A shorter term of notice: Were older employees fired more often after the introduction of the 'Flexwet'?

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Abstract

In this paper the causal relation between firing costs and the probability of being fired is assessed. To do so a Difference-in-Differences approach is applied. The variable of interest is a dummy stating whether someone was fired from her previous tenured job during the year of enumeration. I use employment data of older workers from the Dutch Socio-Economic Panel for the years 1996 until 2002. With the introduction of the so-called 'Flexwet' in The Netherlands on January 1st, 1999, an employee's age ceased to be a determinant in the formula to calculate her legal term of notice. The new formula increased the term of notice for some employees and decreased it for others. The term of notice is an important element of the firing costs an employer faces when she would like to fire a worker. I find evidence that this indeed led to relatively more firings in the treated group.

Keywords: Employment protection, Older workers, Layoffs, Difference-in-differences, Panel data models

JEL-classification: C21, C23, J14, J65

1 Introduction

Employment protection is widely debated in the political arena nowadays. On the one hand employers complain that they are incapable of adapting to economic circumstances because of high firing costs and on the other hand vulnerable groups of employees complain that the firing risk they face is too large. Policy-makers have to decide upon an optimal level of protection in the midst of this. In order to make such decisions it is interesting to see how different types of employment protection affect labor market outcomes. This paper focuses on a particular kind of firing costs, namely the term of

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notice. As specific groups, notably older workers, are often 'protected' by longer terms of notice we would like to know whether this actually lowers the probability to be fired. An attempt is made to establish a causal relationship in this paper.

Whether a profit-maximizing firm decides to fire a worker with a permanent contract primarily depends on the productivity and wage of the worker. When the difference between these entities is negative, an employer loses money. Assuming that all employees receive a minimal wage, the firm will then consider to fire the worker. It is costly however to adjust the number of employees downwards because of firing costs. Furthermore, it is also costly to attract a new employee in the future because of hiring and training costs. So it could be rational to defer the irreversible decision to fire a worker to circumvent firing and hiring costs. The firm will indeed do so when both the probability of a change for the better (e.g. an economic upturn or an individual productivity increase) and the costs of firing the old and hiring a new worker are sufficiently large. So theoretically, higher costs of firing an existent employee lowers the propensity to fire (but also hire) a worker.

An extensive literature looks at this relationship. On the theoretical side, Bentolila and Bertola (1990) for example built a model by assuming linear adjustment costs and they find that firing costs might actually increase overall long-run employment. Addison and Teixeira (2003) provide an overview of the empirical literature. Lazear (1990) was the first to test the relationship between severance pay and unemployment using international macro-data. He finds significant effects of severance pay on labor markets; increases in severance pay both reduced the number of jobs as well as the size of the labor force. The use of micro-data in the employment protection literature is relatively new however. Pfann (2006) develops a model based on the heterogeneity of workers and tests his findings using micro-data of a mass-layoff in one firm. He concludes that individual characteristics help to explain firing probabilities of workers.

This paper would like to empirically establish the causal relation between terms of notice and firing decisions of firms by using longitudinal micro-data. Section 2 lays out the terms of notice regulations in The Netherlands and introduces the so-called 'Flexwet'. Section 3 then explains employed Difference-in-Differences methodology. The data that is used to establish an empirical relation receives attention in Section 4. The specific empirical strategy utilized will be explained in Section 5. Results will finally be presented in Section 6. Section 7 concludes and gives suggestions for further research.

2 Term of notice/ The 'Flexwet'

One of the main components of firing costs is the term of notice (ToN). Employers in most, at least European, countries are legally required to notify a fired employee several weeks or months in advance before that person is actually fired. When it is assumed that firms only wish to fire loss-making employees, the length of the term of notice partly determines the size of the loss. Furthermore, employees who know they will be fired won't be very motivated to add value to the firm. The losses made will therefore be aggravated when a worker is already dismissed. Thus, the longer the term of notice the higher the firing costs of an employee for the employer. Other components of firing costs include the severance pay and legal costs for dissolving a contract.

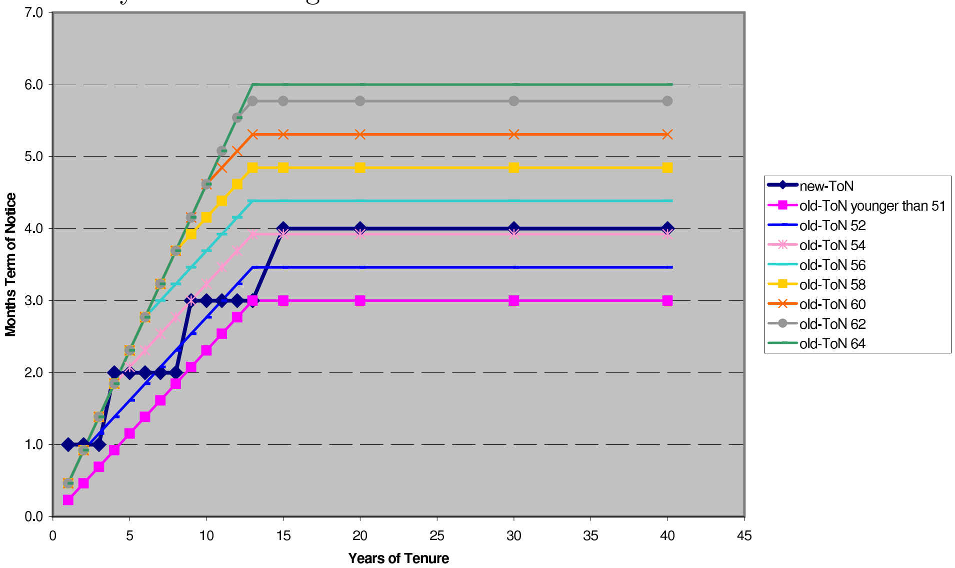
The legal terms of notice in The Netherlands changed on January first, 1999. This policy change will be utilized in the Difference-in-Differences methodology of this paper. A law was introduced that was meant to render the labor market more flexible, appropriately called the 'Flexwet'. On the one hand, it was made more attractive to hire employees with temporary contracts. These workers also received better legal protection than before. On the other hand, employees with permanent contracts lost some rights. See Van der Geest, Koopmans and Stavenuiter (2000) for a more detailed description of the changes. The formula for the legal term of notice of tenured employees was adjusted and simplified.[1] Age was removed from the ToN-calculation and the number of possible terms was lowered from 26 to four. This policy change therefore decreased the terms of notice of some employees but increased it for others.

The legal Term of notice (ToN) before 1st of January, 1999 will from now on be referred to as the old-ToN. The old-ToN had a length of one week for every year of tenure, with a maximum of 13 weeks. On top of this workers received an extra week of notice for each year they had worked while being 50 or older, with a maximum of 13 weeks. So, two workers with identical tenures (say 15 years) but different ages (say 40 and 60) could face different terms of notice (in this case 13 weeks and 24 weeks).

The legal term of notice (ToN) after the 1st of January, 1999 will from now on be referred to as the new-ToN. The new-ToN does not depend on age. Any two workers with identical tenures will have to be told the same number of months in advance about their dismissal. Workers who are employed in between zero and four years will receive a ToN of one month.

[1] Although the term of notice is set out by law it is possible to divert from this in a collective wage agreement. Furthermore, it is possible for an employer to circumvent the terms of notice by going to court to dissolve the employment contract. The Dutch ministry of social affairs reports in its annual 'Ontslagrapportage' that from 1996 to 2002 a little over half of annual firings were applied for at the 'CWI', thereby involving the legal term of notice. Nothing changed in the court-procedure over the analyzed period.

Figure 1: Legal Term of Notice in old (pre-1999) and new (1999 and after) scheme by tenure and age



When employed between five and nine years employees will receive notice two months in advance. If a worker's tenure is between ten and fourteen years, his or her employer will face three months of ToN when he or she would like to fire the worker. Any tenure longer than fourteen years results in a new-ToN of four months.

Whether the old or new term of notice is the shortest thus depends on age and tenure. As the relationship is not so straightforward, figure 1 shows the number of months of notice per number of years of tenure for selected ages. As the new legal term of notice is independent of age, only one line represents this scheme, namely the stepwise increasing line. All employees younger than 51 have an equal old-ToN, depicted by the lowest line. In fact, for those in the youngest age-group the new-ToN is always equal or higher than the old-ToN. This group is thus better protected against firing after the introduction of the 'Flexwet'. Relatively young older workers (aged 51-53) that have been working at a particular firm for a long time (from fourteen years of tenure onwards) are also better protected in the new scheme. Young older workers (aged 51-53) with a lower tenure however might be faced with a lower ToN, depending on the exact years of tenure. Relatively old older workers (aged 55 and older) that have worked in a job for longer than three years enjoyed a higher term of notice in the old scheme. This group of old older workers is hence worse protected through the 'Flexwet' than before. Only old older workers with a very low tenure benefit from the introduction of the 'Flexwet' in terms of a higher term of notice.

3 Methodology

Is it true that employees that are more costly to fire are dismissed less often? Although it is relatively simple to find evidence of an association between firing costs and employment status, it is much harder to infer upon causality. To answer the causality question we need information on firing probabilities of employees with low and high firing costs who are identical in everything else. However, this information is practically unavailable as terms of notice depend on observable characteristics. Longer tenured employees are rewarded for their loyalty with higher legal terms of notice. Also older employees often (or at least in the past) have to be notified longer in advance of their upcoming dismissal. This was installed, at least in The Netherlands, as it was assumed that older workers will have more trouble finding new, suitable employment and therefore they need more time to search for it. Firing costs thus differ across groups of workers, but whether you belong to a group with short or long tenure is in part endogenous and depends on unobservable characteristics such as work attitude. Therefore a specific empirical strategy is needed to answer the research question.

A methodology called Difference-in-Differences (DID) will be employed to assess whether firing costs can be causally linked to firing probabilities. Card (1989) and Card and Krueger (1994) were among the first to employ this technique in the field of labor economics. DID takes into account that in order to find a causal relationship between two variables one has to compare the outcomes of groups that only differ in terms of the variable one is interested in. To do so, one needs to have information on the outcome variable, on the cause-variable and on an exogenous change in the cause-variable (i.e. the treatment). It is essential for a DID analysis to find an exogenous change that affects one group of the population (i.e. the treated group) and that does not affect another group of the population (i.e. the control group). In my analysis, the introduction of the 'Flexwet' will serve as an exogenous policy change.

If being a member of the treated group is fully exogenous one could identify causality by simply comparing the outcomes of the two groups. However, experiencing the 'treatment' is often endogenous and therefore outcomes over time are needed. When observations of both groups are present before and after the change in the cause-variable it is possible to distinguish a causal effect. DID compares the outcomes after and before of the treated group to the outcomes after and before of the control group. Hereby one compares outcomes of individuals that are identical except for the cause-variable. The control group trend is subtracted in order to take out a time trend in the outcome variable. An essential assumption for DID is that the time trend is equal to both groups.

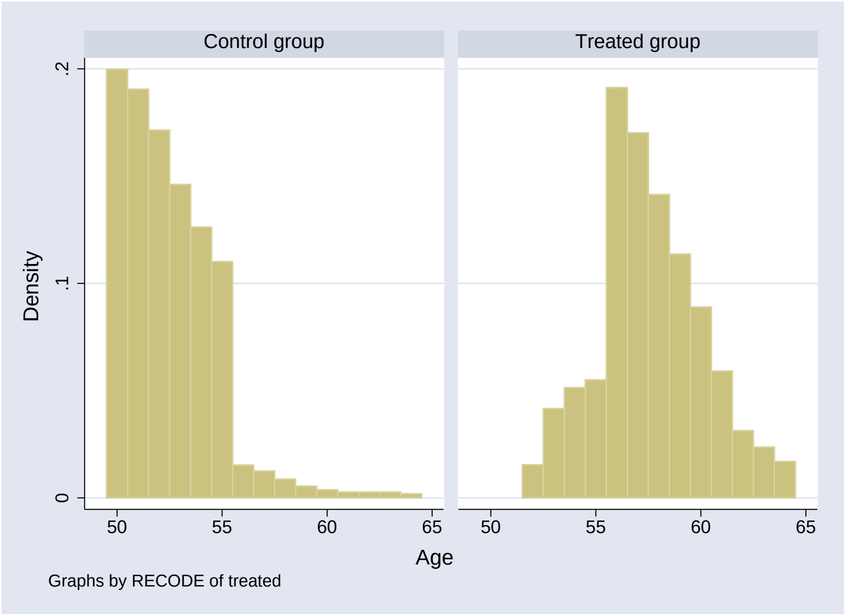
4 Data

Seven waves of the Dutch Socio-Economic Panel (SEP) were used for the empirical analysis of the research question (1996-2002). This longitudinal dataset has been collected annually by the Dutch Central Bureau of Statistics from 1984 onwards. The SEP is collected every April. I have used all available waves after the 'treatment' (after the collection of the 1999 wave) and an equal amount of waves before the 'treatment'. Although the same individuals are ususally observed multiple times in my sample, I do not use a balanced panel. The sample consists of individuals who are 50 to 64 years old. This age-restriction was chosen in order to meet the difference-in-difference assumption that time trends should be identical to both the treated and the control group. If younger respondents would be included in the analysis other age-specific policy alterations, such as changes in early retirement schemes, could render the time trend in firing probabilities between the treated and untreated groups unequal. The treatment effect of the introduction of the 'Flexwet' could then become unidentifiable.

The treated group consists of individuals who had tenure (i.e. were employed with a permanent contract) in the previous survey-period who would, when employed, face a shorter new-ToN in the current survey-year than an old-ToN in the previous year. In the control group, we find the previous survey's tenured employees who would, if employed, face an equal or longer new-ToN this year than an old-ToN last year. To decide whether someone is a member of the treated or the control group I thus compare the terms of notice someone would have received last year under the old legislation and the term of notice someone would receive this year under the new legislation when employed. Membership of the treatment group therefore only depends on a respondent's particular age, tenure vector (denoted as (A_i, L_i)). 3,669 individuals meet the requirements for the control group in my dataset and 1,945 individuals meet the requirements for the treated group in the dataset. Note that only individuals between 50 and 64 that were employed with a permanent contract last year are included in the sample.

For example: a 51-year old with a tenure of ten years would be in the control group. He or she would have received eleven weeks of notice in the old situation in the previous survey-period (ten times one week plus one week extra for the year worked when older than fifty), but would receive three months of notice in the new situation this year (eleven years of tenure gives rise to three months of notice). However, a 52-year old who has been employed for eight years would be in the treated group. Last year in the old situation, the employer would have been forced to provide a ToN of ten weeks (eight times one week plus two weeks for the years worked when older than fifty) whereas this year in the new situation, the legal ToN would be two months (eight years of tenure gives rise to two months of notice). Note that this is not necessarily linear in age. A 57-year old with a tenure of three

Figure 2: Histogram of ages in control and treated group



years will be a member of the treated group and a 58-year old with a tenure of four years will be a member of the control-group. This also means that over time individuals can move back and forth between groups.

Figures 2 and 3 show histograms of the ages and tenure years of individuals in the control (left) and treated (right) groups. Although almost each age and tenure is present in both groups, the treated group contains more older workers and more workers with a long tenure.

Some summary statistics of the treated and control group can be found in tables 1 and 2. Table 3 presents the t-statistics of t-tests on the equality of means in the two groups. The tests indicate that the groups are different in nature. The control group is younger, lower tenured and works longer hours. However, in terms of gross wages there is no significant difference between the groups. If we believe that wages fully represent productivity the groups are hence comparable. Note that the groups are allowed to be different in observed and unobserved characteristics, as long as the time-trend in the groups' outcome-variable is equal. As both groups are relatively equal in age and equal in wages I believe this requirement is met.

The dependent variable on whether someone was fired from his or her previous job during the last year is not directly observed in the SEP. Therefore this dummy is constructed by combining several variables. Someone is considered to be fired in the period between the previous and the current survey when he or she was employed during the last survey, indicates in this survey that his or her last job terminated within a year ago of the current survey and when the reason for that termination was a forced resignation.

Figure 3: Histogram of tenure in control and treated group

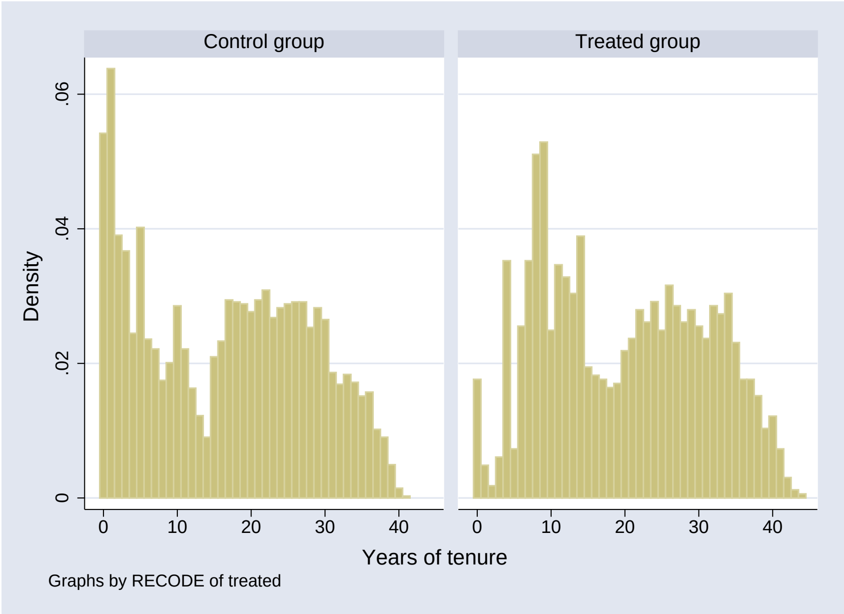


Table 1: Summary statistics for control group

Variable	Obs	Mean	Std. Dev.	Min	Max
Age	3,669	52.5	2.2	50	64
Tenure	3,432	16.6	11.4	0	41
Gross yearly wages	3,267	62,699	41,552	2	733,342
Dummy for working part-time	3,496	.32	.47	0	1
Working hours	3,669	35.0	14.2	0	99
Dummy for permanent contract	3,496	.86	.34	0	1
Dummy for male	3,669	.65	.48	0	1
dummy1996	3,669	.12	.33	0	1
dummy1997	3,669	.14	.34	0	1
dummy1998	3,669	.14	.35	0	1
dummy1999	3,669	.14	.35	0	1
dummy2000	3,669	.15	.36	0	1
dummy2001	3,669	.15	.36	0	1
dummy2002	3,669	.15	.36	0	1

Table 2: Summary statistics for treated group

Variable	Obs	Mean	Std. Dev.	Min	Max
Age	1,945	57.6	2.5	52	64
Tenure	1,647	20.0	11.0	0	50
Gross yearly wages	1,677	61,586	41,548	1	768,534
Dummy for part-time	1,690	.40	.49	0	1
Working hours	1,944	30.9	17.0	0	90
Dummy for permanent contract	1,690	.84	.37	0	1
Dummy for male	1,945	.66	.47	0	1
dummy1996	1,945	.11	.31	0	1
dummy1997	1,945	.13	.34	0	1
dummy1998	1,945	.14	.34	0	1
dummy1999	1,945	.14	.34	0	1
dummy2000	1,945	.14	.35	0	1
dummy2001	1,945	.17	.37	0	1
dummy2002	1,945	.18	.38	0	1

Table 3: T-tests on the equality of means between the treated and control group

Variable	t-value
Age	-78.2
Tenure	-10.2
Gross yearly wage	0.9
Dummy for part-time	-5.8
Working hours	9.6
Dummy for permanent contract	2.1
Dummy for male	0.8

Two different paths lead to being considered as not being fired. Someone is not fired when she was employed during the last survey and when she indicates in this survey that her last job ended longer than a year ago and someone is not fired when her last job terminated within a year ago of the current survey but the reason for that termination was another job or other personal reasons. Of those older than 49 and younger than 65, 3,867 observations in my dataset were not fired according to my definition. Only 49 observations were forced to resign during the previous year. This comes down to 1,236 individuals who are never fired, 26 individuals who enter the dataset as both fired and not-fired and 23 individuals who enter the dataset only as being fired. Although this is a very small number the regressions in the next section render significant results.

Some summary statistics of those who were fired and those who were not fired can be found in table 4 and 5. Table 6 presents the t-statistics of t-tests on the equality of means in the two groups. These tests indicate, like for the treated and control group, that the groups are different in nature. Information collected in the current survey is used to compile these tables (as opposed to lagged information from last year's survey). This explains why some respondents who are fired are at the same time employed. These have started a new job already during the previous year. Overall, recently fired individuals have lower tenure and have a permanent contract much less often. Table 4 shows that there is some measurement error in the data. Individuals who I believe to have been fired during the last year have an average tenure of 5.5 years. Closer inspection reveals that this is due to an outlier of 35 years. Because of the small number of dismissals the observation is not omitted from the sample.

5 Empirical Strategy

Table 3 shows that the treated and control group are different when observable characteristics are taken into account. On top of this it is likely that these differ in unobservable characteristics. Hence, the groups are not directly comparable. When pre- and post-treatment information is available the Difference-in-Differences estimator can then be estimated to answer the research question.

The DID-estimator is defined as

$$\begin{aligned}\beta = [E(Y|T=1, D=1) - E(Y|T=0, D=1)] - \\ [E(Y|T=1, D=0) - E(Y|T=0, D=0)]\end{aligned} \tag{1}$$

where Y denotes whether an individual is fired. $T = 0$ is the period before the new law was introduced and $T = 1$ is the period after the introduction of the law. $D = 1$ denotes an individual that faces lower ToN this year (i.e. is a member of the treated group) and $D = 0$ denotes an individual that faces

Table 4: Summary statistics for fired group

Variable	Obs	Mean	Std. Dev.	Min	Max
Age	46	53.6	3.2	50	62
Tenure	11	5.5	12.5	0	35
Gross yearly wages	42	45,298	28,572	3,653 135,919	
Dummy for part-time	12	.5	.52	0	1
Working hours	46	8.3	15.3	0	48
Dummy for permanent contract	12	.33	.49	0	1
Dummy for male	46	.59	.50	0	1
dummy1996	46	.26	.44	0	1
dummy1997	46	.17	.38	0	1
dummy1998	46	.17	.38	0	1
dummy1999	46	.02	.15	0	1
dummy2000	46	0	0	0	0
dummy2001	46	.09	.28	0	1
dummy2002	46	.28	.46	0	1

Table 5: Summary statistics for not-fired group

Variable	Obs	Mean	Std. Dev.	Min	Max
Age	3,796	54.09	3.2	50	64
Tenure	3,518	16.6	9.8	0 43	
Gross yearly wages	3,752	59,588	42,004	69	768,534
Dummy for part-time	3,796	.40	.49	0	1
Working hours	3,794	33.7	12.8	2	80
Dummy for permanent contract	3,796	.98	.14	0	1
Dummy for male	3,796	.61	.49	0	1
dummy1996	3,796	.13	.33	0	1
dummy1997	3,796	.14	.34	0	1
dummy1998	3,796	.14	.34	0	1
dummy1999	3,796	.14	.35	0	1
dummy2000	3,796	.15	.35	0	1
dummy2001	3,796	.16	.36	0	1
dummy2002	3,796	.16	.37	0	1

Table 6: T-tests on the equality of means between the fired and not-fired group

Variable	t-value
Age	-2.9
Tenure	12.9
Gross yearly wage	3.26
Dummy for part-time	0.44
Working hours	9.6
Dummy for permanent contract	32.7
Dummy for male	-0.07

equal or higher ToN (i.e. is a member of the control group). β measures the difference between the difference in the percentage of fired employees in the treated group after 1998 and before 1999 and the same difference for the control group.
The DID-parameter β can also be obtained as the coefficient of the interaction term in a regression of outcomes on treatment and time dummies. Cameron and Trivedi (2005) explain this in detail.. One can write

$$Y^* = \beta_0 + \beta_1 D + \beta_2 T + \beta_3 (D * T) + \epsilon \tag{2}$$

where Y^* is a latent variable representing the desire of an employer to fire an employee. This latent variable is mapped into a binary variable that represents whether the employee is actually fired or not. In this equation β_3 equals the DID-estimator. I will estimate this equation through a probit regression.

The dataset furthermore has a panel character, i.e. multiple individuals are observed a number of times (even within $T = 0$ and $T = 1$). The data will then be serially correlated by construction as the error terms will be correlated for every year for given individuals. This problem is 'solved' by using random effect estimators. Fixed effects is impossible as the number of individuals with both fired and not-fired observations is very little. So the error terms of the regressions are assumed to consist of a term that is randomly distributed over all observations (ϵ) and an individual-specific term that is randomly distributed over all individuals (υ), i.e. to use a random effects probit model. The following equation is estimated in the remainder of this paper.

$$Y^*_{it} = \beta_0 + \beta_1 D_{it} + \beta_2 T_t + \beta_3 (D * T)_{it} + \upsilon_i + \epsilon \tag{3}$$

$$Y_{it} = \begin{cases} 0 & \text{if } Y^*_{it} < 0.5, \\ 1 & \text{if } Y^*_{it} >= 0.5. \end{cases} \tag{4}$$

6 Results

To provide some additional information on the association between firing costs and the probability of being fired I have first estimated an RE probit model of being fired on last year's term of notice, tenure and several other relevant independent variables, such as age and gross wage. The results of this exercise can be found in table 7. Column one shows the coefficients of the probit regression of all observations, so also of individuals younger than 50. Column two presents the results of the same regression but now including lagged wage information as a regressor. Column three presents the results of the same regression as in column one, but the ages of the respondents are restricted to higher than 49 and lower than 65.

The coefficient of the variable of interest, the lagged term of notice, is significant and negative in both the regressions involving observations of all ages as in the regression for the elderly. A higher term of notice and consequently higher firing costs are thus associated with higher firing rates, even when we control for age, wage, tenure and year. The variable on lagged years of tenure is furthermore positive and significant in the regression on older workers only.

This paper was written however to establish more than a relationship between the terms of notice and the odds of being fired. I would like to find a causal relationship. Table 8 therefore shows the coefficients and t-statistics of the Difference-in-Differences-probit model. Column one does not condition on any observable characteristic, column two just conditions on the years of enumeration and column three conditions on some other independent variables as well. All three regression utilize the same observations. The degree of the error term attributable to time-invariant individual-specific error term is negligible in all regressions ($\rho = 0\%$).

β_1 is negative in all columns and significant in the first column. This means that taking into account some controls, those in the treated group are not very different from the control groups in terms of the propensity to get fired. β_2 is however negative and significant on the one percent level for all regressions. After the first of Jan, 1999, apparently significantly less dismissals occurred in both the treated and the control group. This could be an economic cycle-effect. When the year dummies are included this effect even grows in magnitude. Note that the dummy for 1996 is omitted to prevent estimation with an identity matrix and note that the dummy for 2000 has no standard error because no one was fired in that year. There have been a significant higher number of firings in 2001 and 2002, perhaps because of an economic downturn.

The DiD-parameter, β_3, which we are primarily interested in, is positive and significant at the five percent level for all regressions. This indicates that there might indeed be a causal effect of lower firing costs on higher firing rates. In fact, the treatment effect for the treated is estimated to be as

Table 7: Coefficients and t-statistics RE probit model to test association

	(1 - all ages) Prob. of being fired	(2 - all ages) Prob. of being fired	(3 - only age 50-64) Prob. of being fired
Lagged term of notice	-0.163**	-0.166**	-0.184*
	(-2.94)	(-2.98)	(-1.98)
Lagged tenure	0.0140	0.0144	0.0276**
	(1.74)	(1.78)	(2.67)
Age	-0.0353	-0.0288	-0.238
	(-1.41)	(-1.12)	(-0.41)
Squared age	0.000449	0.000373	0.00214
	(1.32)	(1.08)	(0.40)
Dummy being over 49	-0.0140	-0.00545	
	(-0.11)	(-0.04)	
Dummy 1997	-0.0883	-0.0817	-0.191
	(-1.08)	(-0.99)	(-0.99)
Dummy 1998	-0.196*	-0.190*	-0.259
	(-2.24)	(-2.16)	(-1.30)
Dummy 1999	-0.370***	-0.362***	-0.567*
	(-3.79)	(-3.70)	(-2.34)
Dummy 2000	-0.243*	-0.235*	-6.841
	(-2.48)	(-2.39)	(-0.00)
Dummy 2001	-0.262**	-0.254**	-0.486*
	(-2.69)	(-2.59)	(-2.17)
Dummy 2002	0.00849	0.0102	-0.00783
	(0.10)	(0.12)	(-0.05)
Lagged gross yearly wage		-0.0000	
		(-0.44)	
N	17,742	17,633	3,608
Ind	5,224	5,187	1,205
aic	2,580	2,564	490
rho	0.000	0.000	0.000

t-statistics in parentheses

constant and controls for gender included but not displayed

* p<.05, ** p<.01, *** p<.001

Table 8: Coefficients and t-statistics RE probit model to test causality

	(1) Prob. of being fired	(2) Prob. of being fired	(3) Prob. of being fired
Treated (D)	-0.407*	-0.393	-0.385
(β_1)	(-2.00)	(-1.91)	(-1.56)
After treatment (T)	-0.487***	-2.357***	-2.386***
(β_2)	(-3.42)	(-3.98)	(-4.02)
Treated*After Treatment	0.588*	0.572*	0.563*
(β_3)	(2.22)	(2.08)	(2.03)
Dummy 1997		-0.175	-0.177
		(-0.91)	(-0.92)
Dummy 1998		-0.236	-0.244
		(-1.20)	(-1.23)
Dummy 1999		1.267*	1.291*
		(2.51)	(2.55)
Dummy 2000		-5.900	-6.089
		(-0.00)	.
Dummy 2001		1.688**	1.710**
		(2.70)	(2.73)
Dummy 2002		2.168***	2.195***
		(3.56)	(3.60)
Age			-0.246
			(-0.44)
Squared age			0.0022
			(0.44)
N	3,608	3,608	3608
Ind	1,205	1,205	1,205
aic	496	476	479
rho	0.000	0.000	0.000

t-statistics in parentheses

constant (all three columns) and control for gender (column 3) included but not displayed

* $p<.05$, ** $p<.01$, *** $p<.001$

high as .59. This means that the probability of being fired for those whose terms of notice was lowered by the the introduction of the ‘Flexwet’ was increased by .59. This is a major marginal effect. The t-statistics are however still low. The low number of positive firing observations make the inference unreliable. Bertrand, Duflo and Mullainathan (2002) furthermore indicate that standard errors might be underestimated because of serial correlation problems in the DID-procedure.

7 Conclusion and suggestions for further research

This paper performs a DiD-estimation of the propensity to be fired on a policy change in the length of the term of notice for workers with a permanent contract. The relevant policy change is the introduction of the ‘Flexwet’ in 1999 in the Netherlands that altered the legal terms of notice for tenured employees. An employee’s age no longer determined his or her ToN and the number of ToN options was lowered from 26 to four. To determine who is in the treated and who is in the control group, a comparison is made of the terms of notice that would have been granted last year under the old legislation and the terms of notice that would have been granted this year under the new legislation when someone would have been employed. The treated group consists of individuals new-ToN is lower than their old-ToN. The control group consists of (former) employees that face an equal or higher new-ToN. The Dutch Socio-Economic Panel dataset was used to estimate a random effects probit regression on a dummy representing whether someone who had a permanent labor contract last year was fired during this year. The analysis was restricted to individuals that were over 49 and under 65 years of age as general time-trends in firing probabilities are most likely to be identical for these employees and as younger employees all faced a higher ToN under the new legislation.

The DiD-parameter β_3, the coefficient of the interaction term in a regression of outcomes on treatment and time dummies, is positive and significant in different specifications of the model. Workers who faced lower terms of notice than before were apparently sooner fired than comparable peers in the past. This can be interpreted as evidence for the positive causal effect of lower firing costs on the probability to be fired.

First, the number of positive firing observations in my sample is very low, 49 to be precise. The test-statistic on the basis of which we accept the null-hypothesis that lower terms of notice lead to more firings is hence unreliable. Further research on larger datasets should be conducted in order to give an conclusive answer to the research question. Second, the causal relation between employment protection and hiring decisions of employers deserves attention. Even if lower firing costs leads to more firings, the negative effects of such a policy change in employment protection could be mitigated by an increase in hirings.

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