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Returns to university education

Evidence from an institutional reform

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Abstract in English

In 1982 duration of university education in the Netherlands decreased from five to four years. This institutional reform is exploited for estimating the causal effect of one year of university education on wages in 1997. Wages of employees who enrolled just before or after the reform are compared using data from the Dutch Wage Structure Survey of 1997. We find that the fifth year of university education increased wages with 7 to 9 percent. This wage differential is found for employees enrolling four years before or after the reform. Confounding factors like time-effects, typical age-effects or ability-bias do not seem to bias the main results. The findings suggest that there is scope for increasing private contributions of students. Moreover, the reform may have harmed total welfare. Alternative policies of sticking to five-year duration and increasing private contributions for higher education could have given a more favourable outcome.

Keywords: private returns to university education, natural experiment,

Abstract in Dutch

Met de invoering van de zogenoemde Tweefasen Structuur in 1982 is de nominale duur van een opleiding in het wetenschappelijk onderwijs verminderd van vijf naar vier jaar. Deze institutionele verandering is in deze studie gebruikt om het causale effect te bepalen van één jaar wetenschappelijk onderwijs op de lonen in 1997. Daarvoor zijn de lonen vergeleken van personen die studeerden net voor en net na de verandering. De gegevens zijn afkomstig uit het Loon Structuur Onderzoek 1997 van het CBS. Het empirisch onderzoek wijst erop dat het vijfde jaar wetenschappelijk onderwijs geleid 7 tot 9 procent meer loon oplevert. Dit loonverschil wordt gevonden door de vier ‘jaargangen’ van voor de hervorming te vergelijken met de vier ‘jaargangen’ van na de hervorming. De resultaten lijken niet te worden vertekend door mogelijk verstorende factoren als tijdseffecten, leeftijdseffecten of selectieve instroom van studenten. De bevindingen suggereren dat er ruimte is voor een verhoging van private bijdragen van studenten. Daarnaast lijkt de hervorming nadelig te zijn geweest voor de Nederlandse welvaart. Het handhaven van de vijfjarige nominale duur en het verhogen van private bijdragen voor hoger onderwijs had mogelijk een betere uitkomst gegeven.

Steekwoorden: privaat rendement van wetenschappelijk onderwijs, natuurlijk experiment

Nederlandse samenvatting beschikbaar op www.cpb.nl.

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Summary

A prominent policy option in Dutch higher education is to increase private contributions in combination with a reform of the student support system. Insight into the private returns to higher education can be important for the current debate as it indicates the scope for changing the level of contributions. This paper estimates the wage effect of one year of university education.

It is well known that wage differences between higher educated and lower educated employees might not reflect the causal effect of schooling. Higher educated employees differ in many respects from lower educated employees and not all differences are observed. Unobserved factors, such as motivation or intelligence, can bias the estimated effects of education on earnings. Exogenous variation in education can solve this problem. This paper uses a major reform, introduced in 1982, which reduced the duration of university education from five to four years. For identifying the causal effect of one year of university education wages of graduates who enrolled four years before and after the reform are compared. Data from the Dutch wage structure survey of 1997 have been used.

The main finding of this paper is that graduates in the five-year regime earn on average 7 to 9 percent more than graduates in the four-year regime. Confounding factors such as time-effects, typical age-effects or ability-bias do not seem to threaten the main results. First, a sudden change in the wages of higher educated employees might bias our results. However, a comparison with wages of graduates of higher professional education shows that only wages of university graduates increased in the relevant years. Second, the discontinuity in wages might be the result of ‘typical age-effects’ of university graduates. However, we do not find sudden increases of wages of university graduates around the age of 35 in earlier years (1979 and 1985). We only find a discontinuity in wages of university graduates in 1997. Third, if the four-year regime attracted more low ability students the discontinuity in wages might be caused by differences in ability. If ability bias would be important we only expect to find a wage difference in the lowest parts of the wage distribution. However, the wage difference between the two regimes is found over most parts of the wage distribution.

The results of this paper indicate that a fifth year of university education can be very profitable for participants. The high private returns suggest that there is scope for private contributions. The analysis of this paper can also be seen as a partial evaluation of the reform of 1982. We find that the reform decreased earnings opportunities for all graduates from the four-year studies. This suggests that the reform of 1982 was harmful for total welfare as it is unlikely that the government savings from the reduction of the duration of education outweigh the earnings reduction. An alternative policy of sticking to the five-year duration and increasing private contributions for higher education might have given a more favourable outcome.

1 Introduction

Dutch higher education is undergoing some major restructuring. Recently the bachelor/ master structure has been introduced to facilitate international mobility of students in the European Union. The discussion centres on the length of a Masters education and the issue of public or private financing. Moreover, there is discussion about the current system of the financial student support and the central regulation of college fees.

This report analyses the causal effect of university education on wages. It is well know that wage differences between higher educated and lower educated employees might not reflect the causal effect of schooling because of endogeneity problems. Higher educated employees differ in many respects from lower educated employees and not all differences are observed. Unobserved factors, such as motivation or intelligence, can bias the estimated effect of education on earnings. Exogenous variation in education can solve the endogeneity problem. Institutional changes might be a source of exogenous variation in education because all participants, irrespective of motivation or intelligence, are confronted with these changes. This paper uses a major reform in Dutch university education, introduced in 1982. The reform reduced the length of university education for all students from five to four years. We exploit this reform as a natural experiment for estimating the causal effect of one year of university education on wages in 1997. For this, we compare wages of employees who studied just before or after the reform.

This report aims to contribute to current policy discussions. The evidence on the returns to an extra year of university education is relevant for the current discussion on the length of a Masters Education. Second, the results of this paper can be important for students deciding on enrolling in higher education. What are the benefits compared to the costs. The results can also be relevant for financing issues of higher education if the private returns could be compared with the social returns. In addition, this paper contributes to the economic literature on the private returns to education. First, we estimate the returns for a specific group: university graduates. Second, we study the effects of a reduction of the length of education whereas the literature concerns the effects of increasing the length of education. Does reducing the length of education give similar results as the effects of increasing the length of education from the literature? Third, estimates for the Netherlands that take endogeneity into account are scarce.[1] This study provides new evidence for the Netherlands.

For identifying the causal effect of one year of university education we compare wages of graduates who enrolled four years before and after the reform. As we do not directly observe whether a graduate has been treated with the fifth year of university education we use age to distinguish between graduates from the four and five-year regime. We find a discontinuity in the wage level of university graduates, which coincides with the reduction of duration of university education. On average graduates assigned to the five-year regime earn 7 to 9 percent

[1] Two Dutch studies take endogeneity into account: Levin and Plug (1999), Kalwij (2000)

more than graduates in the four-year regime. This wage difference is robust for different specifications and different years around the reform. We consider three potential confounding factors: time-effects, 'typical age-effects' or ability-bias. First, a sudden change in wages of higher educated employees might bias our results. We checked whether these time-effects played a role in a difference in difference analysis in which graduates from higher professional education are used as a counterfactual. Higher vocational education is the second type of higher education in the Netherlands and it's duration of four years did not change in the relevant period. This analysis shows that only wages of university graduates increased in the relevant years. Second, the discontinuity in wages might be the result of 'typical age-effects' of university graduates. Do university graduates typically receive a sudden wage increase at the age of 35? If this is the case our findings may not be caused by the extra year of education. However, we do not find this 'typical age-effect' on wages of university graduates in 1979 and 1985. We only find a discontinuity in wages of university graduates in 1997. Third, if the four-year regime attracted more low ability students the discontinuity in wages might be caused by the difference in ability. Enrolment actually increased after the introduction of the new regime but this might be the result of demographic factors because the number of graduates from pre-university education also increased. The lowest estimates of the wage difference drop some 1.5 percent point if we remove 10 percent of the lowest wages from the relevant age-groups of the new regime. This implies the strong assumption that the introduction of the new regime induced these graduates to enter university. A comparison of trends in enrolment and graduation in pre-university education suggest that this simulation might overrate the increase in enrolment resulting from the reform. Moreover, we performed a quantile regression on the reduced sample to check whether the wage difference only occurs in the lowest parts of the wage distribution. The analysis shows substantial wage differences between graduates from the two regimes over most parts of the wage distribution. We conclude that all three confounding factors do not seriously bias our results.

2 Background of the reform

The discussion on the structure of Dutch higher education dates back to the end of the Second World War (the state committee Reinink (1949)). Many commissions spoke out concerns about overeducation. But until the early eighties not much changed and the standard duration of university education was five years. In August 1982 duration changed to four years. The main reason for this reform was a financial one and had nothing to do with overeducation. Increasing enrolment had put pressure on the budget for higher education. The Dutch Minister of education and science mr. Pais solved this problem by shortening duration of studies. In this way costs were reduced without restricting enrolment. The reduction of duration of studies with one year was not implemented by skipping the fifth year of studies. In general studies reduced many parts of the curriculum, ‘key elements’ survived and ‘voluntarily’ elements were brought back in duration.

The reform intended to change university education in a so-called two-stage-structure. The first stage consisted of the standard university degree of fours years that included a propedeutic exam after one year. The second stage was meant as preparation for scientific research. The duration of this stage should be no more than 2 years. This second part of the reform was in fact never implemented. In 1984 the new Minister Deetman introduced a new system for PhD-students, the so-called AIO-system with duration of four years.

3 Review of literature and identification strategy

The traditional way of identifying the wage effect of an extra year of schooling is to estimate a Mincer type of wage equation. In this equation (log) wages is used as the dependent variable and regressed on the number of schooling years (and several covariates). It can be shown that the coefficient of the schooling variable represents the private return to one year of education. With this approach we find the private return of a year of university education in the Netherlands to be nearly 11 percent (see appendix). It is well known that the estimated coefficient for the schooling variable might not reflect the causal effect of education on earnings because of endogeneity problems. A huge literature deals with this problem and describes new approaches such as natural experiments or twin studies (see for instance, Card, 1999 or Ashenfelter et al. (1999)). Several studies used institutional changes of the education system as a source of exogenous variation in education. The seminal study in this line of research is Angrist & Krueger (1991). They use the fact that the school year for all 6-year olds in the US starts on the same date together with compulsory schooling laws binding students to schools up till the age of 16. As a result, students born early in the year on average follow less education than students born later in the year. They use this variation in schooling for identifying the causal effect of education on earnings. One year of education is estimated to give 6 to 8 percent higher wages. Harmon & Walker (1995) use the raising of the minimum school-leaving age in the United Kingdom (which occurred in 1947 and 1973) as a source of exogenous variation in education. They find an estimate of schooling return of more than 15 percent. The same approach is used by Levin & Plug (1999) and Vieira (1999). Meghir & Palme (2001) use a major education reform in Sweden, which was implemented in the 1960s. Before the country-wide implementation a proportion of municipalities started with the new school system. They find that the reform increased educational attainment of individuals from poorer backgrounds and that the returns to schooling depend on the ability of individuals. The returns for high ability individuals are estimated to be 7.5 percent. Low ability individual have lower returns. Aakvik, Salvanes & Vaage (2003) exploit a major reform in the comprehensive school system in Norway in the 1960s. As in Sweden, there is a staged implementation. They find that the returns to schooling are strongly nonlinear, depending on the type of education.

Identification Strategy

We use the university reform of 1982 for identifying the wage effect of one year of university education. The reform creates a discontinuity in educational attainment for graduates who started studying in higher education before or after the reform. All graduates who started higher education in august 1982 or later entered the four-year regime. Graduates who started before august 1982 entered the five year regime. They had to follow the fifth year of university education. This type of discontinuity can be exploited with a regression-discontinuity design (RD design). The RD design exploits a known discontinuity in the treatment assignment to identify the treatment effect (the fifth year of higher education). There are two types of designs. In a sharp design the discontinuity completely determines the treatment. In case of a fuzzy design the discontinuity determines the probability of being treated. Literature on these methods dates back to Campbell (1969), for a recent application see Leuven and Oosterbeek (2004, forthcoming). This paper would be an application of a sharp design if we could observe whether a graduate has been treated with the fifth year of education. However, this is not the case. We use age to distinguish between graduates from the four and five-year regime. Our design translates to the fuzzy type because age does not completely determine the treatment. In Dutch education the way to university runs through pre-university education, which is a six-year type of secondary education. Students who follow the fastest way can enter university at the age of 18. As many university students in Dutch education have some delay age is not a perfect predictor of the university regime taken by the graduate. All graduates born on or after the first of August 1963 were not treated with the extra year. Graduates born earlier had a probability of being treated with the fifth year. The probability becomes larger when the distance between the date of birth and the first of august 1963 increases. The probability of receiving the extra year of university education (t_i) depends on the date of birth (d_i) with a known discontinuity at point $\bar{d}$ (in our case August 1 1963).

$$\Pr(t_i) = f(d_i, 1\{d_i \leq \bar{d}\} \tag{3.1}$$

The outcome is the wage increase from the fifth year of university education and can be described as follows

$$\ln W_i = \alpha + \beta t_i \tag{3.2}$$

where α is the wage without the fifth year of university education, and β is the wage change due to the fifth year of university education.

Our empirical strategy for identifying the wage effect of the fifth year of university education comes down to comparing the wages of employees born around the cut-off date related to the introduction of the reform. This comparison will give an unbiased estimate of the treatment effect if there is no reason to believe that persons close to $\bar{d}$ are different. The major identifying assumption is that there are no other discontinuities around $\bar{d}$. The treatment effect in a fuzzy regression discontinuity design has been estimated with an instrumental variable approach (Van der Klaauw (2002), Leuven and Oosterbeek (2004)). In the first stage equation the treatment is regressed on the discontinuity. Our application differs as we do not observe whether the graduate has been treated with the fifth year of university education (we can not estimate the first stage equation). We therefore estimate a reduced form equation

$$\ln W_i = \alpha + \beta T + \delta X \tag{3.3}$$

in which T is the treatment variable and X is a set of control variables. If students did not have any delay in previous education and we could observe the month of birth we would define T=1 if the employee is born before August 1 1963, and T=0 is born after August 1 1963. We deviate from this for two reasons. First, we only observe the year of birth. Second, general statistics suggest that at least half of the students have a delay of one year in pre university education. Moreover, after enrolment, 30 to 50 percent of students drop out or switches to another study. As a consequence, we expect that the majority of students born in 1963 entered the new regime and that students born in 1962 where allocated over both regimes (see appendix for more details). We therefore excluded graduates born in 1962 from the analysis and define

T=1 if the employee is born before 1962;
T=0 if the employee is born after 1962.

We estimated this reduced form equation using four so-called discontinuity samples. Discontinuity sample ±1 (abbreviated DS±1) consist of the treatment group of graduates born in the first year before 1962 and the control group of employees born in the first year after 1962. Discontinuity sample ±4 (abbreviated DS±4) consists of the treatment group of graduates born in the first four years before 1962 and the control group of graduates born in the first four years after 1962.

4 Data

The data we use come from the so-called Wage Structure Survey (Loon Structuur Onderzoek (LSO)) held by Statistics Netherlands (CBS). Data on wages are obtained through the annual survey on employment and wages among firms (Enquête naar Werkgelegenheid en Lonen) and partly through Administrations on insured people (Verzekerden Administratie (VZA)). This means that all information on wages comes from administrative sources (firms or administrations on insured people). This dataset also contains information on gender, age and job characteristics. Data on education are obtained from the annual labour force survey (Enquête Beroepsbevolking (EBB)) and matched with the wage data. This matched dataset is called the Wage Structure Survey.

The main data we use come from the survey of 1997. The total sample consists of nearly 120.000 employees. We use sub samples of graduates from university education and higher professional education born between 1958 and 1966. Table 1 shows statistics for graduates assigned to the old and new regime. The dependent variable in the analysis is gross hourly wage in 1997. As independent variables we use: gender, education (4-digit coding according to the Standard Education Coding), age (measured in years in December of the year of the survey), age squared, experience (the difference between age and year of graduation), experience squared and potential experience (the difference between age and the nominal duration of education). The data on the year of graduation seem to contain measurement errors. In the analysis we corrected the experience variable based on the year of graduation by using a maximum experience level. All employees with experience above this level where assigned a maximum based on the nominal duration of the study. Moreover, in the analysis we use potential experience which is the standard approach in estimating the Mincerian wage equation. We also control for firm characteristics: type of industry measured at 2-digit level (SBI-code) and firm size measured as the (log) number of employees. We restrict the analysis to employees with a Dutch nationality. We do not have information whether non-Dutch employees were educated in the Netherlands or elsewhere. Moreover, we exclude self-employed and employees with wages below the minimum wage for 23 year-olds.

Table 4.1 Sample statistics of graduates in old and new regime in 1997

University	Treatment group Graduates born in 1958-1961 (n=1158)				Control group Graduates born in 1963-1966 (n=1192)			
	mean	st. dev.	min	max	mean	st. dev.	min	max
(Ln)wage	3.78	0.34	2.66	5.18	3.61	0.32	2.44	4.82
Female (%)	32.9	47.00			40.9	49.2		
Graduation date	1987	3.6	1965	1998	1990	2.8	1953	1998
Experience	10.5	3.3	0	15	7.3	2.4	0	11
Potential experience	13.4	1.1	12	15	9.5	1.1	8	11
Age	37.4	1.1	36	39	32.5	1.1	31	34
Firm size	5613	11086	1	56804	4770	10274	1	64974
Higher professional	Graduates born in 1958-1961 (n=2840)				Graduates born in 1963-1966 (n=2837)			
	mean	st. dev.	min	max	mean	st. dev.	min	max
(Ln)wage	3.58	0.32	2.40	5.21	3.47	0.28	2.42	4.95
Female (%)	44.0	49.6			45.3	49.8		
Graduation date	1985	5.1	1920	1998	1989	3.5	1963	1998
Experience	12.3	4.4	0	17	8.6	2.8	0	12
Potential experience	15.5	1.1	14	17	10.5	1.1	9	12
Age	37.5	1.1	36	39	32.5	1.1	31	34
Firm size	4031	10094	1	64974	4151	10515	1	64974

We also use data from the Dutch Wage Structure surveys of 1979 and 1985. The sample statistics of university graduates from these years are shown in table 4.2.

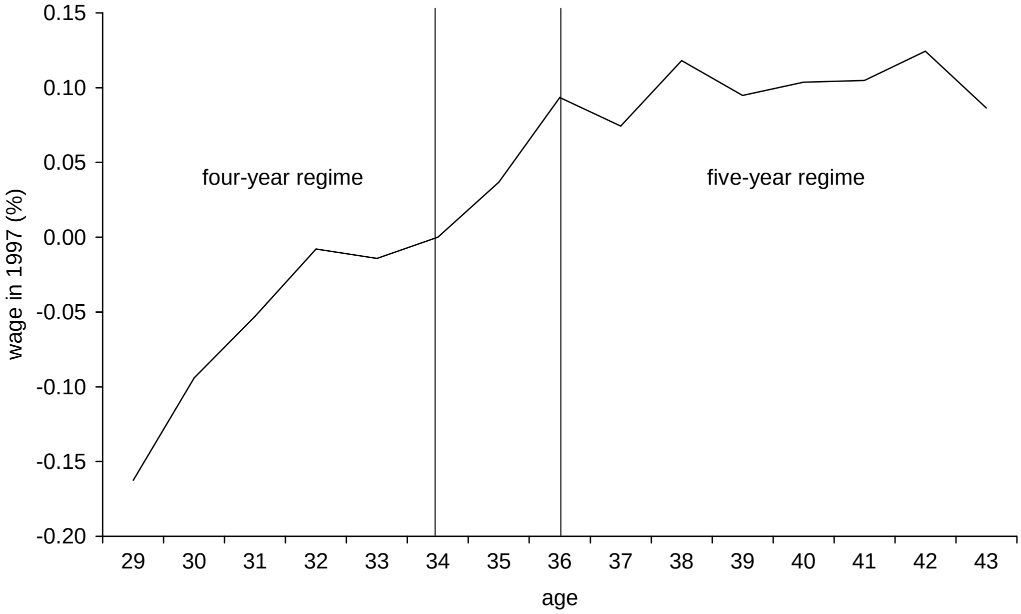
Table 4.2 Sample statistics for 1979 and 1985

1979	Treatment group Graduates 36-39 years (n=90)				Control group Graduates 31-34 years (n=141)			
	mean	st. dev.	min	max	mean	st. dev.	min	max
(Ln)wage	3.51	0.17	3.03	4.07	3.32	0.22	2.52	4.06
Female (%)	6.6	25			9.9	30.0		
Age	37.5	1.1	36	39	32.4	1.1	31	34
1985	Graduates 36-39 years (n=53)				Graduates 31-34 years (n=47)			
	mean	st. dev.	min	max	mean	st. dev.	min	max
(Ln)wage	3.46	0.24	2.69	4.29	3.33	0.27	2.80	3.87
Female (%)	13.2	34.2			29.8	46.2		
Age	37.5	1.2	36	39	32.9	1.1	31	34

5 Results

For a first impression of the wage effects of the reform we regress wages on age controlling for experience, experience squared, type of education and gender. University graduates aged 34 are the first year-cohort in the four-year regime. We take this group as a reference in figure 5.1.

Figure 5.1 Earnings of university graduates by age



Wages increase sharply for university graduates in the four-year regime until the age of 32. Then they stabilise for two years. This pattern is typical for new entrants on the labour market. In the first years wages increase more than the average effect of experience (and experience-squared) from our specification. After a few years wages come in line with the average effect of experience on wages. Between the age of 34 and 36 there is again a sharp rise in earnings. The timing of this increase coincides with the reform in university education. Wages of 35-year olds lie between the two regimes, conform expectations based on general statistics. After the age of 36 wages are more stable and move around the 10 percent line. This pattern suggests a clear wage difference between graduates in the four- and five-year regime.

Next, we more closely examine wage differences between age groups around the regime change. We estimate wage differences using four discontinuity samples around the regime change. In table 5.1 we present estimates using four different specifications. The first column controls for type of education, gender and experience measured by year of graduation. In the second column we use potential experience, which is the standard experience variable from the Mincer equation. The third and fourth column also control for firm characteristics and labour market conditions in the year of graduation. In column (3) we use type of industry and firm size. In column (4) we control for labour market conditions in 1987 and 1988. In these two years a double flow of university graduates (from both regimes) entered the labour market.[2] We estimated the model separately for each discontinuity sample. Standard errors are given in parentheses.

Table 5.1 Wage difference between graduates born before and after 1962

	(1)	(2)	(3)	(4)
DS+/-1	0.086			
	(0.027)			
N	604			
DS+/-2	0.078	0.116	0.096	0.100
	(0.020)	(0.040)	(0.039)	(0.038)
N	1221	1221	1221	1221
DS+/-3	0.076	0.095	0.074	0.077
	(0.017)	(0.030)	(0.030)	(0.030)
N	1813	1813	1813	1813
DS+/-4	0.089	0.088	0.070	0.072
	(0.015)	(0.025)	(0.025)	(0.025)
N	2350	2350	2350	2350
Controls				
Experience	yes	no	no	no
Potential experience	no	yes	yes	yes
Firm characteristics	no	no	yes	yes
Dummies 87/88	no	no	no	yes

Notes: Coefficients shown in columns 1 to 4 are coefficient of a dummy in an OLS regression, standard errors in parentheses. Individuals born in 1962 are excluded from the estimations. All models control for gender, education (4-digit level). Controls denoted by 'experience' are experience and experience squared based on year of graduation. Controls denoted by 'potential experience' are potential experience and potential experience squared based on age. Firm characteristics are type of industry (2-digit level) and (log) number of employees. Dummies 87/88 are dummies for graduation years 1987 and 1988.

The estimates in column (1), using basic controls and experience based on the year of graduation, indicate that employees who graduated in the old regime earn some 7.6 to 8.9 percent more that employees who enrolled after the reform. The wage difference increases if we include potential experience (column (2)) instead of measured experience. In the four year discontinuity sample we find almost the same wage difference. Including firm characteristics lowers the estimates (column (3)). Adding dummies for the high outflow years 1987 and 1988 gives slightly higher estimates (column (4)). In general, unemployment was higher in the first half of the eighties when graduates from the old regime entered the labour market. These unfavourable starting conditions could have a downward bias on the estimated wage difference.[3] All specifications indicate wage differences between graduates from the four and five-year regime between 7 and 9 percent. The smallest estimate comes from the largest

[2] In 1985 there were 20,100 university graduates, in 1986 29,500, 1987 25,000, 1988 19,400 and in 1989 20,100.

[3] We constructed a time series of the unemployment rate for those aged 25-34 years based on the Dutch Labour Force Surveys 1970-1997 and included this variable in the analysis. However, the results seemed not very plausible. Higher unemployment at the time of entering the labour market increased wages.

discontinuity sample. Moving further from the regime change offers more ground for confounding factors. Therefore, it seems safe to conclude that the extra year of university education caused a wage increase between 7 and 9 percent.

In table 5.1 we leave out graduates born in 1962. This decision is based on general statistics on drop out and the delay of students in pre-university education. However, it might be argued that employees born in 1963 could also be divided over both regimes and should be left out from the analysis. We therefore repeated the analysis after excluding graduates born in 1963 and adding graduates born in 1967 to the four-year discontinuity sample. A drawback of this approach is that we move further from the cut-off date and to left and steeper part of the earnings function in figure 5.1. With this approach the range of findings increases and the lowest estimates of the wage difference are close to 6 percent (appendix A.3).

6 Confounding factors

Several factors might bias the results in the previous section. We investigate the bias of three factors: time effects around the reform, typical age-effects for university graduates and ability-bias

6.1 Time effects

The major identifying assumption in the previous analysis is that there are no other discontinuities around the reform. In other words, the wage difference between graduates from the old regime and graduates from the new regime can only be related to the reform and not to other differences in the years just before and after the regime change. For instance, if there was a sudden change in wages for higher educated employees at the time of the reform we could falsely ascribe this to the reform. In order to control for time effects we compare wage changes of university graduates with wage changes of graduates of higher professional education of the same age. Dutch higher education consists of two levels: university education and higher professional education. The latter has duration of four years and this did not change in the period of the university reform. If time effects occurred we expect this to be reflected in the wages of graduates of higher professional education. First, we repeated the previous analysis for graduates from higher professional education of the same age. Next, we performed a difference in difference analysis. This analysis compares the wage difference between university graduates from the two regimes with the wage difference between graduates from higher professional education from the two age-groups. The treatment effect (β) of one year of university education then can be found as:

$$\beta = (\ln W^{u}_{58-61} - \ln W^{u}_{63-66}) - (\ln W^{hp}_{58-61} - \ln W^{hp}_{63-66}) \qquad (6.1)$$

with $\ln W^{u}_{58-61}$ is the wage of university graduates from the old regime (born in 1958-1961) and $\ln W^{hp}_{63-66}$ is wages of graduates from higher professional education from the age-groups parallel to the new university regime (born in 1963-1966). In table 6.1 we present estimation results for the most extended specification (column 4 in table 5.1). The first column repeats the results for university graduates. The second column shows results for graduates from higher professional education. The third column shows results of the difference in difference analysis.

Table 6.1	Wage difference between graduates from university and higher professional education born before and after 1962		
	University (1)	Higher Professional (2)	(1) – (2)
DS+/-2	0.100	– 0.036	0.136
	(0.038)	(0.031)	(0.048)
N	1221	2888	4109
DS+/-3	0.077	0.004	0.073
	(0.030)	(0.021)	(0.035)
N	1813	4316	6129
DS+/-4	0.072	0.007	0.065
	(0.025)	(0.017)	(0.029)
N	2350	5677	8027

Notes: Controls used are gender, type of education, potential experience, potential experience squared, type of industry, (log) number of employees in firm, year dummies if graduated in 1987 and 1988.

Time-effects do not seem to bias previous estimates. We do not find wage differences for graduates from higher professional education (column 2). Therefore, the difference in difference estimates, which captures the difference between column 1 and column 2, are comparable to the estimates in column 1. The estimates for the 2-year discontinuity sample are even higher.

6.2 Typical age-effects for university graduates

If the age-experience profiles of university graduates typically have a discontinuity at the age of 35 we might falsely relate this to the reform. To check whether this is the case we repeat the analysis of section 5 on data from the Wage Structure Surveys of 1979 and 1985. The first column of table 6.2 gives the results for 1979, the second column for 1985 and the third column for 1997. The surveys of 1979 and 1985 do not have a detailed education variable. In table 6.2 we control for gender, potential experience, potential experience-squared and education at two-digit level.

Table 6.2	Wage difference between university graduates older and younger than 35 years in 1979, 1985 and 1997		
	1979	1985	1997
DS+/-1	– 0.017	0.021	0.124
	(0.050)	(0.083)	(0.025)
N	54	35	665
DS+/-2	– 0.257	– 0.159	0.116
	(0.115)	(0.186)	(0.040)
N	107	57	1356
DS+/-3	– 0.173	– 0.091	0.091
	(0.082)	(0.128)	(0.030)
	173	78	2013
DS+/-4	– 0.033	– 0.075	0.080
	(0.065)	(0.108)	(0.026)
N	231	100	2607

Note: Controls used are gender, type of education (2-digit), age and age squared.

Both for 1979 and 1985 we do not find a typical discontinuity at the age of 35. In both years older graduates on average earn less than younger graduates. This clearly deviates from the findings for 1997.

6.3 Selection effects

Despite the fact that we focus on an institutional reform to solve the endogeneity problem there still might be self-selection of students. The reduction in duration of university education might have attracted more or different students because of lower investment costs and lower drop-out risk. In fact, enrolment in university education increased after the reform. In the years before the reform approximately 23,500 students yearly enrolled. After the reform this went up to nearly 27,000 students.[4] However, a substantial part of this increase might have a demographic cause. The number of graduates from pre-university education, which gives direct access to university, increased from approximately 29,600 in 1981 to 33,000 in 1985.[5] If the 4-year regime attracted more low-ability students the results in section 5 might suffer from ability-bias. To check whether selection bias confounds our results we excluded graduates from the new regime with the lowest wages. We left out 10 percent of graduates of each separate birth year of the new regime, except 1963. This exclusion is roughly based on the figures on enrolment and graduation in pre-university education. We did not leave out low earners born in 1963 because the increase of graduates from pre-university education is larger than the increase in enrolment.

[4] The exact figures on freshmen enrolling are: 23,513 in 1981, 24,243 in 1982, 26,816 in 1983, 26,845 in 1984 and 26,707 in 1985.

[5] The exact figures are: 29,600 in 1981, 30,800 in 1982, 31,200 in 1983, 31,700 in 1984 and 33,100 in 1985.

This exclusion of low wage graduates probably overrates the change in enrolment caused by the introduction of the new four-year regime. On this reduced sample we first estimated the wage effect of the extra year of university education. Next, we performed a quantile regression to investigate whether the wage difference only occurs in the lowest part of the wage distribution or can also be found in the upper parts of the wage distribution. We expect no wage difference between the two regimes in the higher parts of the wage distribution if our previous findings are the result of ability bias. In a quantile regression the coefficients are estimated for different quantiles of the wage distribution (see Buchinsky, 1994). The quantile regression model can be written as:

$$\ln W_i = X_i \beta_\theta + u_{\theta i} \text{ with } \mathrm{Quant}_\theta(\ln W_i | X_i) = X_i \beta_\theta \tag{6.2}$$

where X_i is the vector of exogenous variables and β_θ is the vector of parameters. Quant $\theta(\ln W|X)$ denotes the θ th conditional quantile of lnW given X. The θ th regression quantile, $0 < \theta < 1$, is defined as a solution to the problem:

$$\min_{\beta \varepsilon R^t} \{ \sum_{i:y_i \geq X_i \beta} \theta |\ln W_i - X_i \beta_\theta| + \sum_{i:y_i > X_i \beta} (i-\theta) |\ln W_i - X_i \beta_\theta| \} \tag{6.3}$$

By variation of θ, any quantile of the conditional distribution can be obtained.

Table 6.3 shows estimates of the wage difference for the reduced sample. Column 1 shows the standard regression results for the reduced sample, column 2 shows results for the 10-th quantile, the lower part of the wage distribution and column 6 for the 90-th quantile, the upper part of the wage distribution.

Table 6.3 Wage differences between university graduates born before and after 1962 for various quantiles of the wage distribution after excluding low wage of new regime

	Standard	10	25	50	75	90	N
DS+/-2	0.131	0.125	0.149	0.127	0.142	0.139	1190
	(0.037)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	
DS+/-3	0.077	0.110	0.091	0.051	0.056	0.033	1754
	(0.028)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	
DS+/-4	0.055	0.104	0.061	0.061	0.057	0.020	2356
	(0.024)	(0.028)	(0.003)	(0.000)	(0.008)	(0.010)	

Notes: 10 % lowest wages of those born in 1964, 1965 and 1996 are excluded. Controls used are gender, type of education, potential experience, potential experience squared, type of industry, (log) number of employees in firm, dummies if graduated in 1987 or 1988.

Removing low-wage graduates from the new regime lowers the estimated wage differences in the three- and four-year sample to 7.7 and 5.5 percent. This can be seen as lower bound results given the fact that we probably removed too many graduates. In addition, our assumption that the reform induced these graduates to enter university may not hold. The estimated wage difference for the two-year sample increases. Removing graduates with the lowest wages also changes other coefficients in the model. The model for the two-year sample is sensitive for this because of the smaller sample and the smaller period for estimating the coefficient on experience. As the results for the three and four-year samples are less sensitive for this they can be better compared with previous estimates. The results of the quantile regression indicate that the extra year of university education pays more in the lowest part of the wage distribution. However, substantial wage differences are found over most parts of the wage distribution. For instance, we find a wage difference of more than 5.5 percent for the 75-th percentile of the wage distribution. We think this excludes ability bias as the main reason for the wage differences found in the previous section. If we accept the strong assumption that the introduction of the new regime induced the low-wage graduates to enter university we find a wage difference of 5.5 percent.

7 Conclusions and discussion

The reduction of duration of university education coincides with a discontinuity in the wage level of university graduates. Graduates in the five-year regime earn on average 7 to 9 percent more than graduates in the four-year regime. Confounding factors such as time-effects, typical age-effects or ability-bias do not threaten the main results. First, a sudden change in the wages of higher educated employees might bias our results. However, a comparison with wages of graduates of higher professional education shows that only wages of university graduates increased in the relevant years. Second, the discontinuity in wages might be the result of 'typical age-effects' of university graduates. However, we do not find sudden increases of wages of university graduates around the age of 35 in earlier years (1979 and 1985). We only find a discontinuity in wages of university graduates in 1997. Third, if the four-year regime attracted more low ability students the discontinuity in wages might be caused by differences in ability. Enrolment increased after the introduction of the new regime but this might be the result of demographic factors as the number of graduates from pre-university education also increased. The lowest estimate of the wage difference drops to 5.5 percent if we remove 10 percent of the lowest wages from the relevant age-groups of the new regime. This implies the strong assumption that the introduction of the new regime induced these graduates to enter university. Moreover, the wage difference between the two regimes is found over most parts of the wage distribution. If ability bias would be important we only expect to find a wage difference in the lowest parts of the wage distribution.

Policy implications

This paper shows that a fifth year of university education can be very profitable for participants. In the current policy context this might imply that a second year of a masters education can yield high returns. At the moment government subsidies cover about 88 percent of costs of higher education. The high private returns suggest that there is scope for private contributions. An income contingent loan system can increase private contributions without loss of accessibility of higher education (see CPB, 2003).

The analysis of this paper can also be seen as a partial evaluation of the reform of 1982. We find that the reform decreased earnings opportunities for all graduates from the four-year studies. This reduction of earnings and productivity was not intended and should be weighted against government savings from the reduction of duration. Jacobs and Canton (2003) calculate that increasing private contributions from the current 12 percent to 37.5 percent corresponds to an average of 3.5 percent of lifetime income of Dutch university graduates. The government savings of one year higher education are about this size and clearly do not outweigh the earnings reduction. We conclude that the reduction of 1982 was harmful for total welfare. It is likely that an alternative policy of sticking to the five-year duration and increasing private contributions for higher education would have given a more favourable outcome.

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Appendix 1: OLS-estimation of private returns to education

The traditional way of identifying the private returns to education is estimating a Mincer type of wage equation:

$$\ln W_i = \alpha + \beta S + \gamma E + \lambda E^2 + \delta X_i + \varepsilon_i \quad (7.1)$$

with W is wages, S is years of education, E is experience, X is a vector of covariates and ε_i are unobserved factors of individual i. Table A1 presents estimations results. Coefficients smaller than 0.10 can be interpreted as percent differences. Coefficients larger than 0.10 can be translated in percent difference with exp(C)-1, with C is coefficient. The models in table A1 control for gender, age and age-squared. The average yearly returns can be obtained by dividing with the duration of education. The sample is restricted to employees of 37 to 55 year which only includes the old regime of university education.

Table A.1 OLS-estimates of private returns to education in 1997

	Total	Male	Female
Lower education	– 0.195	– 0.203	– 0.196
	(0.007)	(0.010)	(0.011)
Middle general (mavo3)	– 0.006	– 0.003	– 0.001
	(0.031)	(0.050)	(0.040)
Lower vocational (vbo)	– 0.113	– 0.117	– 0.125
	(0.007)	(0.009)	(0.010)
Middle general (mavo4)	0.000	0.000	0.000
Intermediate vocational (mbo)	0.101	0.084	0.112
	(0.006)	(0.009)	(0.009)
Higher general (havo)	0.092	0.078	0.102
	(0.011)	(0.016)	(0.015)
Higher professional (hbo)	0.371	0.369	0.355
	(0.007)	(0.009)	(0.009)
Pre university (vwo)	0.249	0.264	0.195
	(0.013)	(0.016)	(0.021)
University (wo)	0.600	0.592	0.599
	(0.008)	(0.010)	(0.012)
N	63471	40819	22652

Notes: Controls used are age and age squared. Sample consists of employees of 37 to 55 year-olds in 1997. Standard errors in parentheses.

In 1997 male graduates from university earned on average 80.8 percent more than graduates from middle general education (exp(0.592)-1). Graduates of pre university education earned on average 28.3 percent more than graduates from middle general education. The returns to a year of university education can be calculated by the difference between university and pre university graduates and dividing with the nominal duration. This gives a return to a nominal year of university education of: 10.8 percent for all, 10.5 percent for men and 12 percent for women.

Appendix 2: Timing of the cut-off in the analysis

Table A.2 **Freshmen in Dutch universities by age (row percentages)**

Age	18	19	20	21
College-year				
1980-‘81	30	26	12	32
1981-‘82	31	26	11	32
1982-‘83	30	26	12	32
1983-‘84	27	26	13	34

Source: CBS (1981-1985)

Without delay the average age of entering university education is 18 years and 6 months. In table A2 age is measured at December 31 of the college-year. Students who entered without delay could have a maximum age of 19 years and 5 month. Table A.2 shows that a large share of freshmen is 19 years or older. If we assume that 19 year-olds are equally distributed over the year more than half of the students has a delay of one year or more. Moreover, a large share, approximately 30 to 50 percent, of first time enrollers drops out or switches to another study within higher education. The new regime started at August 1, 1982. All students born on or after August 1, 1963 entered the new regime. We expect that also a large share of students born earlier in 1963 entered the new regime due to delay in pre university education or switching to other studies. This implies that the majority of students born in 1963 entered the new regime. As we only observe year of birth in our data we assume that all university graduates born in 1963 entered the new regime. We expect that university graduates born in 1962 are divided over both regimes. Therefore, we excluded these graduates from the analysis. We expect that the majority of university graduates born in 1961 entered the old regime.

Appendix 3: Excluding graduates born in 1962 and including graduates born in 1967

Table A.3 **Wage difference between graduates born before and after 1962**

	(1)	(2)	(3)	(4)
DS+/-1	0.107			
	(0.033)			
N	613			
DS+/-2	0.086	0.154	0.136	0.134
	(0.022)	(0.056)	(0.054)	(0.054)
N	1204	1204	1204	1204
DS+/-3	0.097	0.075	0.060	0.057
	(0.019)	(0.037)	(0.036)	(0.036)
N	1805	1805	1805	1805
DS+/-4	0.109	0.075	0.063	0.062
	(0.017)	(0.030)	(0.029)	(0.029)
N	2380	2380	2380	2380
Controls				
Experience	yes	no	no	no
Potential experience	no	yes	yes	yes
Firm characteristics	no	no	yes	yes
Dummies 87/88	no	no	no	yes

Notes: Graduates born in 1962 and 1963 are excluded. Graduates born in 1967 are included in the four-year sample. Controls used are the same as in table 5.1.