absences due to sickness is thus not clear.



a) Insured population (IP)



b) Long-term sick insured population (LSIP)

Figure A6 Average ST sickness duration (ASD) and regional unemployment rate (RUR), by gender, IP and LSIP, 1986 -1991