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Estimating the Impact of Experience Rating on the Inflow into Disability Insurance in the Netherlands

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Abstract
This paper examines the effects of experience rating on the inflow into disability insurance (DI) in the Netherlands, using unique longitudinal administrative data from the social benefit administration. We follow a difference-in-differences approach to identify the impact of changes in DI premiums. Due to unawareness of the experience rating system, employers seem to have been triggered to increase preventative activities, once they have experienced increases in DI premium. We find this impact to be substantial, amounting to a 15% reduction of the DI inflow.

Keywords: experience rating, disability insurance, panel data

JEL classification: H22, I12, C23

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1. Introduction

As of 1998, disability insurance (DI) in the Netherlands is financed by premiums that are experience rated. This means that, in principle, employers bear the costs of the first five years of DI benefits. Initially, the experience rating system did not cause substantial controversy among employers and policy makers. It appears this has been due to the nature of the incentive system itself: each year, a new cohort of disability benefit costs has been added to the disability premium. In 2003 the experience rating incentive had reached its maximum impact. At the same time, criticism against the experience rating has grown steadily, as an increasing group of employers had been confronted with substantial increases in their premiums. The major argument of the opponents against experience rating – both employer organizations and politicians – is that employers, in particular those with small employer size, cannot be held responsible for (on average) about 31% of the DI costs of their employers.

Empirical evidence on the effects of experience rating is limited. The most rigorous empirical analyses in this field are with respect to the US unemployment insurance (UI) system (see Meyer 2002 for a survey). Evaluations of experience rating for DI arrangements are even more scarce, and in the specific case of the Netherlands, virtually absent. In a way, this is not surprising: in the case of perfect experience rating, the (ex ante) incentive is equal for all employers – thus ruling out any variation that could potentially help to identify the impact. In practice however, experience rating systems are mostly incomplete, in the sense that premiums are bounded by minimum and/or maximum rates. This offers opportunities for the identification of incentive effects, as the (marginal) incentive will vary among employers. The empirical literature on UI experience rating largely relies upon this, as well as inter state variation in the design of experience rating.

This paper examines the effects of experience rating on the inflow into DI in the Netherlands, using a unique longitudinal administrative data set from the Dutch social benefit administration. This data set covers employers for the years 2000, 2001 and 2002. This means that about 370 thousand employers, employing roughly six million insured workers (about 75% of the working population in the Netherlands), are followed during these three years. For each employer, we observe employer specific characteristics, the worker composition, as well as information on disability cohorts, the disability risk and the experience rated premium.

Our analysis contributes to the empirical literature on experience rating in two aspects. First, we use longitudinal data, taking changes in movements of individual firms along premium schedules to be exogenous. This approach is similar to e.g. Anderson and Meyer (1994), who estimate the impact of experience rating on the incidence of unemployment insurance in the US. We follow a differences-in-differences approach to identify the impact of changes in premiums on the inflow into DI. More specifically, we compare the DI inflow of employers that have experienced premium increases with those who have not. This is in contrast to most studies that examine the impact of experience rating, and where identification follows from (inter state) cross sectional variation in the data.

Second, in our analysis we explicitly distinguish between ex ante and ex post effects of experience rating. Ex ante incentives arise if employers are aware of the incentive system and preventative actions have substantial impact. The identification of ex ante effects follows from the incompleteness of experience rating – that is, the marginal incentive to limit the inflow into disability is zero if a maximum premium is reached. In contrast to this, ex post effects of experience rating result from the unawareness of employers – that is, unanticipated premium increases may trigger employers to increase preventative activities. In particular from the perspective of small and medium-sized firms, the experience rating system is complex, and is seemingly unimportant – as long as there is no inflow into DI. This corresponds to e.g. Hyatt and Thomason (1998), who find the awareness of experience rating among individual firms to be limited. This seems particularly relevant for the first years of the experience rating system in the Netherlands.

The paper proceeds by first presenting an overview of the evidence on experience rating, as well as briefly discussing some theory and its design. Section 3 describes the Dutch DI system, with particular attention to the premium system that was effective from 1998 to 2002 – the time period under consideration in the empirical analysis. In Section 4 we discuss the data. Section 5 presents our estimation results, following estimation strategies that are based on cross sectional data first, and then taking advantage of the panel structure of the data. Section 6 concludes.

2. Experience rating: theory and evidence

Experience rating follows from the basic principle that employers may be given the discretion to their actions, provided that the corresponding costs are not imposed on other agents. Within the context of UI, this means that firms should pay for the societal costs of dismissals. As Blanchard and Tirole (2004) argue, these costs may even exceed the worker benefit level, if psychological and other costs borne by the worker are high, or in case of externalities – for example increases of crime rates in depressed areas. In practice, there are several strong reasons to abstract from the principle of full experience rating. Firms may be too fragile to handle large layoffs and the accompanying costs. This may result in extra layoffs (‘snowball effects’). A second complication is that employers may be triggered to evade taxes – employers may let their firms go bankrupt, so as to avoid further future layoff costs, and relocate new activities. Third, experience rating is likely to induce upward wage pressure by strengthening the position of workers, facing a less credible risk of becoming fired. On the whole, this results in higher unemployment.

For the inflow into DI the design of experience rating comes with more or less similar arguments.[1] This means that there is no strong case for full experience rating of DI, just like there is no strong case for any experience rating at all. In addition to this, the arguments for the degree of experience rating for DI to be stronger or weaker that UI are mixed. From the perspective of the employer, the inflow into disability is – for an important part – driven by

[1] We do not discuss all issues related to the design of experience rating in great detail. Hyatt and Thomason (1998) list a number of issues that should be addressed: the problem of small employers, the appropriate measure of disability experience, the appropriate time window and the issue of prospective versus retrospective rating.

worker specific risks, whereas layoffs result from the decisions of employers. Thus, employers cannot be held responsible for the full disability risk, and the degree of experience rating therefore should be lower for DI than for UI. On the other hand – and in contrast to the DI risk – one should be aware that the unemployment risk cannot be (fully) insured by private insurers. In this respect, the government should be more cautious with the design of incentives for UI than for DI.

In order to measure the effect of experience rating on DI, there should be substantial variation in extent of experience rating among employers. Moreover, this variation should – at least to some extent – be exogenous to the disability risk. Clearly, creating a setting in which these conditions are met is a difficult task: one has to rely on policy changes over time, or (seemingly) arbitrary institutional variation in the design of experience rating (natural experiments). The rarity of such variation explains why research in this field is scarce, and – for an important part – based on indirect inference. In this respect, it seems the empirical literature on experience rating for UI is better developed than that for DI.[2]

Hyatt and Thomason (1998) survey the empirical literature on DI experience rating for the US and Canada. They distinguish between studies describing the effects on the incidence of worker injuries, the severity of claims and the employers (preventative) activities. They conclude that experience rating is clearly associated with a reduction in the incidence of workplace

[2] Since the seminal paper of Topel (1983), a number of empirical studies have appeared on the effects of experience rating on the UI incidence, all exploiting the US inter state variation in experience rating. Meyer (2002) surveys the literature, and concludes that all studies find large effects of incomplete experience rating on lay-offs.

injuries. At the extreme, the estimated impact can even amount to about 40% (see Bruce and Atkins, 1993). Evidence of the effects on the severity of claims is however less convincing. This lends credence to the idea that, once an injury has occurred, employers have limited ability to rehabilitate workers. Finally, it seems that experience rating induces employers to appeal more often to workers' compensations claims. This suggests that part of the reduction in the incidence rate is driven by more aggressive claims management by employers, next to increased preventative activities. In this respect, Hyatt and Thomason also stress the fact that the employer awareness of experience rating is limited. In particular, it seems that employers who become subject to the demerits of experience rating are most likely to become aware of the system.

For the Netherlands, there is virtually no evidence on the impact of experience rating, either on the incidence or on the length of DI spells. Considering the literature that is somehow related to this issue, there are indications that the potential effect of experience rating is substantial. Empirical research on the size of hidden unemployment in DI suggests that the DI scheme has been used by employers as an exit route for residual workers. Hassink et al. (1997) find that about 10% of the DI inflow is due to redundancy of workers. A second strand of literature addresses the sickness behaviour of workers. Since the DI program follows after a period of absenteeism, it may well be that the effectiveness of preventative measures on absenteeism has important spillovers to DI. Following this line of research, Van Lomwel and Nelissen (2003) use a difference in difference-methodology to investigate the effectiveness of these activities. They conclude that the short term impact equals about 10% of the absenteeism rate. This confirms the idea that there is a potential for experience rating to discourage the inflow into DI.

3. Experience rating in the Netherlands

In the Netherlands, the provision of DI is mandatory and financed by pay-as-you-go contribution rates. In principle, the program covers all workers against all incomes losses that result from injuries. This, combined with the public monopoly provision of DI, makes the disability determination system rather susceptible to moral hazard problems (see e.g. Aarts en de Jong, 2000). Public monopoly insurance has a record of minimising erroneous denials, at the cost of more erroneous admissions. Moral hazard problems are further aggravated by the generosity of the DI system, which is based on the individual earnings capacity.[3] This means that disability is measured as a percentage, rather than an all or nothing condition.[4]

Over the years, the Dutch DI program has repeatedly been subject of public debate. This is not surprising, as the DI enrolment in the Netherlands has remained high and persistent. Expressed as a percentage of the insured population, DI enrolment peaked at 16% in the mid eighties, and since then declined and stabilised at about 13%.[5] At the same time, various reform plans have been introduced. For an important part, these plans have aimed at improving employer incentives. To start with, the sickness benefit program has been (fully) privatised in 1996, making employers fully responsible for these

[3] The statutory replacement of DI benefit is set at 70%. In many collective agreements, statutory DI benefits are supplemented with non-statutory benefits, in particular in the first one or two years of benefit recipiency.
[4] Buddelmeyer (2001) discusses the way the level of disability is determined in the Netherlands.
[5] See Buddelmeyer (2001) for an international comparison of DI enrolment in the Netherlands.

costs. In 1998, employer's incentives further have been enhanced by the system of DI experience rating. Finally, in 2002, the (potential) impact of incentives was further extended by a more stringent system of gate keeping, and an extension of the sickness benefit period. In order to be eligible for a medical DI assessment both workers and employers have to meet several conditions, so as to convince the benefit administration that disability was unavoidable.

Basically, the calculation of the DI premium of employers consists of two steps: the determination of the disability risk, and that of the disability premium. In both steps the exact calculation can be rather complex and experience rating may be incomplete. This does not only mean that the disability costs of the first five years of benefit years are not fully passed on, employers may also pay more than the costs involved, as we will show later on.

The calculation of the employer disability risk combines information on the disability costs of the first five worker cohorts and the (average) wage sum(s) over a five-year period. Both types of information are reported with a delay of two years. The disability risk d_t of an employer at time t is calculated as the costs of the first five DI worker cohorts, divided by the average wage sum. Both the nominator and the denominator are registered with a delay of two years:

$$(1) \qquad d_t = \frac{\sum_{u=0}^{T} S_{t-2,\, t-2-u}}{\sum_{u=0}^{T} W_{t-2-u} \,/\, (T+1)}$$

where $S_{a,b}$ are the disability costs of an employer in year a for recipients that entered into the program at b, and W_a the insured wage sum at time a. $T = 4$, as we have a five year time-window.

Averaging the wage sums over a time window of five years helps to diminish the effect of the volatility in wage sums.[6] Without averaging, e.g. a sudden drop in the wage sum in 2000 may cause the disability premium, which follows from the product of the risk in 2000 and the wage sum in 2002, to be too high. At the same time, this way of smoothing also results in incompleteness of the experience rating system – that is, fast growing firms will cross-subsidise downsizing firms. To make this clear, we derive the marginal payment that follows from one extra euro of the total DI costs for a firm, with constant growth rate g (where the growth rate is corrected for a discount rate).[7] That is, we calculate the marginal incentive e_t as the derivative of the premium rate with respect to one extra euro of DI costs, and multiply this by the wage sum at time t:

$$(2) \qquad e_t = W_t \times \frac{\partial \; d_t}{\partial \; \sum_{u=0}^{T} S_{t-2,\, t-2-u}} = \frac{(1+g)^6 \times g \times 5}{(1+g)^5 - 1}$$

[6] In some cases, the information that is needed to calculate the disability risk may be incomplete, e.g. when employers have started their businesses recently, or when for some period there are no workers at a particular firm. This means that the disability risk has to be calculated over less than five years, but rescaled to a five year period.
[7] Under the assumption that the wage sum has a constant yearly growth rate, W_t cancels out of the expression.

For g close to zero, e_t can be approximated by $1+4g$.[8] This means that cross subsidisation between firms is modest and equal to $4g$.[9]

The calculation of the DI premium is based on the disability risk, together with the average DI premium rate. Each year, the social benefit administration estimates the average premium that is needed to compensate for the expected costs for the first five DI benefit cohorts. Then, the individual premium is calculated as

$$(3) \qquad p_t \quad = \quad \min\left(p_{min} + d_t \ , \ p_{max} \right)$$

where p_t is the premium rate, and p_{min} and p_{max} the minimum and maximum premiums – indicating the range over which p_t may vary.

Note that the minimum premium is in fact a uniform premium that is paid by all employers, also for $d_t = 0$. Furthermore, values for both p_{min} and p_{max} are different for small firms on the one hand, and medium sized and large firms on the other hand. In particular, if the wage sum in t exceeds 15 times the average wage sum across all employers, then p_{min} is set lower, and p_{max} higher, thus extending the support of premiums. For large firms, the maximum premium is set equal at four times the average premium, whereas that for small firms is three times the average premium. Next, the minimum premiums are set

[8] For small growth rates, the wage sum growth can be approximated by a linear function. Thus, the average wage sum then equals the third year of the five year time window. This, together with the two year delay, implies that the total delay equals four years. Thus, the disability risk is overestimated with 4 x g % .

[9] For larger growth rates however, the impact of smoothing can be substantial. For example, suppose a yearly growth rate of the wage sum of a firm of 25%. Then, the resulting marginal incentive equals 2.3, implying that about 232% of the marginal costs are passed on to the employer. Similarly, if we would assume the yearly growth rate to be negative, say – 20%, then the marginal incentive would equal 0.39.

at the level that balances the disability costs with the collected premiums. This has to be done iteratively, as, due to increases in the minimum, more premiums will be bounded by the maximum.[10] Since premiums of small firms have a higher probability to be bounded by the maximum, the minimum premium is higher for small firms.

In the first five years (1998 to 2003) the experience rating system has gradually been brought into force. This means that the average premium first was based on the DI costs of new DI cohorts in 1998, and extended with an additional cohort in the subsequent years. Accordingly, the maximum and minimum premium rates have increased over time. However – and in contrast to this – as early as the introduction of experience rating in 1998, the calculation of the disability risk has been based on a five year time window. Thus, the individual disability risks had to be adjusted to match them with the premium rates.[11]

In sum, the experience rating system gives a clear incentive for employers to decrease their DI costs, as it covers an important part of those DI costs.[12] Still, this does not mean that incentives have been strong for all employers at all time periods. First, due to the gradual introduction of the system, incentives were only modest in the first years. Second, the way the disability risk is calculated accommodates cross subsidies from (fast) growing to downsizing firms. And third, similar cross subsidies (still) exist from high risk to low risk employers.

[10] Apart from that, minimum premium revenues are also needed to finance the disability costs of firms that have gone bankrupt, or to make up past reserve deficits

[11] For example, for 1998 the adjustment rate was 0.20, and in 2001 it equaled 0.83.

[12] In 1998, the costs of the first five years of the DI program were estimated at 37%.

4. Data

The data set we use combines yearly information of all employers in three administration systems of UWV, the Dutch social benefit administration for 2000-2002. First, we have all registered information that is needed for the calculation of disability risks and premium rates. Second, we have (limited) worker information (like age and gender) that can be aggregated to the level of employers. Here, it should be noted that – in the current setup of the data – individual workers are not followed over time. Thus, in the estimation we cannot allow for individual worker specific effects. Third, data are derived from the social security records, containing information on the DI benefit recipients for various cohorts. This also includes the inflow into the DI program.

We have merged these three administration systems using employer codes, resulting in a panel data set of about 370 thousand employers, employing roughly six million workers. Table 1 summarises the main characteristics of this data set. To start with, it is important to note that the data set does not cover all employers and workers in the Netherlands. In particular, experience rated premium rates cannot be calculated for firms that exist for less than one year, or firms that have merged or split up recently. This induces some variation in the number of employers that is observed in the data over time, as well as the number of employees.

< Table 1 here >

Table 1 shows that the impact of experience rating has increased in the time period under consideration. This increase is most prominent in terms of the number of benefit recipients that has been assigned to employers (from 1.5 to 2.6% of the working population). In terms of the DI risk rates, the increase is less pronounced (from 1.2% tot 1.4%). This is not surprising, as the risk measures are rescaled to a five year window if information is incomplete. Finally, we see that the premium rates have increased gradually, from 1.4% to 1.6% of the wage sum.

The inflow rate into DI can be derived from the inflow variable, together with the employer size.[13] This rate equals about 1.0% of the working population.[14] Moreover, the inflow consists for about 50% of female workers, indicating that the individual inflow rate is somewhat higher for women (as they make up about 44% of the working population). Finally, in the time period under consideration, the inflow rate into full disability has increased slightly, whereas the opposite holds for the inflow into partial disability.

Some of the DI inflow is measured with a delay – causing the observed inflow into DI to be susceptive to measurement errors. In particular, when the DI spell of new recipients starts at the end of 2000, the assignment to a specific employer can take place in 2001. This measurement problem is limited to small employers. Therefore, in the empirical analysis, we select employers with at least ten employees. This reduces our sample size substantially – about 80% of

[13] Additionally, the DI outflow rate can also be derived from the cohorts of benefit recipients, but this is accompanied by measurement errors – in particular in the first year of benefit recipiency. Moreover, we do not have information on the exact destination following the DI spell.

[14] This rate is averaged over employers, not employees. As the inflow rate is somewhat lower for smaller firms than for medium sized and large firms, the resulting average is lower than the employee average.

the employer population is left out – but, in terms of the number of workers, the loss is only about 20%. In order to construct a panel that is balanced, we also select employers that are observed for all relevant years. As a result, the proportion of sectors with small employers – such as the trade sector and the primary sector – decreases, as well as sectors where firm turnover is high.

< Table 2 here >

Table 2 provides more detailed summary statistics on the dynamics of the experience rating system. Here, we distinguish between three employer types: (i) employers with small employer size, for which the maximum premium equals three times the average premium; (2) employers with no more than 100 employees, for which the maximum premium equals four times the average premium; and (3) employers having similar experience rating parameters, employing 100 workers or more. The vast majority of small employers in the sample has no DI costs that can be assigned to them, paying the minimum premium. Over time, we see that this group has decreased only slightly, indicating that the risk of premium increases is limited. At the same time, the number of employers paying the maximum premium has increased to the same extent. Obviously, this pattern would be far more prominent for the full sample, also including employers with less than 10 workers. For this group the DI costs of a benefit recipient constitute an important part of the wage costs – thus making it more likely that the maximum premium will be reached. For medium sized employers the picture becomes more dispersed – i.e. there is a higher probability of paying a premium between the minimum and maximum rate. Comparing the premium rates for 2000 and 2001, we also see that the probability of moving along the premium schedule is larger than for small employers. Still, we still find more than half of the employers paying the minimum premium. The likelihood of paying the minimum premium – just as the maximum – is very small for large employers.

In equation (2) we have calculated the experience rating incentive measure assuming a constant growth rate for the wage sum. As a result, the marginal incentive of an extra euro of DI costs was constant over time and over the various benefit cohorts. Given the data at hand, this equation now can be generalised by allowing for the growth rate to vary over time – thus inducing the marginal incentive for the specific cohort years to be time dependent. This means that the marginal incentive can be calculated as a (discounted) weighted average over a time period of five consecutive years of DI costs:

$$(3) \qquad e_t = \frac{\sum_{c=0}^{T} \psi^c \, e_{t,c}}{\sum_{c=0}^{T} \psi^c}$$

with T = 4 and $e_{t,c}$ representing the expected marginal incentive of DI costs in the c+1[th] year of benefit recipiency. Discounting occurs at rate ψ. The marginal incentive for the c+1[th] year of benefit recipiency is calculated as the derivative of the (future) disability risk d_{t+c} with respect to the future disability costs in the c+1[th] year, multiplied by the wage sum in the c+1[th] year of benefit recipiency:

$$(4) \qquad e_{t,c} = W_{t+2+c} \frac{\partial \; d_{t+c}}{\partial \; S_{t-2,\, t-2}} = \frac{W_{t+2+c}}{\sum_{u=0}^{T} W_{t+c-u} \,/\, (T+1)}$$

with $c = 0,..\ T$.

Note that for a constant growth rate of the wage sum, we obtain the same expression as in equation (2), and the marginal incentive for the separate years of benefit recipiency are all equal. Furthermore, equation (4) makes apparent that, in order to determine e_t , we need information on the past and future values of W.[15] Therefore, we have calculated sector specific (average) growth rates to extrapolate expected future values of W. Further, we set ψ equal to 85%, corresponding to a yearly discount rate of 15%.[16]

< Table 3 here >

Table 3 provides summary statistics on the resulting marginal incentive measures, averaged over sectors. The first column of the table – expressing the percentage of employers with $p < p_{max}$ – shows that only a small fraction of the sample consists of employers for whom the marginal incentive is equal to zero. Not surprisingly, this is most likely to occur in sectors with small employer size, such as the primary sector and service industries. The third column of the

[15] Again, note that, due to the delay in the administrative system, in the sample wage sums are observed up to 2000. In order to obtain observations for 2001 and onwards, we extrapolated the wage sums for subsequent years by using the average sector specific wage sum growth rates for 1998-2000. These average growth rates were calculated on the basis of two-digit-sectoral information (at about 70 sectors in total).

[16] Within the context of the experience rating system, the discount rate consists of the real interest rate, together with the exit rate out of DI. The yearly real interest rate is set equal to 5%, and the average exit rate is set equal to 10%.

table represents the sectoral (average) marginal incentive, multiplied by the probability of $p < p_{max}$. Here we find employers in sectors with (on average) high wage sum growth rates – such as the service and the financial sector – to pay more than 100% of their (additional) DI costs – thus cross-subsidising e.g. employers in the primary sector.

5. Empirical strategy and estimation results

5.1. Identification issues

In the empirical literature, the identification of incentive effects of experience rating is typically built on three major assumptions: (i) exogeneity in marginal incentive variation; (ii) local linearity of the incentive system; and (iii) rationality of employers.

First, and as argued in the previous sections, it is assumed that variation in the marginal experience incentive is – at least to some extent – exogenous. Exogeneity may arise in the case of (inter state) variation in the experience rating systems – such as the UI programs in the US, where the degree of completeness of experience rating varies between states. For the Netherlands, such an approach is not applicable, since the only institutional variation comes from the distinction between small employers and medium and large employers. Exogeneity can also be derived from variation in the experience rating incentive over time, using longitudinal employer data. Identification then follows from movements along tax or premium schedules. Anderson and Meyer

(1994) use this approach, so as to provide first difference estimates of the impact of UI experience rating.

Second, the local linearity assumption is needed to ensure that employer decisions are driven by the current marginal incentive. Stated differently, the incentive the employer is faced with is well approximated by the slope that determines how the employer's costs would change in case of a small one-time increase in DI costs. For the US system of the UI program, the local linearity assumption does not seem too restrictive. In most states benefit exhaustion occurs after 26 weeks of unemployment, whereas the marginal incentive changes on a yearly basis. However, for programs with a longer time horizon – such as the Dutch DI system – the relevant marginal incentive may change over time, thereby invalidating the local linearity assumption.

Third, empirical evidence on the effects of experience rating is mostly based on variation in ex ante incentives. The presence of such effects – in particular for the UI system – suggests that employers are fully aware of the experience rating system and its incentives. Thus, the (implicit) assumption is that employers are driven by ex ante incentives, not ex post. Again, for the Dutch DI system, it can be argued that this assumption is too restrictive. In the time span covered by our analysis the system was introduced only recently, the vast majority of employers having no DI costs and paying the uniform (minimum) premium rate. Therefore, it may well be that the ex ante incentive – in particular for small and medium sized employers – was limited.

5.2. Empirical strategy and implementation

Estimating the effect of experience rating using standard techniques is likely to result in severe estimation biases. In order to show this, let us start by specifying the probability that worker *i* of employer *j* enters into the DI program at time *t* , using the familiar Logit specification:

$$(5a) \quad DI_{ijt} \quad = \quad \Pr (DI_{ijt} = 1) + \varepsilon_{ijt}$$

with

$$(5b) \quad \Pr (DI_{ijt} = 1) \quad = \quad \frac{\exp [\boldsymbol{X}_{jt} \boldsymbol{\beta} + \gamma \, \mathrm{I} (p_{jt} < p_{max, jt}) + \alpha_j]}{1 + \exp [\boldsymbol{X}_{jt} \boldsymbol{\beta} + \gamma \, \mathrm{I} (p_{jt} < p_{max, jt}) + \alpha_j]}$$

where $DI_{ijt} = 1$ if the individual enters into the DI program, and 0 otherwise. Further, $\boldsymbol{X}$ denotes employer specific characteristics and I is an indicator function denoting the event between parentheses. $\boldsymbol{\beta}$ is a vector of parameters describing the effect of $\boldsymbol{X}$, and γ denotes the impact of the (ex ante) experience rating incentive. α_j is a fixed component of the error term that captures permanent firm characteristics that affect the DI inflow probability. Finally, ε_{ijt} denotes the error component.

As $\mathrm{I} (p_{jt} < p_{max, jt})$ is a function of past DI inflow (with a delay of two years), it can easily be shown that the marginal incentive of experience rating is correlated with α_j , thus resulting in biased estimates of γ. Given the data at hand, there are two main empirical strategies to correct for this. The first is to exploit the (cross sectional) variation in the experience rating schedule for different employer types. The assumption that this variation is exogenous is too restrictive, but there is valuable information on past DI costs that can be used as control variables. In the second estimation approach, we exploit the panel character of the employer data, taking changes in the premium schedule to be exogenous. If ex ante effects are important, we would expect the inflow into DI to increase if an employer jumps up to the maximum premium, having no (marginal) incentive. Ex post effects are particularly relevant for employers without DI costs, paying the minimum, uniform premium, and are subsequently confronted with (partially) unanticipated raises in the premium.

5.3. Cross sectional estimation results

The first strategy to estimate the impact of experience rating is to include control variables for the worker specific effect, leaving us with variation in the incentive that is due to institutional variation. An obvious candidate for this is the observed DI risk of employers d_{jt} , which is needed to calculate the premium rate. We may expect the DI inflow probability to be an increasing function in d_{jt} , as this variable is an overall summary of the past DI inflow probabilities – hence approximating the employer specific effects. Due to the institutional setup of the experience rating system, we would expect the inflow probability to have an upward jump if the risk equals $p_{max}– p_{min}$. Thus, equation (5b) can be rewritten as:

$$(6) \qquad \Pr(DI_{ijt} = 1) = \frac{\exp[\boldsymbol{X}_{jt}\boldsymbol{\beta} + \gamma\, \mathrm{I}(p_{jt} < p_{max,jt}) + f(d_{jt})]}{1 + \exp[\boldsymbol{X}_{jt}\boldsymbol{\beta} + \gamma\, \mathrm{I}(p_{jt} < p_{max,jt}) + f(d_{jt})]}$$

where f denotes a spline function of the DI risk percentage of an employer. Further, the assumption that is needed for identification – d_{jt} is an appropriate control for α_j – can be formulated as:

$$(7) \qquad \mathrm{E}(\mathrm{I}(p_{jt} < p_{max,jt}) \mid \alpha_j, d_{jt}) \quad = \quad \mathrm{E}(p_{jt} < p_{max,jt}) \mid d_{jt}).$$

Now, if f is sufficiently flexible, we have two sources of identifying variation for the incentive effect. First, under the assumption that the local linearity assumption holds, the effect is identified by the discrete jump at the point of support where $d_t = p_{max} - p_{min}$. Second, premium schedules are different for small and medium/large employers. For some part of the support of the disability risks, the incentive is limited to medium and large employers, with small employers paying the maximum premium. This range is defined as:

$$(8) \qquad \underline{p}_{max} - \underline{p}_{min} < d_{jt} < p_{max} - p_{min}$$

with $\underline{p}_{min}$ and $\underline{p}_{max}$ denoting the minimum and maximum premium for employers with small employer size, and p_{min} and p_{max} for minimum and maximum premium for employers with medium and large employer size.[17]

[17] From 2000–2002, the support increased from 1.96 to 2.80%. For example, in 2002 $\underline{p}_{min}$ and $\underline{p}_{max}$ were equal to 1.24% and 6.06%, and p_{min} and p_{max} were set at 0.45% and 8.08%. Thus, the

< Table 4 here >

Table 4 presents the Maximum Likelihood estimation results of the Grouped Logit model[18] for equation (6) for 2000, 2001 and 2002. When looking at the sectoral dummies, we find the DI inflow to be highest in the primary sector (i.e. the reference sector) and for temporary employment agencies, whereas the inflow is lowest in the financial sector. For the other sectors, the relative position seems to change over the years. For the wage sum variable, we also find coefficients to change over the years. In particular, for 2000 the DI inflow probability increases with respect to the average wage sum. This can be explained by the relative attractiveness of the DI program for workers with high wages – DI benefits are based on the last earned wage – as well as the fact that part time workers have a lower DI risk. This picture fades out when looking at 2001 and 2002, which may be due to the setup of the data: as we have a panel of employers that are observed for three consecutive years, the number of employers with a low wage sum and small employer size decreases considerably. Employers performing poorly – employing less workers and paying lower wages over time – sort into the reference group. This group of employers is more inclined to use the DI program as a substitute for

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premium range with different marginal incentives for employers with small employer size and those with medium and large employer size was 4.82% (6.06–1.24) < .. < 7.63% (8.08–0.45).

[18] STATA is used for the estimation of the model. Although the DI risk is expressed in terms of individual workers, all explanatory variables of $\boldsymbol{X}$ are measured at the level of employers – otherwise the data could not be grouped according to this level. As worker data are grouped according to employers (with employer size N), the model is estimated by Grouped Logit estimation. In the pooled estimation (section IV-D) we use the cluster option in STATA as well, so as to combine yearly observations of employers and obtain robust standard errors.

the UI program, explaining the changing pattern in the coefficient estimates describing the effect of employer size.

Age and gender seem to be most important in explaining the inflow into DI. On average, the DI probability is somewhat higher for female workers. Remarkably, the inflow probability seems to be hump-shaped function of age. This suggests that sorting effects are important here – that is, unhealthy workers becoming non-participant over time.

In the estimation of the model, we have extended the number of polynomials in the spline function to a maximum of eight.[19] The inclusion of this spline function affects our estimate of the incentive coefficient substantially – indicating the size of the estimation bias that would occur otherwise. In particular, without the spline function we find the coefficient estimate to equal –1.03 (0.023) for 2002, whereas including polynomials results in a coefficient estimate of – 0.20 (0.031). Thus, including *f* helps us to correct for substantial estimation biases – concentrating on the employers with a premium rate that lies in the support that is defined by equation (8). In this respect, note that the control group in this range – existing of employers with small employer size – is very limited. In 2000, this group contains only about 3,000 individuals working for 175 employers, as opposed to a treatment group – employers with medium and large employer size – of about 23,0000 individuals (working for 2,500 employers). This is also reflected in the relatively large size of the standard error of the incentive coefficient for 2000. We only find the impact of experience rating to be significant for 2002. This suggests that either the

[19] Higher order terms did not result in a substantial improvement of the fit of the model.

experience rating has only led to a substantial reduction in the inflow into DI in 2002, or that the incentive effect has gradually increased over time.

< Figure 1 here >

Figure 1 illustrates the estimated impact of the (ex ante) experience rating incentive in 2002, averaged over all workers in the sample. The dotted line shows the premium schedule for small employers, as opposed to the premium schedule of employers with medium and large employer size. For 4.82% < *d* < 7.63% these schedules do not intersect. Here, the incentive for employers with small employer size is equal to zero, resulting in an inflow probability effect of 22%. This corresponds to an increase in the DI inflow probability of 0.2%-point.

Clearly, our estimation strategy helps us in correcting for substantial endogeneity biases. However, when interpreting our results we have to be aware of two (implicit) assumptions that have been made so far. First, our estimation strategy strongly relies on the local linearity assumption. We have argued earlier that this assumption is probably too restrictive in the context of the DI program in the Netherlands. Within the relevant time horizon of five years, employers are likely to move along the premium schedule. This means the marginal incentive will only gradually decrease with respect to the disability risk percentage, rather than being an all-or-nothing condition. Thus, we may expect possible incentive effects to be picked up by the spline function. The second (implicit) assumption that underlies our relates to the specification of the incentive effect. We assume this effect to be equal for all employers, irrespective of the number of employees. Thus the difference in the DI probability in the support of equation (6) can be fully attributed to the experience rating incentive. This assumption may be too restrictive if there are economies of scale in preventative activities.

If both assumptions are doubtful, the question arises how we then should interpret our results. Suppose the incentive decreases only gradually, and is virtually complete for all types of employers with $p < p_{max}$ – also including employers with small employer size. Then, if the incentive effect is equal for all types of employers, we would expect the coefficient estimate of γ to be equal to zero. This means that γ is informative on the *relative effectiveness* of the experience rating incentive of employers with large employer size, compared to those with small employer size. The coefficient estimate of – 0.20 for 2002 thus implies that the incentive effect is higher for employers with medium and large employer size.

5.4. Panel model estimation results

The advantage of the first estimation strategy is that only (yearly) cross sectional data are needed to estimate the effect of experience rating. However, identification of the model strongly relies on the local linearity assumption, and – despite the size of the data set – only a few observations can actually be used to identify the incentive effect. The second estimation strategy – exploiting the panel character of the data – is less susceptible to these problems.

Within the context of panel data models, a number of techniques can be used to account for individual (employer) specific effects. Anderson and Meyer (1994) use a linear probability model, because of their large sample size and the large number of explanatory variables they use – enabling them to obtain first difference estimates of the effect of experience rating. For our model such an approach is less appealing, as the DI risk is only small and the number of time variant explanatory variables in our data is limited.[20] As an alternative, following Mundlak (1978), we therefore pool the data from 2000–2002 and include the average values of $\boldsymbol{X}$ as explanatory variables in our model. The intuition behind this approach is that – by including averages – the effect of changing $\boldsymbol{X}$ is estimated, holding the time average fixed. For the parameter describing the impact of experience rating, we allow this effect to be asymmetric.[21] This means that we rewrite equation (5b) into:

$$(9)\quad \Pr(DI_{jt} = 1) = \frac{\exp[\boldsymbol{X}_{jt}\boldsymbol{\beta} + \alpha_{uv} + \Delta_{uv}\,\mathrm{I}\,(t = 2001) + \Delta'_{uv}\,\mathrm{I}\,(t = 2002)]}{1 + \exp[\boldsymbol{X}_{jt}\boldsymbol{\beta} + \alpha_{uv} + \Delta_{uv}\,\mathrm{I}\,(t = 2001) + \Delta'_{uv}\,\mathrm{I}\,(t = 2002)]}$$

with

[20] We also have estimated the DI inflow using a linear probability model. For the cross sectional data, this resulted in similar estimation results as for the Logit (and Probit) model, but the results differed substantially when applying first difference estimation on the combined (panel) data.

[21] We also could have used Random Effects Probit or Logit model. However, given the size of our data, the estimation of such models is too cumbersome and time consuming. As is shown by Mundlak (1978), the estimated parameters of pooled estimation give very similar results, provided that the average values of time varying parameters are included as explanatory variables and that a robust variance matrix is used to account for within-employer variation.

(10a) $u_j = I(p_{j,\, t=2000} > p_{min,\, j,\, t=2000})$

and

(10b) $v_j = I(p_{j,\, t=2001} > p_{min,\, j,\, t=2001})$

u_j and v_j are indicator dummies that are equal to one if employer j pays a premium exceeding the minimum premium in 2000 and 2001, respectively. The combination of u_j and v_j results in four employer categories for which α , Δ and Δ' are allowed to vary. As a reference group, we take employers that have paid the minimum premium in both years, thus having $\{u_j = 0 , v_j = 0\}$. Obviously, we would expect the DI inflow probability to be higher for employers for which $u_j = 1$ and/or $v_j = 1$, as employers with high α are more likely to have disability costs. Thus, α_{uv} can be interpreted as the mean employer specific effect for employers. The assumption here is that including α_{uv} eliminates the correlation between the employer specific effect and the incentive measure. In this respect, the parameter values of Δ and Δ' that denote the change in the DI inflow probability for 2001 and 2002, can be interpreted as difference-in-difference estimates of the (ex post) effect of experience rating.

< Table 5 here >

Table 5 presents the baseline values of the DI inflow probability of the four employer groups for 2000-2002, where the baseline is normalised to zero for $\{u = 0, v = 0\}$ in 2000. We take $\{u = 0, v = 0\}$ as reference group for all years, with $\Delta_{2001\text{-}2000}$ and $\Delta_{2002\text{-}2000}$ as the respective time trend dummies. Difference-in-difference estimation follows from the comparison of Δ and Δ' between the employer groups. Under the null hypothesis that there is no impact of ex post incentives of the experience rating system, we would expect Δ_{01} and Δ_{01}' to be equal to zero.

In our estimation strategy we take advantage of the fact that past inflow into DI only affects the premium rate with a lag of two years. The intuition behind this is as follows. Suppose that an employer has no relevant DI recipients until 1998, but in 1999 one or more workers in fact do enter into the DI scheme. Due to the two-year time delay, the premium rate increase occurs in 2001. The ex post effect then follows from the comparison of the DI inflow between 2001 and 2000, as well as 2002 and 2000. Obviously, this is likely to result in a substantial estimation biases if, for example, the time delay of the experience rating system would be equal to one year.

< Table 6 here >

Table 6 shows the estimated parameter coefficients that result from the difference-in-difference strategy, using equation (9). To start with, it should be noted that employer specific effects indeed are important. Due to the inclusion of average values of the worker composition variables, the parameter estimates describing the impact of age and gender change substantially. In particular, we find the DI probability to increase with age for almost all age categories –

whereas the pattern was hump-shaped according to our earlier estimates. This suggests that older workers sort into jobs with low DI risks. The importance of unobserved employer specific effects is also mirrored by the coefficient estimates of α. Employers with high DI risks will sort into α_{01} , α_{11} and α_{10}. Not surprisingly, we find the estimate of α_{11} to be the highest, whereas the DI inflow is lowest for the reference group of employers paying the minimum premium in 2000 and 2001.

When looking at the difference-in-difference estimates, we see that the time trends differ substantially between employer categories. To start with, we find the inflow into DI to have increased in 2001 with about 17%, and stabilised in 2002. For the other groups, this upward trend is less prominent, or even in the opposite direction. In particular, the increase in the DI inflow rate of employers that experience a rise in their premium rate in 2001 is about 4% lower than the reference group in 2001, and about 15% lower in 2002 (with a coefficient of – 0.16 (0.021) [22]). This suggests that ex post incentives of the experience rating system are important. Apparently, premium rate increases are unanticipated by employers – making them (more) aware of the experience rate incentives, and triggering them to increase preventative activities. As such activities take time to become effective, it is not surprising that the incentive effect is found to be increasing over time.

Remarkably, we find the difference estimates for employers with DI costs in 2000 even to exceed that of employers no longer paying the uniform minimum premium in 2001. The estimated coefficients for these groups are

[22] This implies that, comparing 2000 and 2002, the DI inflow probability has remained constant for this group.

equal to – 0.25 (0.017) and – 0.26 (0.032). However, when interpreting this result, one should be aware that in the time period under consideration, the experience rating system was introduced only recently. If the experience rating system has triggered employers to increase preventative activities, it is likely that these have become effective over time, explaining the decrease in the DI probability in 2000–2002. Still, caution is needed for such an interpretation, as there may be other factors explaining the decrease of the DI risk for these particular groups.

Table 6 also presents estimation results for the separate inflow into partial and full DI schemes. The picture that emerges is that the relative importance of both schemes differs substantially across sectors and employer and worker types. Most strikingly, the inflow probability into partial DI schemes increases with respect to the average wage sum of an employer, whereas the opposite holds for the inflow into full disability. This can be explained by the fact that the degree of disability is based on the residual earnings capacity of workers, compared to the minimum wage. As low wage workers are more likely to have an earnings capacity that is lower than the minimum wage, they are more likely to enter into the full DI scheme.

Similar as for the total inflow into the DI schemes, we find substantial ex post effects of experience rating – measured by the difference-in-difference estimators Δ_{01} and Δ'_{01} . For both schemes we find a similar effect of experience rating, at about 15% after one year. This suggests that either the inflow into full disability are equally susceptible to moral hazard problems, and/or that part of the former inflow has been shifted to the partial DI scheme – i.e. that both schemes are substitute pathways. In both cases, restricting the experience rating to the partial DI scheme is not a wise thing to do.

5.5. Robustness tests and model extensions

So far, the outcomes of the difference-in-difference approach have been subject to a number of specifying assumptions. In particular, we have adopted a Logit specification and assumed the inflow into DI to be serially uncorrelated. We have re-estimated the model using various alternative estimation methods, so as to test for the potentially biasing impact of these assumptions. To start with, an obvious robust test was to re-estimate the model using a Probit specification. This resulted in very similar average derivatives of the parameter estimates. We also have addressed the potential bias impact through serial correlation. To clarify this issue, let us look at two moment restrictions that are needed for the identification of the immediate experience rating effect in 2001 and the one-year-effect in 2002, respectively:

(11a) $$E[\ v_j\,(\varepsilon_{ij,\,t+1} - \varepsilon_{ijt})\ ] = 0$$

and

(11b) $$E[\ v_j\,(\varepsilon_{ij,\,t+2} - \varepsilon_{ijt})\ ] = 0$$

These moment conditions state that the probability of a premium increase in 2001is uncorrelated with changes in the error term. We know that v_j

is a function of the past inflow into the DI program. Due to the delay of two years, and under the assumption that there is no serial correlation, there is no estimation bias. However, if there is serial correlation between ε_{ijt} and $\varepsilon_{ij,\ t-1}$, then, using a first order Taylor approximation of v_j with respect to $\varepsilon_{ij,t-1}$ it can be shown that the estimated effect will be biased:

$$\text{(12a)} \quad E\left[\ v_j\left(\varepsilon_{ijt+1} - \varepsilon_{ijt}\right)\right] = E\ \frac{\partial v_j}{\partial \varepsilon_{ij,t-1}}\ \sigma_\varepsilon^2\ \left(\rho^2 - \rho\right)$$

and

$$\text{(12b)} \quad E\left[\ v_j\left(\varepsilon_{ij,t+2} - \varepsilon_{ijt}\right)\right] = E\ \frac{\partial v_j}{\partial \varepsilon_{ij,t-1}}\ \sigma_\varepsilon^2\ \left(\rho^3 - \rho\right)$$

where ρ is the degree of serial correlation of ε_{ijt} .

To test for the presence of serial correlation, we have calculated the degree of (serial) correlation between the first differences of the estimation residuals of equation (9). This resulted in a coefficient estimate of 0.065. What this suggests is that (positive) serial correlation declines over time considerably, and the expected (negative) bias is virtually equal for the effect in 2001 and 2002 (since $\rho^3 \approx \rho^2$). We do not find such a pattern to be supported by the estimation results. In particular, the immediate effect of experience rating is insignificant, whereas the effect is substantial after one year. From this we conclude that the biasing impact serial correlation is not substantial.

< Table 7 here >

Our results indicate that ex post incentive effects of experience rating are important – i.e. employers are triggered to increase preventative activities, once they have experienced (unanticipated) increases in their premiums. It is interesting to see whether this result varies across employer types and alternative incentive measures. First, one may argue that the impact of experience rating is larger for employers with high wage sum growth rates, paying more than 100% of additional DI costs. To test for this, we have replaced the incentive indicators Δ_{01} and Δ'_{01} by the marginal incentive measures, where account is taken of the employer-specific wage sums. If employers respond to this more accurate incentive measure, rather than the mere fact of paying a premium that exceeds the minimum, this is likely to improve the fit of the model. The estimation results that follow from this approach – which are presented in the first column of Table 7 – indicate that this is not the case.[23] Therefore we conclude that the decision to increase preventative activities is mainly an issue of being aware of the experience rating system, rather than responding in line with more accurate measures of financial incentives.

Second, we have extended the model by allowing for difference-in-difference estimates for employers that have paid the maximum premium in

[23] We also re-estimated the model with both the indicator dummy and the more precise marginal incentive measure. The coefficient estimate for the two-year impact of the dummy indicator was equal to – 0.16 (0.021), whereas the impact of the other measure was not found to be significant.

2000-2001. If ex ante incentives are important, we expect employers that start paying the maximum premium in 2001 – having no (marginal) incentive to prevent further DI inflow – to have a higher DI inflow in 2002. In terms of the parameter coefficients in the second column of Table 7, this means that we expect $\Delta_{11} < \Delta_{12}$ and $\Delta'_{11} < \Delta'_{12}$. In a similar vein, we expect the reverse to hold for employers that no longer pay the maximum premium in 2001: $\Delta_{22} > \Delta_{21}$ and $\Delta'_{22} > \Delta'_{21}$. When looking at the parameter estimates, we find all four hypotheses not to be supported by the data. This suggests that incompleteness of the experience rating system due to the maximum premium does not affect the behaviour of employers. Moreover, and in contrast to this, in some cases there is weak evidence of the opposite effect to hold. For example, employers that start paying the maximum premium in 2001 are found to have a parameter estimate of – 0.35 (0.043) for 2002, compared to an estimate of – 0.26 (0.017) of the (control) group that continued paying a premium rate below the maximum.

Finally, the third column of Table 7 shows that the estimated impact of experience rating is only significant for employers with more than 25 workers – in particular with a delay of two years.[24] Apparently, economies of scale are important here, lending credence to the idea that the degree of experience rating should be lower for small employers. This also confirms our earlier finding that resulted from the first estimation strategy – using cross sectional data only.

[24] We also re-estimated the model, allowing the effect of experience rating to differ between sectors. For most sectors, the estimated effect corresponded to the average impact of about 16% (after two years). We only found the impact to be relatively low in the trade sector and the social services sector.

6. Conclusions

The overall picture that emerges from our empirical analysis is that the impact of experience rating on DI inflow is substantial. This is in accordance with findings in other international studies. More strikingly, we find the decision of employers to increase preventative activities to be mainly an issue of being aware of the experience rating incentive. In the time period under consideration, it seems that the awareness of employers of the experience rating system was limited – providing us with the key ingredient of our difference-in-difference approach. Employers have been triggered to increase their preventative activities ('ex post incentives'), once they have experienced (substantial) increases in their premium rates. Apparently, employers were not well informed, either on the experience rating system or on the nature and the size of their disability risk.

We have estimated the ex post effect of the experience rating system to amount to a 15% reduction in the DI inflow after one year. This finding is robust with respect to various specification alternatives. In particular, we have found the potential bias through serial correlation in DI risks to be unimportant. Other robustness tests confirm the idea that preventative activities are driven by ex post incentives, rather than responses that are in line with more accurate measures of financial incentives (ex ante incentives).

When taking a policy perspective, there are a number of lessons that follow from our analysis. First, an obvious option is to increase the awareness of employers of the experience rating system. Similar to the analysis of Hyatt and Thomason (1998), employers seem to be less aware of the merits of experience rating – i.e. a gradual reduction in the premiums they pay, as long as there is no inflow into DI – than the disadvantages, once the premium increases occur. To increase awareness, one may think of the provision of internet services and software options to calculate the financial consequences of DI costs. Second, moral hazard problems are equally important for the inflow into partial and full disability. Therefore, restricting the experience rating incentive to partial DI schemes only does not seem a wise thing to do. Third, lowering the incentive for small employers is a more obvious policy option, since we found experience rating incentive no substantial impact of the incentive for employers with less than 25 workers. This suggests that economies of scale in preventative activities are important.

All in all, our results point out the effectiveness of the DI experience rating system. Still, for a more comprehensive evaluation, we need to know what the effects are on the outflow of the DI scheme, as well as the consequences for the hiring behaviour of employers. Until now, the literature has largely ignored these issues, so these are important avenues for future research.

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Table 1 *Employer characteristics: total and selected sample*

	Full sample			Selected sample		
	2000	2001	2002	2000	2001	2002
# Employers	309174	315314	312656	66063	66063	66063
Number of employees	6524458	6972086	6922609	5181148	5470448	5462871
Average employer size	21.1	22.1	22.1	78.4	82.8	82.7
Sectors (%)						
Primary sector	6.2	5.9	5.9	3.9	3.9	3.9
Industrial sector	21.3	21.3	21.4	28.6	28.6	28.6
Trade sector	28.1	27.6	26.8	21.0	21.0	21.0
Service industries	3.3	3.4	3.5	5.1	5.1	5.1
Transport	3.3	3.3	3.3	4.8	4.8	4.8
Catering	8.2	8.3	8.3	7.6	7.6	7.6
Social services / cultural	12.7	12.5	12.5	10.3	10.3	10.3
(Semi-)public	1.6	1.6	1.5	5.4	5.4	5.4
Financial sector	13.9	14.7	15.3	11.3	11.3	11.3
Temp. empl. agencies	0.8	0.7	0.7	1.1	1.1	1.1
Unknown	0.7	0.8	0.8	0.7	0.7	0.7
Age and gender (%)						
15-25	23.6	23.3	23.2	22.5	21.9	21.4
26-35	32.3	31.5	30.3	31.6	30.5	29.0
36-45	22.0	22.5	23.1	23.2	23.7	24.3
46-55	16.6	16.7	17.1	17.4	18.0	18.7
56-65	5.6	6.0	6.5	5.4	6.0	6.6
male	56.1	56.3	56.1	62.4	62.1	61.7
female	43.9	43.7	43.9	37.7	37.9	38.3
Experience rating variables						
Average wage sum	34983	–	–	41319	–	–
Disability risk (%)	1.18	1.37	1.35	1.29	1.48	1.69
DI premium (%)	1.39	1.51	1.56	1.52	1.57	2.01
$p_{min} - p_{max}$, small empl.	1.24 – 4.17	0.98 – 4.77	1.24 – 6.06	1.24 – 4.17	0.98 – 4.77	1.24 – 6.06
$p_{min} - p_{max}$, large empl.	0.67 – 5.56	0.41 – 6.36	0.45 – 8.08	0.67 – 5.56	0.41 – 6.36	0.45 – 8.08
% p_{min}	86.2	84.5	83.2	55.6	51.7	47.7
% p_{max}	5.1	4.9	4.5	2.8	5.4	8.3
% between p_{min} and p_{max}	8.7	10.7	12.2	41.5	43.0	44.0
Disabled workers as % of employer size						
Total	1.52	2.31	2.63	1.74	2.94	4.07
- Male	1.03	1.28	1.40	1.24	1.73	2.28
- Female	0.48	1.04	1.24	0.50	1.21	1.79
Inflow into DI (%)						
Total	0.99	1.01	0.94	1.10	1.08	1.05
- Male	0.49	0.51	0.48	0.60	0.60	0.56
- Female	0.50	0.50	0.46	0.50	0.48	0.49
- Fully disabled	0.30	0.36	0.34	0.40	0.46	0.44
- Partially disabled	0.69	0.65	0.60	0.70	0.62	0.61

Table 2 *Transitions within the experience rating system; selected sample, 2000-2001*

2000	**2001**			
small employer size	p_{min} = 0.98%	p_{min} < .. < p_{max}	p_{max} = 4.77%	total
p_{min} = 1.24%	8027	300	193	8520
p_{min} < .. < p_{max}	169	338	192	699
p_{max} = 4.17%	75	66	537	678
	8271	704	922	10076
medium sized employers (N<100)	p_{min} = 0.41%	p_{min} < .. < p_{max}	p_{max} = 6.36%	
p_{min} = 0.67%	20555	4068	217	24840
p_{min} < .. < p_{max}	2537	16050	1300	19877
p_{max} = 5.56%	46	264	683	993
	23138	20382	2200	45583
large employer size (N>99)	p_{min} = 0.41%	p_{min} < .. < p_{max}	p_{max} = 6.36%	
p_{min} = 0.67%	511	731	6	1248
p_{min} < .. < p_{max}	170	6093	95	6358
p_{max} = 5.56%	0	24	10	34
	681	6848	111	7640

Table 3 *Marginal incentives, average per sector in 2001 (standard deviations between brackets)*

Sector	**Pr ($p < p_{max}$)**	**%Δ wage sum**	**e_t × Pr ($p < p_{max}$)** **the average marginal incentive**
Primary sector	91.6	4.5	0.813 (0.163)
Industrial sector	94.4	5.4	0.967 (0.141)
Trade sector	94.2	6.4	0.980 (0.200)
Service industries	90.3	7.5	1.042 (0.252)
Transport	95.2	6.1	0.972 (0.177)
Catering	93.4	6.1	0.984 (0.190)
Social services / cultural	93.3	7.8	1.041 (0.177)
(semi-)public	95.7	4.7	0.966 (0.147)
Financial sector	96.5	8.5	1.020 (0.282)
Temp. empl. agencies	93.5	5.5	0.985 (0.271)
Unknown	94.5	7.6	1.016 (0.092)
Total	94.7	6.4	0.983 (0.197)

Table 4 *Grouped Logit estimates on cross sectional data (standard errors between parentheses)*

	2000		2001		2002	
Log likelihood	– 330625.1	(N=5181149)	– 379935.3	(N=5470778)	– 352624.8	(N=5462871)
Constant	– 5.84	(0.093)	– 5.87	(0.080)	– 6.34	(0.077)
Industrial sector	– 0.28	(0.036)	– 0.31	(0.035)	– 0.25	(0.034)
Trade sector	– 0.42	(0.037)	– 0.42	(0.035)	– 0.34	(0.035)
Service industries	– 0.12	(0.039)	– 0.38	(0.038)	– 0.51	(0.038)
Transport	– 0.30	(0.039)	– 0.25	(0.038)	– 0.22	(0.038)
Catering	– 0.36	(0.044)	– 0.37	(0.043)	– 0.40	(0.043)
Social services / cultural	– 0.18	(0.038)	– 0.48	(0.037)	– 0.48	(0.037)
Semi-public	– 0.35	(0.038)	– 0.29	(0.036)	– 0.42	(0.036)
Financial sector	– 0.53	(0.038)	– 0.47	(0.037)	– 0.39	(0.037)
Temp. empl. agencies	0.18	(0.038)	0.049	(0.038)	0.27	(0.038)
Public sector	– 0.37	(0.043)	– 0.31	(0.040)	– 0.53	(0.042)
Sector Unknown	– 0.076	(0.058)	– 0.11	(0.055)	– 0.15	((0.055)
10000 < wage < 20000	0.37	(0.025)	0.23	(0.023)	0.031	(0.024)
20000 < wage < 30000	0.40	(0.023)	0.39	(0.021)	0.21	(0.023)
30000 < wage < 40000	0.46	(0.023)	0.38	(0.022)	0.11	(0.023)
40000 < wage < 50000	0.47	(0.024)	0.38	(0.022)	0.035	(0.024)
50000 < wage < 60000	0.47	(0.025)	0.30	(0.023)	0.0060	(0.024)
60000 < wage < 70000	0.34	(0.028)	0.22	(0.026)	– 0.11	(0.028)
wage > 70000	0.82	(0.032)	0.32	(0.030)	– 0.0094	(0.032)
$15 < N \leq 25$	0.022	(0.027)	0.037	(0.028)	– 0.37	(0.027)
$25 < N \leq 50$	0.013	(0.025)	0.048	(0.027)	– 0.47	(0.025)
$50 < N \leq 100$	0.091	(0.025)	0.12	(0.027)	– 0.52	(0.024)
$100 < N \leq 250$	0.13	(0.025)	0.18	(0.026)	– 0.57	(0.024)
$250 < N \leq 1000$	0.13	(0.025)	0.22	(0.026)	– 0.53	(0.023)
$N > 1000$	0.071	(0.050)	0.25	(0.026)	– 0.47	(0.023)
% male, 25-35	0.95	(0.085)	0.98	(0.085)	1.13	(0.092)
% male, 35-45	1.40	(0.084)	1.79	(0.081)	1.93	(0.084)
% male, 45-55	1.17	(0.091)	1.89	(0.088)	2.02	(0.089)
% male, 55-65	0.23	(0.14)	0.67	(0.13)	1.68	(0.12)
% female, 15-25	0.20	(0.092)	0.26	(0.090)	0.56	(0.094)
% female, 25-35	1.53	(0.074)	1.65	(0.074)	1.77	(0.078)
% female, 35-45	1.21	(0.097)	1.56	(0.10)	1.98	(0.098)
% female, 45-55	2.10	(0.11)	2.07	(0.10)	1.52	(0.10)
% female, 55-65	0.31	(0.22)	1.22	(0.20)	1.81	(0.18)
$I(p < p_{max})$	0.096	(0.065)	– 0.081	(0.044)	– 0.20	(0.031)

Figure 1 *Estimated effect of experience rating on DI inflow for small en medium sized/large firms in 2002*

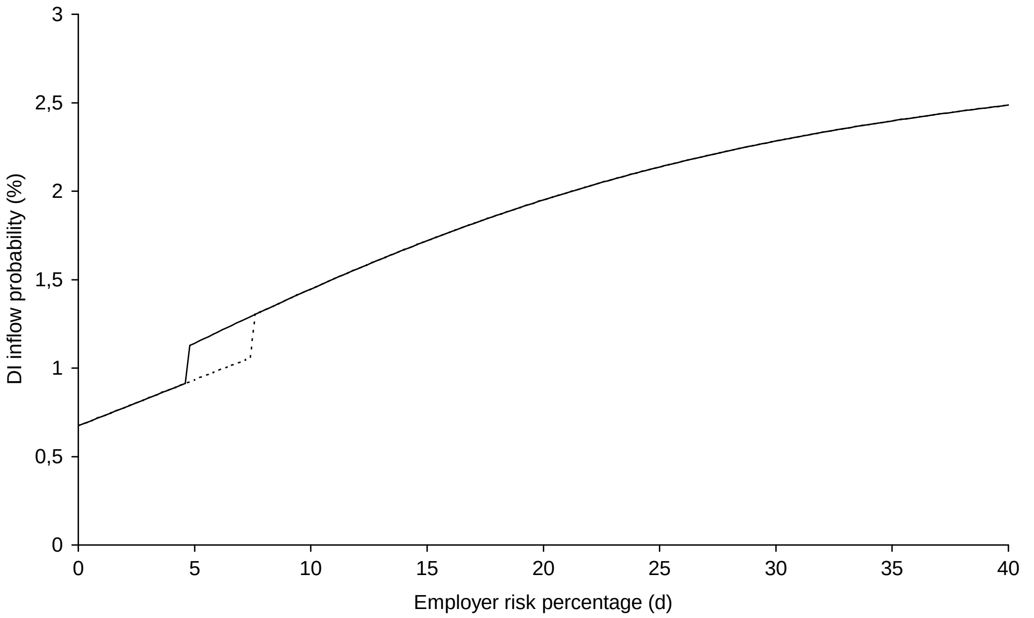

Table 5 *Baseline parameter values for the employer categories for 2000-2002*

	2000	2001	2002
$u = 0, v = 0$	0	$\Delta_{2001\text{-}2000}$	$\Delta_{2002\text{-}2000}$
$u = 0, v = 1$	α_{01}	$\alpha_{01} + \Delta_{2001\text{-}2000} + \Delta_{01}$	$\alpha_{01} + \Delta_{2002\text{-}2000} + \Delta'_{01}$
$u = 1, v = 0$	α_{10}	$\alpha_{10} + \Delta_{2001\text{-}2000} + \Delta_{10}$	$\alpha_{10} + \Delta_{2002\text{-}2000} + \Delta'_{10}$
$u = 1, v = 1$	α_{11}	$\alpha_{11} + \Delta_{2001\text{-}2000} + \Delta_{11}$	$\alpha_{11} + \Delta_{2002\text{-}2000} + \Delta'_{11}$

Table 6 *Grouped logit estimates on panel data (standard errors between parentheses), Part 1*

	Total DI inflow		Partial disability		Full disability	
Log likelihood	–1070818.6	(N=16114420)	– 538545.5		– 660922.6	
Constant	– 6.03	(0.041)	– 7.49	(0.072)	– 6.26	(0.050)
Industrial sector	– 0.28	(0.020)	– 0.15	(0.034)	– 0.35	(0.025)
Trade sector	– 0.45	(0.021)	– 0.40	(0.036)	– 0.46	(0.025)
Service industries	– 0.22	(0.022)	– 0.55	(0.040)	– 0.12	(0.027)
Transport	– 0.30	(0.022)	– 0.026	(0.036)	– 0.49	(0.028)
Catering	– 0.37	(0.025)	– 0.24	(0.044)	– 0.43	(0.030)
Social services / cultural	– 0.40	(0.021)	– 0.045	(0.036)	– 0.58	(0.026)
Semi-public	– 0.37	(0.021)	0.0085	(0.036)	– 0.69	(0.027)
Financial sector	– 0.53	(0.021)	– 0.32	(0.036)	– 0.63	(0.027)
Temp. empl. agencies	0.29	(0.022)	– 0.15	(0.040)	0.42	(0.026)
Public sector	– 0.50	(0.024)	– 0.18	(0.038)	– 0.81	(0.032)
Sector Unknown	– 0.081	(0.032)	– 0.025	(0.052)	– 0.084	(0.041)
10000 < wage < 20000	0.28	(0.014)	0.14	(0.028)	0.28	(0.016)
20000 < wage < 30000	0.40	(0.013)	0.52	(0.023)	0.33	(0.015)
30000 < wage < 40000	0.31	(0.013)	0.60	(0.022)	0.15	(0.016)
40000 < wage < 50000	0.26	(0.013)	0.66	(0.022)	– 0.029	(0.017)
50000 < wage < 60000	0.20	(0.013)	0.67	(0.022)	– 0.17	(0.018)
60000 < wage < 70000	0.058	(0.015)	0.50	(0.024)	– 0.30	(0.020)
wage > 70000	0.30	(0.018)	0.70	(0.028)	– 0.0017	(0.024)
N > 15	– 0.070	(0.016)	– 0.063	(0.025)	– 0.069	(0.020)
25 < N ≤ 50	– 0.10	(0.015)	– 0.11	(0.023)	– 0.095	(0.019)
50 < N ≤ 100	– 0.090	(0.015)	– 0.087	(0.023)	– 0.086	(0.019)
100 < N ≤ 250	– 0.0091	(0.014)	– 0.064	(0.023)	– 0.11	(0.019)
250 < N ≤ 1000	– 0.085	(0.014)	– 0.027	(0.022)	– 0.13	(0.019)
N > 1000	– 0.096	(0.014)	– 0.0036	(0.022)	– 0.17	(0.019)
% male, 25-35	1.00	(0.13)	0.77	(0.22)	1.09	(0.17)
% male, 35-45	0.96	(0.14)	0.47	(0.22)	1.17	(0.17)
% male, 45-55	1.25	(0.14)	1.45	(0.22)	0.85	(0.18)
% male, 55-65	1.27	(0.17)	0.99	(0.25)	1.66	(0.23)
% female, 15-25	0.31	(0.16)	– 0.32	(0.28)	0.52	(0.20)
% female, 25-35	0.88	(0.14)	1.02	(0.22)	0.75	(0.18)
% female, 35-45	0.65	(0.14)	0.80	(0.22)	0.41	(0.18)
% female, 45-55	1.73	(0.15)	1.75	(0.23)	1.65	(0.19)
% female, 55-65	2.20	(0.24)	1.35	(0.36)	2.65	(0.32)
% male–mean, 25-35	– 0.012	(0.14)	0.19	(0.23)	0.073	(0.18)
% male–mean, 35-45	1.11	(0.14)	1.67	(0.23)	0.92	(0.18)
% male–mean, 45-55	0.49	(0.15)	0.92	(0.23)	0.40	(0.19)
% male–mean, 55-65	– 0.088	(0.17)	0.022	(0.26)	– 0.25	(0.24)
% female–mean, 15-25	0.13	(0.17)	0.079	(0.29)	0.17	(0.21)
% female–mean, 25-35	0.77	(0.15)	0.68	(0.23)	0.91	(0.19)
% female–mean, 35-45	1.23	(0.15)	1.15	(0.24)	1.39	(0.19)
% female–mean, 45-55	0.66	(0.15)	0.67	(0.24)	0.77	(0.20)
% female–mean, 55-65	– 0.82	(0.26)	– 0.58	(0.39)	– 0.91	(0.34)

Table 6 *Grouped Logit estimates on panel data (standard errors between parentheses), Part 2*

	Total DI inflow		Partial disability		Full disability	
Year = 2001	0.17	(0.015)	0.34	(0.023)	0.051	(0.020)
Year = 2002	0.17	(0.015)	0.30	(0.024)	0.077	(0.020)
Average in 2000						
α_{01}	0.31	(0.016)	0.23	(0.024)	0.39	(0.020)
α_{11}	0.47	(0.013)	0.35	(0.020)	0.55	(0.016)
α_{10}	0.18	(0.022)	0.26	(0.032)	0.12	(0.031)
Difference estimator 2000 – 2001						
Δ_{01}	– 0.044	(0.021)	– 0.067	(0.031)	– 0.068	(0.028)
Δ_{11}	– 0.14	(0.017)	– 0.14	(0.026)	– 0.14	(0.022)
Δ_{10}	– 0.011	(0.032)	0.038	(0.042)	– 0.017	(0.046)
Difference estimator 2000 – 2002						
Δ'_{01}	– 0.16	(0.021)	– 0.17	(0.032)	– 0.19	(0.028)
Δ'_{11}	– 0.25	(0.017)	– 0.21	(0.026)	– 0.28	(0.022)
Δ'_{10}	– 0.26	(0.032)	– 0.24	(0.044)	– 0.35	(0.048)

Table 7 *Grouped Logit estimates of panel models (standard errors between parentheses), Part 1*

	Model (i)		Model (ii)		Model (iii)	
Log likelihood	–1070826.8	N=16114420	– 1070521.3		– 1070810.8	
Constant	– 6.03	(0.041)	– 6.09	(0.041)	– 6.04	(0.041)
Industrial sector	– 0.28	(0.020)	– 0.28	(0.020)	– 0.28	(0.020)
Trade sector	– 0.45	(0.021)	– 0.45	(0.021)	– 0.45	(0.021)
Service industries	– 0.22	(0.022)	– 0.22	(0.022)	– 0.22	(0.022)
Transport	– 0.30	(0.022)	– 0.29	(0.022)	– 0.30	(0.022)
Catering	– 0.37	(0.025)	– 0.37	(0.025)	– 0.37	(0.025)
Social services / cultural	– 0.39	(0.021)	– 0.39	(0.021)	– 0.40	(0.021)
Semi-public	– 0.37	(0.021)	– 0.40	(0.021)	– 0.37	(0.021)
Financial sector	– 0.53	(0.021)	– 0.53	(0.021)	– 0.53	(0.021)
Temp. empl. agencies	0.29	(0.022)	0.30	(0.021)	0.29	(0.022)
Public sector	– 0.50	(0.024)	– 0.51	(0.024)	– 0.50	(0.024)
Sector Unknown	– 0.079	(0.032)	– 0.079	(0.032)	– 0.081	(0.032)
10000 < wage < 20000	0.28	(0.014)	0.30	(0.014)	0.28	(0.014)
20000 < wage < 30000	0.40	(0.013)	0.42	(0.013)	0.40	(0.013)
30000 < wage < 40000	0.31	(0.013)	0.35	(0.013)	0.31	(0.013)
40000 < wage < 50000	0.26	(0.013)	0.30	(0.013)	0.26	(0.013)
50000 < wage < 60000	0.20	(0.014)	0.25	(0.014)	0.20	(0.014)
60000 < wage < 70000	0.060	(0.015)	0.10	(0.015)	0.058	(0.015)
wage > 70000	0.30	(0.018)	0.34	(0.018)	0.30	(0.018)
N > 15	– 0.069	(0.016)	– 0.044	(0.016)	– 0.070	(0.016)
25 < N ≤ 50	– 0.10	(0.015)	– 0.061	(0.015)	– 0.098	(0.015)
50 < N ≤ 100	– 0.088	(0.015)	– 0.032	(0.015)	– 0.083	(0.015)
100 < N ≤ 250	– 0.089	(0.015)	– 0.026	(0.015)	– 0.085	(0.015)
250 < N ≤ 1000	– 0.084	(0.014)	– 0.016	(0.015)	– 0.079	(0.015)
N > 1000	– 0.095	(0.014)	– 0.024	(0.014)	– 0.090	(0.014)
% male, 25-35	0.99	(0.13)	1.03	(0.13)	1.00	(0.13)
% male, 35-45	0.94	(0.14)	0.95	(0.14)	0.96	(0.14)
% male, 45-55	1.24	(0.14)	1.24	(0.14)	1.25	(0.14)
% male, 55-65	1.26	(0.17)	1.36	(0.16)	1.27	(0.17)
% female, 15-25	0.29	(0.16)	0.37	(0.16)	0.31	(0.16)
% female, 25-35	0.87	(0.14)	0.92	(0.14)	0.88	(0.14)
% female, 35-45	0.64	(0.14)	0.70	(0.14)	0.66	(0.14)
% female, 45-55	1.72	(0.15)	1.73	(0.15)	1.73	(0.14)
% female, 55-65	2.19	(0.24)	2.36	(0.24)	2.20	(0.24)
% male–mean, 25-35	0.020	(0.14)	– 0.037	(0.14)	0.015	(0.14)
% male–mean, 35-45	1.12	(0.14)	1.083	(0.14)	1.10	(0.14)
% male–mean, 45-55	0.50	(0.15)	0.49	(0.15)	0.49	(0.15)
% male–mean, 55-65	– 0.082	(0.17)	– 0.20	(0.17)	– 0.088	(0.17)
% female–mean, 15-25	0.15	(0.17)	0.058	(0.17)	0.13	(0.17)
% female–mean, 25-35	0.78	(0.15)	0.72	(0.15)	0.77	(0.15)
% female–mean, 35-45	1.25	(0.15)	1.19	(0.15)	1.23	(0.15)
% female–mean, 45-55	0.67	(0.15)	0.64	(0.15)	0.66	(0.15)
% female–mean, 55-65	– 0.81	(0.26)	– 1.04	(0.25)	– 0.82	(0.26)

Table 7 *Grouped Logit estimates of panel models (standard errors between parentheses), Part 2*

	Model (i)		Model (ii)		Model (iii)	
Year = 2001	0.17	(0.014)	0.17	(0.015)	0.17	(0.015)
Year = 2002	0.15	(0.014)	0.17	(0.015)	0.17	(0.015)
Average in 2000						
α_{01}	0.30	(0.012)	0.27	(0.016)	0.30	(0.016)
α_{11}	0.46	(0.012)	0.42	(0.013)	0.46	(0.013)
α_{10}	0.18	(0.021)	0.16	(0.023)	0.19	(0.022)
α_{12}			0.81	(0.029)		
α_{22}			0.71	(0.030)		
α_{21}			0.54	(0.034)		
Difference estimator 2000 – 2001						
Δ_{01}	– 0.052	(0.019)	– 0.036	(0.021)	– 0.078	(0.045)
Δ_{01} *x* I($N > 25$)					0.036	(0.042)
Δ_{11}	– 0.14	(0.016)	– 0.13	(0.017)	– 0.14	(0.017)
Δ_{10}	– 0.0078	(0.032)	0.0053	(0.033)	0.011	(0.032)
Δ_{12}			– 0.23	(0.044)		
Δ_{22}			– 0.24	(0.043)		
Δ_{21}			– 0.16	(0.045)		
Difference estimator 2000 – 2002						
Δ'_{01}	– 0.14	(0.020)	– 0.16	(0.021)	– 0.020	(0.042)
Δ'_{01} *x* I($N > 25$)					– 0.16	(0.040)
Δ'_{11}	– 0.24	(0.016)	– 0.26	(0.017)	– 0.26	(0.017)
Δ'_{10}	– 0.24	(0.032)	– 0.25	(0.033)	– 0.26	(0.032)
Δ'_{12}			– 0.35	(0.043)		
Δ'_{22}			– 0.30	(0.044)		
Δ'_{21}			– 0.18	(0.045)		