CPB Discussion Paper

No 12

August 2002

The Demand for Higher Education in the Netherlands, 1950-‘99

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[*] We would like to thank Peter Dekker, Hans Vossensteyn, the HRM-department of CPB, the Ministry of the Interior and Kingdom Relations and especially Wout de Bruin for their kind help in the data collection process. The complete data set is available from the authors upon request (e-mail: canton@cpb.nl). Helpful comments from Rudy Douven, Pierre Koning, Edwin Leuven, Hessel Oosterbeek, Marc Pomp, Richard Venniker and seminar participants at the NAKE Day 2001, CPB and University of Athens are appreciated.

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ISBN 90-5833-111-3

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Abstract

This paper investigates the role of economic factors in the university enrollment decision for the post-war period in the Netherlands. We include those factors standing at the heart of the idea that education is an investment. Collecting student enrollment data for eight subject categories results in a large data set, as a cross-section dimension is added to the time-series. The econometric results suggest that students are not responsive to tuition fees, but financial support (the sum of loans and grants), the college premium on future labor market earnings, and the alternative wage are important in the enrollment decision.

JEL-codes: C23, D10, I20

Keywords: higher education, student enrollment, human capital approach, pooled data

1 Introduction

This paper explores the demand for higher education in the Netherlands. University attainment has witnessed a profound expansion after the second world war. In 1950, the inflow of first-year students amounted to about 3% of the 18-year old population cohort. This has risen to 17% in 1999. Perhaps some satiation point has been reached, as the majority of people who completed upper secondary school nowadays decide to take up a higher education program. But it is also claimed that as the Netherlands, like other modern countries, is moving towards a knowledge-based society, university enrollment may expand even further in the time ahead of us.

To gain insight into the factors behind the development of Dutch university enrollment, we study the demand for higher education within the analytical framework provided by human capital theory (*cf.* Mincer, 1974; Becker, 1975). According to this theory, people consider education as an investment in human capital. The investment is attractive when the benefits exceed the costs associated with the training program. The benefits are typically expressed in terms of the wage premium connected with the training program, whereas the costs include tuition fee payments and foregone labor market earnings. In addition, we incorporate unemployment data to proxy for the business cycle effect of human capital investments. This latter effect is due to the opportunity cost hypothesis: it is more attractive to invest in human capital when times get bad, as the opportunity costs of such investments in terms of foregone earnings on the labor market are lower in a recession (*cf.* Bean, 1990; Saint-Paul, 1993). Besides the investment view, education may also have a consumptive value. We therefore also consider per capita income as a potential explanation of the demand for university training. Collecting student enrollment data for eight subject categories over the 1950-'99 period results in a large data set, as a cross-section dimension is added to the time-series. The econometric results suggest that students are not responsive to tuition fees, but financial support, the college premium and the alternative wage are important in the enrollment decision.

Related studies for the Netherlands include Kodde and Ritzen (1984), Huijsman *et al.* (1986), De Jong *et al.* (1990), and Sterken (1995), adopting time-series analysis to study the role of financial variables in the decision to take up a higher education program. There is remarkably little consensus among these studies about the econometric results. The only study reporting a significantly negative impact of tuition fees on student enrollment is Sterken (1995). Using a simple error correction specification, he finds a short-run price elasticity of approximately -0.04, and a long-run price elasticity in the order of -0.5. However, as several potentially important variables (such as financial support to students and the alternative wage) are missing, the regression coefficients may suffer from omitted variables bias.[1] Kodde and Ritzen (1984) and Huijsman *et al.* (1986) find empirical support for consumption and investment motives in the university enrollment decision. The latter study reports income elasticities in the range of 0.53-1.02 for males and in the range of 0.15-1.03 for females; the enrollment elasticity with respect to financial aid enters with a positive sign in the model for first-year male enrollment (with a magnitude between 0.09 and 0.15), but with a negative sign in the model for first-year female enrollment; the enrollment elasticity with respect to foregone earnings is negative in all cases, and ranges from -0.15 to -0.33 for males and from -0.19 to -0.52 for females; and the enrollment elasticity with respect to future earnings is positive in all cases, and ranges from 0.12 to 0.34 for males and from 0.12 to 0.52 for females. In contrast, De Jong *et al.* (1990) find that the economic variables hardly affect the decision to enroll in an academic program. The only significant variable is per capita income. The income elasticity equals 1.20 for males and 2.01 for females. The alternative wage, the salary of a university graduate, financial support and the unemployment rate all appear with regression coefficients insignificantly different from zero and often with the "wrong" sign.

Other studies for the Netherlands are based on micro-data or questionnaires. The apparent advantage of such data is that account can be taken of individual characteristics that might be important for the enrollment decision, such as scholastic aptitude and socio-economic status. Oosterbeek and Webbink (1995) use a micro-data set on cohorts of students who were in the final year of secondary education in 1982 (the 1982 cohort) and in 1991 (the 1991 cohort). They use foregone earnings, future earnings, ability scores in language and mathematics, gender, level of secondary education, educational attainment of the parents and family earnings as explanatory variables for higher education enrollment. Three models are investigated: an investment model, a consumption model and an integrated model. The investment model only includes foregone earnings, future earnings and ability scores in the regression analysis. The consumption model includes all variables except future earnings. And the integrated model uses all explanatory variables. For the 1982 cohort it is found that the integrated model is superior to both the consumption and the investment model. But for the 1991 cohort the pure investment model can no longer be rejected in favor of the other two models. Only the future earnings variable appears with a significant coefficient in the regression model. Although tuition fees are not included in the regression analysis (as there is no cross-sectional variation in tuition fees at any point in time), Oosterbeek and Webbink try and identify the effect of tuition fee changes on higher education enrollment by looking at the budget constraint. The increase in tuition fees from 1982 to 1991 rotates the budget constraint in the "other consumption versus schooling"

[1] This error correction specification is also adopted in McIntosh (2001), studying the demand for post-compulsory education in Germany, the Netherlands, Sweden and England. But also this study may suffer from an omitted variables problem, as variables such as tuition fees, financial support to students, and the alternative wage are not included in the regression analysis.

space. Oosterbeek and Webbink disentangle the effects into an income-effect and a substitution-effect, enabling them to draw inferences on the effects of a tuition fee increase. They conclude that "... the demand for higher education is completely inelastic" (pp. 377).

Berkhout and Van Leeuwen (2000) explore factors behind the decision to enroll in science or engineering from a questionnaire among second-year students in 1997. Abolishment of tuition fees for science and engineering would yield 1,600 additional students in these fields (at the expense of other disciplines). This implies a tuition fee elasticity of approximately -0.073. Likewise, an increase of the monthly grant to students in science and engineering by € 340 is expected to yield an additional 1,150 students, corresponding to an enrollment elasticity of 0.053 (under the assumption that the policy change corresponds to a doubling of the monthly grant to students in science and engineering; this holds approximately). This is a strange result, as it suggests that, say, a € 100 increase in tuition fees combined with a € 100 increase in student grants would still result in a rather substantial reduction in student enrollment, whereas students should be indifferent to such a policy change.

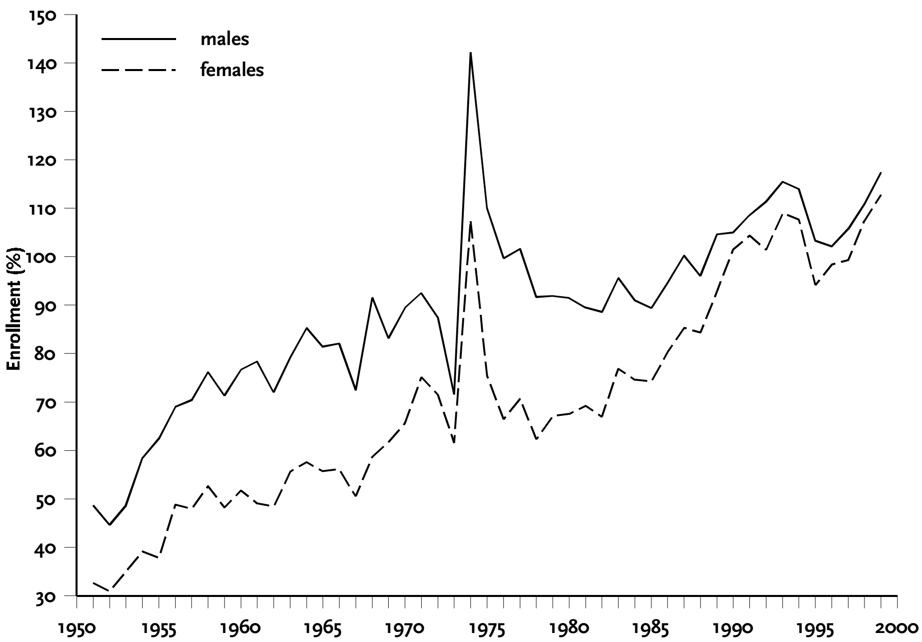
While the Dutch evidence typically suggests that students hardly respond to tuition fee changes, most US studies find larger effects. Leslie and Brinkman (1987) provide a meta-analysis on student price responses in American higher education, updated in Heller (1997). Their major conclusion is that students are responsive to prices and that – ceteris paribus – for every $ 100 increase in tuition price one would expect the participation rate to drop by about 0.7%-point. For an average weighted tuition fee of $ 3,420 and a national higher education participation rate of 0.33 in 1982/83 (*cf.* Leslie and Brinkman, 1987), this corresponds to a price elasticity of -0.73. Others (Manski and Wise, 1983; McPherson and Schapiro, 1991; Kane, 1995; Dynarski, 1999) add that particularly low-income students are more sensitive to tuition price levels than higher income students.

The plan of this paper is as follows. In Section 2 we describe the data. Section 3 outlines the econometric methodology and presents the results. Some concluding remarks are made in Section 4.

2 Demand for higher education in the Netherlands, 1950-'99

The variable of interest in this study is the demand for university training.[2] Figure 2.1 shows first-year university enrollment of male and female students as percentage of the number of qualified male and female secondary school graduates over the 1950-'99 period. It can thus be interpreted as the transition probability from upper secondary school to university training. However, the enrollment rate can exceed unity as there are other gateways to the university (for instance via post-compulsory upper vocational training programs). In addition, the inflow of students also includes foreign students and people that have postponed academic education. Male enrollment increased from 49% in 1950 to 117% in 1999, while female university enrollment increased from 33% in 1950 to 113% in 1999. So during the past half century male student inflow has expanded by about a factor 2.4 and female student inflow by a factor 3.4. It should be noted that the peak in university enrollment in 1975 is due to a system change (to be discussed in more detail below).

Figure 2.1 First-year university enrollment (as percentage of the number of qualified secondary school graduates)



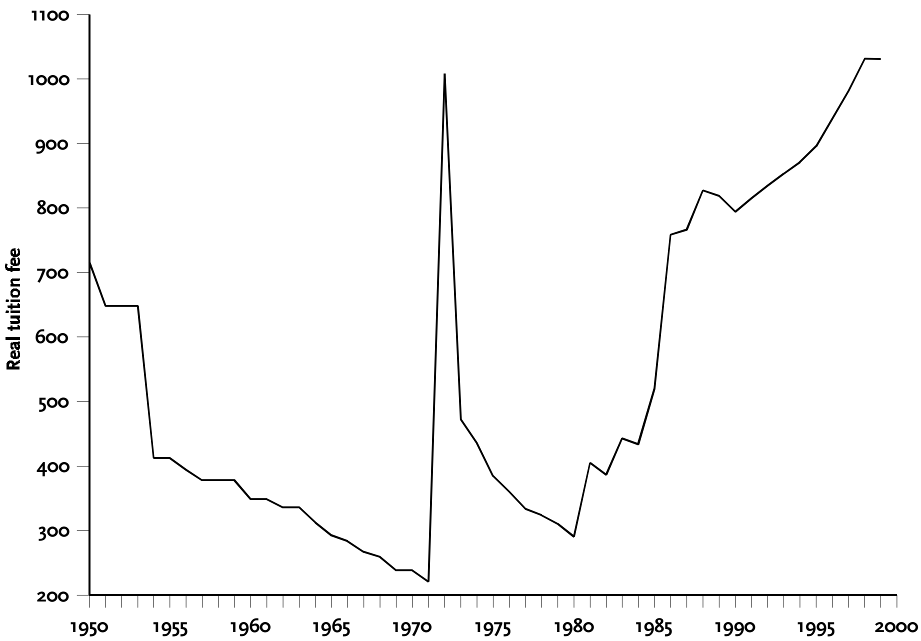
[2] Two types of training are distinguished in the Dutch higher education system, namely upper vocational training and academic training. This paper focuses on academic training, and if we talk about higher education we only refer to academic education. A natural suggestion for further research is therefore to carry out a similar analysis for upper vocational training.

Statistics Netherlands provides enrollment data for eight subject categories: Agriculture, Economics, Engineering, Health, Language & Culture, Law, Science, and Social Science. Separate series for male and female students are available. This allows us to construct a large data set. On the cross-section dimension we use 2×8=16 observations (male students in Agriculture, female students in Agriculture, male students in Economics, and so forth), and we have data for 50 years. In principle, this yields 800 observations for analysis. This is much more than in previous time-series studies on Dutch student enrollment.

To investigate the factors behind the enrollment decision, we recapitulate a number of potential explanations as suggested by the literature.

Tuition fees: Private contributions to the costs of a higher education program are expected to discourage university attainment, as tuition fees tend to lower the expected private return to schooling. Tuition fees for regular full-time students are centrally determined by the government in the Netherlands. Figure 2.2 illustrates the development of tuition fees (in 1990-prices) in the post-war period. After a gradual decrease (in real terms), tuition fees were drastically increased to a nominal amount of Dfl. 1000 (€ 454) in 1972. One year later the government reduced the tuition fee to a nominal amount of Dfl. 500 (€ 227). During the eighties and nineties, tuition fees more than tripled (in real terms).

Figure 2.2 Real tuition fees (€, in 1990-prices)

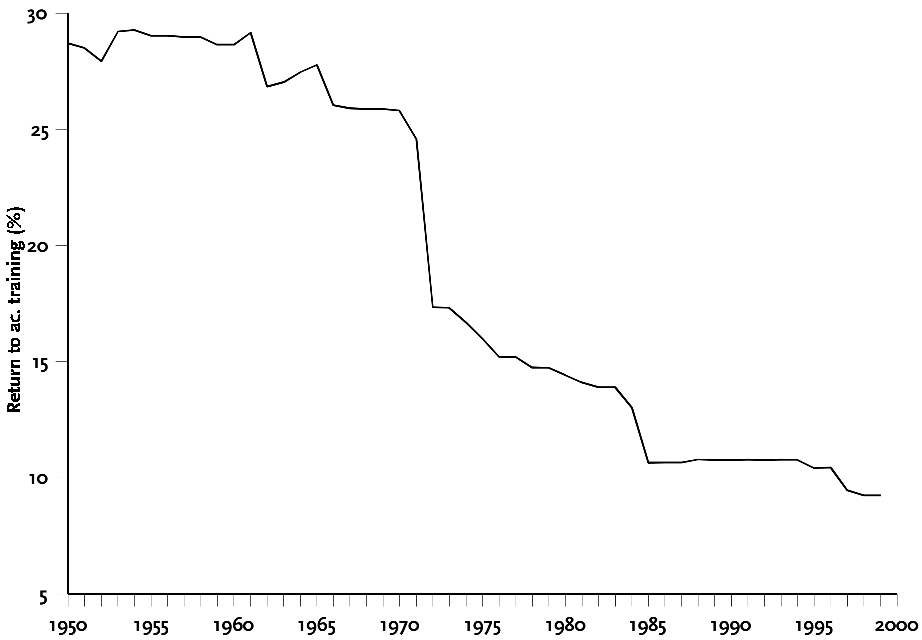


Student support: Financial aid to students by the government encourages university enrollment. We use a series on total public outlays on financial aid to university students and divide it by the total number of enrolled students to get a series for the average financial support per student.

Unfortunately, no distinction could be made between student grants and student loans. This is a shortcoming as grants and loans affect the enrollment decision in different ways. Grants lower the net private contribution to the educational investment, and thereby raise the expected private return to schooling. Loans, on the other hand, do not affect the expected return to schooling (when the loans carry the market interest rate), but may help to alleviate the credit market imperfection prevalent in the higher education market (*cf.* Chapter 2 in CPB and CHEPS, 2001).
Per capita income: The average income in society can influence the demand for higher education when education has a consumptive value. An alternative interpretation is that credit market problems are alleviated when the average income increases. Both stories predict a positive correlation between university enrollment and per capita income.
College premium: It has been argued that teenagers use information on the current wage gap between recent university and high school graduates to gauge the size of their own expected returns to schooling (*cf.* Freeman, 1976; Topel, 1997; Card and Lemieux, 2000). According to Webbink and Hartog (2000), Dutch students can predict their starting salary quite accurately. We therefore consider the actual return to higher education to be a proxy for the expected return. Secondary school graduates are encouraged to enroll in an academic training program when the expected returns to university training are high. Figure 2.3 shows the development of the annualized return to academic training for public sector employees (the "college premium") over the post-war period. There is a profound compression of the wage structure, especially during the seventies and early eighties.[3] It should be noted, however, that some studies point at an increase in the college premium in recent years (*cf.* Leuven and Oosterbeek, 2000), due to shortages on labor markets and skill-biased technological change. In the construction of the time-series we have assumed that the starting salary bracket does not change, whereas employers may offer more generous starting salaries in a tight labor market. But as the most profound changes in the college premium take place in the seventies, we are confident that the modest recent increase in the returns to schooling will not affect our main conclusions.

[3] Similar time patterns of the wage structure are observed in other countries. See, for instance, Goldin and Margo (1992) for the US experience.

Figure 2.3 College premium



Alternative wages: The wage that a person would have earned on the labor market after completing secondary school is a proxy for the opportunity costs associated with enrollment in a higher education program. It can be expected that an increase in this alternative wage discourages university attainment.

Unemployment: On the one hand one could argue that unemployment tends to discourage university attainment as it may lower the expected private return to schooling when graduates end up unemployed, but two counterarguments can be made. First, unemployment rates among skilled workers are typically lower than among the unskilled. Schooling can thereby help to escape (or at least reduce the possibility of) unemployment. Second, the opportunity cost hypothesis predicts that human capital accumulation tends to be counter-cyclical. Schooling becomes more attractive when times get bad, as the associated opportunity costs in terms of foregone labor market earnings are relatively low during recessions (*cf.* Bean, 1990; Saint-Paul, 1993).

The data appendix provides details on how the series are constructed and on data sources.

3 Econometric analysis

3.1 Model

There are a number of possibilities to specify the regression model. If we estimate the relationship between student enrollment and the explanatory variables in levels, we might face problems concerning stochastic trends present in for instance the time series for per capita income and financial support. In most cases such series are integrated of order 1, so that it is sufficient to take first differences to remove the stochastic trend. The disadvantage of doing so is that this is a rather restrictive method to remove the unit root, and that the regression analysis only picks up short-run responses to changes in the exogenous variables. A better alternative is found in an error correction specification, where both short-run and long-run effects are estimated and where account is taken of the stochastic trend problem. We therefore propose the following error correction model:

$$\Delta\ln(E_i^g(t)) = \phi_i^g[\ln(E_i^g(t-1)) - \ln(x'(t-1))\theta_i^g] + \Delta\ln(x'(t))\gamma_i^g + \varepsilon_i^g(t)$$
$$x' = [TF\ FS\ Y\ R\ W\ U] \tag{1}$$

where E_i^g is the inflow of students into university programs as a fraction of the number of qualified secondary school graduates for i=1, 2, ..., N cross-sectional units, and for g=male students, female students, all students; x_t is the vector of the independent variables. Each cross-section unit is observed for dated periods t=1, 2, ..., T. Changes in university enrollment are explained from changes in tuition fees (*TF*), financial support (*FS*), per capita income (*Y*), returns to schooling (*R*), opportunity costs (*W*), and the unemployment rate (*U*). An advantage of a log-linear model is that the vector of regression coefficients can directly be interpreted as elasticities.[4] γ is the vector of short-run elasticities; θ is the vector of long-run elasticities; $-1 \le \phi < 0$ is the convergence coefficient.

A closer inspection of the variables in the regression model reveals two extraordinary historical events, both in the early seventies. First, the nominal tuition fee was raised from Dfl. 200 (€ 91) in 1971 to Dfl. 1000 (€ 454) in 1972, *cf.* Figure 2.2. As a reaction, students refused to register at their institution. Because of this student boycott, enrollment data for 1972 and 1973 are missing. In 1973, the government reduced the nominal tuition fee to Dfl. 500 (€ 227). These shocks in tuition fees are large, especially when expressed in first differences. In addition, we constructed enrollment data for 1972 and 1973 through interpolation, but this method may be unreliable. We therefore introduce two year dummies for 1972 and 1973.

[4] We also ran the model with a logistic transformation as proposed in Pissarides (1982); results hardly changed.

Second, the secondary education system was drastically reformed in the late sixties (the so-called "Mammoetwet"). Because of structural changes, the reported number of school leavers qualified to enter university programs shows a strong downturn in 1973 and 1974, five and six years after implementation of the Mammoetwet in 1968. This is due to the separation of the former "hbs" ("hogereburgerschool") into "atheneum" (entitling people to enter university) and "havo" (which does not allow direct entrance to university programs). As this Mammoetwet has been introduced gradually (in 1968-'69 only the first year in secondary education; in 1969-'70 the first and second year and so forth), the full effect on the number of secondary school graduates eligible for university training materializes after five years. By including year dummies for 1973 and 1974, we try to take account of this structural reform of the education system. Also, to take account of the lag structure in our error correction model, we include a year dummy for 1975.

3.2 Aggregated data

As a first step in the econometric analysis, we aggregate the enrollment data across the eight cross-sections and run time-series regressions for male, female and total student inflow. Table 3.1 presents the results for male students. Equation [1] is the basic error correction model. In Equation [2] we added the lagged dependent variable as an additional explanatory variable. And in Equation [3] we shut down the adjustment mechanism and only include the short-run variables. The convergence coefficients in [1] and [2] are large, pointing at fast adjustment to the long-run equilibrium relationship in response to short-run deviations from that equilibrium. For instance, an adjustment coefficient of -0.679 means that half-way convergence to the equilibrium path only takes half a year. The long-run coefficients for male students suggest an important role for per capita income and the alternative wage rate in the university enrollment decision. For example, a 10% increase in per capita income leads to a 7-8% increase in male student inflow. The coefficients on tuition fees are negative, but insignificantly different from zero in both error correction specifications. Financial support only appears with a significantly positive coefficient in the first equation. A counterintuitive result is the negative long-run coefficient on the return to schooling, but this coefficient does not significantly differ from zero in Equations [1] and [2].The effect of unemployment on enrollment is weak and insignificant.

The short-run coefficients all have the "right" sign, but only financial support and the return to schooling appear with robustly significant coefficients. About 80% of the time-series variation of the dependent variable is explained by the models. Model [3], where the error correction mechanism is repressed, yields short-run responses more or less comparable to those in the first two equations.

Table 3.1	**University enrollment of male students, aggregated data (dependent variable: Δln(*E*))**					
	Equation [1] coefficient	standard error	Equation [2] coefficient	standard error	Equation [3] coefficient	standard error
Convergence coefficient						
ln(*E*(-1))	-0.679***	0.151	-0.573***	0.129		
Long-run coefficients						
ln(*TF*(-1))	-0.143	0.192	-0.377	0.292		
ln(*FS*(-1))	0.161**	0.067	0.084	0.092		
ln(*Y*(-1))	0.722***	0.189	0.836***	0.267		
ln(*R*(-1))	-0.457	0.412	-0.981	0.663		
ln(*W*(-1))	-0.804*	0.449	-1.398*	0.684		
ln(*U*(-1))	-0.038	0.041	-0.030	0.048		
Short-run coefficients						
Δln(*TF*)	-0.228*	0.131	-0.318**	0.154	-0.044	0.130
Δln(*FS*)	0.343**	0.126	0.304**	0.119	0.244**	0.097
Δln(*Y*)	0.212	0.558	0.558	0.463	-0.078	0.583
Δln(*R*)	0.438**	0.179	0.436**	0.164	0.464**	0.226
Δln(*W*)	-0.360	0.296	-0.590**	0.235	-0.125	0.274
Δln(*U*)	0.008	0.059	0.058	0.040	-0.037	0.044
Δln(*E*(-1))			-0.260**	0.125	-0.204	0.191
Diagnostic statistics						
N	48		47		47	
R^2_{adj}	0.83		0.86		0.74	
D.W.	2.33		2.49		2.13	

Note: Sample period is 1950-1999. The regression method is OLS, and standard errors are White heteroscedasticity-consistent. Time dummies for 1972, 1973, 1974, and (in [1] and [2]) 1975, and an intercept are included. *, **, *** means significant at the 10%, 5%, and 1% level, respectively. *E*, *TF*, *FS*, *Y*, *R*, *W*, and *U* stand for student enrollment, tuition fees, financial support, per capita income, return to university training, alternative wage, and unemployment, respectively.

The picture for female students is roughly similar, *cf.* Table 3.2. Financial support, per capita income and the alternative wage are the only variables appearing with significant long-run coefficients. But while the coefficients on financial support are comparable to those obtained in Table 3.1, per capita income and the alternative wage exert stronger effects on enrollment for female students. And also for female students we find that the short-run coefficients all have the "right" sign, except for the positive but insignificant coefficient on tuition fees in Equation [6]. The coefficients on financial support, the return to schooling, the alternative wage and the unemployment rate are significantly different from zero, and the short-run coefficient on the alternative wage is again much larger than we found for male students. So female students are more responsive than male students to changes in the alternative wage, both in the short-run and the long-run.

The results for all students, shown in Table 3.3, are logically in between the results for male and female students.

Table 3.2 University enrollment of female students, aggregated data (dependent variable: Δln(*E*))

	Equation [4] coefficient	standard error	Equation [5] coefficient	standard error	Equation [6] coefficient	standard error
Convergence coefficient						
ln(*E*(-1))	-0.779***	0.126	-0.774***	0.133		
Long-run coefficients						
ln(*TF*(-1))	-0.010	0.136	-0.076	0.159		
ln(*FS*(-1))	0.161**	0.064	0.121*	0.069		
ln(*Y*(-1))	1.312***	0.202	1.323***	0.210		
ln(*R*(-1))	-0.351	0.325	-0.507	0.411		
ln(*W*(-1))	-1.283***	0.322	-1.442***	0.365		
ln(*U*(-1))	-0.006	0.045	0.000	0.043		
Short-run coefficients						
Δln(*TF*)	-0.149	0.142	-0.185	0.143	0.026	0.121
Δln(*FS*)	0.356***	0.118	0.318**	0.130	0.238**	0.110
Δln(*Y*)	0.625	0.479	0.887*	0.497	0.144	0.481
Δln(*R*)	0.537***	0.194	0.512***	0.179	0.465**	0.202
Δln(*W*)	-1.006***	0.254	-1.112***	0.239	-0.558**	0.270
Δln(*U*)	0.093	0.055	0.135***	0.046	0.010	0.052
Δln(*E*(-1))			-0.110	0.084	-0.105	0.151
Diagnostic statistics						
N	48		47		47	
R^2_{adj}	0.80		0.82		0.64	
D.W.	2.03		2.07		2.02	

Note: Sample period is 1950-1999. The regression method is OLS, and standard errors are White heteroscedasticity-consistent. Time dummies for 1972, 1973, 1974, and (in [4] and [5]) 1975, and an intercept are included. *, **, *** means significant at the 10%, 5%, and 1% level, respectively. *E*, *TF*, *FS*, *Y*, *R*, *W*, and *U* stand for student enrollment, tuition fees, financial support, per capita income, return to university training, alternative wage, and unemployment, respectively.

Table 3.3 University enrollment of all students, aggregated data (dependent variable: Δln(*E*))

	Equation [7] coefficient	standard error	Equation [8] coefficient	standard error	Equation [9] coefficient	standard error
Convergence coefficient						
ln(*E*(-1))	-0.688***	0.135	-0.619***	0.127		
Long-run coefficients						
ln(*TF*(-1))	-0.102	0.167	-0.294	0.247		
ln(*FS*(-1))	0.166**	0.066	0.101	0.081		
ln(*Y*(-1))	0.957***	0.178	1.040***	0.234		
ln(*R*(-1))	-0.416	0.366	-0.854	0.567		
ln(*W*(-1))	-1.067***	0.387	-1.554**	0.577		
ln(*U*(-1))	-0.034	0.042	-0.026	0.045		
Short-run coefficients						
Δln(*TF*)	-0.201	0.130	-0.287*	0.147	-0.016	0.128
Δln(*FS*)	0.349***	0.122	0.312**	0.115	0.242**	0.092
Δln(*Y*)	0.375	0.486	0.705	0.424	-0.074	0.550
Δln(*R*)	0.512***	0.173	0.497***	0.156	0.489**	0.213
Δln(*W*)	-0.589**	0.262	-0.797***	0.215	-0.246	0.262
Δln(*U*)	0.030	0.052	0.079**	0.035	-0.030	0.044
Δln(*E*(-1))			-0.225**	0.107	-0.147	0.177
Diagnostic statistics						
N	48		47		47	
R^2_{adj}	0.83		0.86		0.72	
D.W.	2.27		2.43		2.11	

Note: Sample period is 1950-1999. The regression method is OLS, and standard errors are White heteroscedasticity-consistent. Time dummies for 1972, 1973, 1974, and (in [7] and [8]) 1975, and an intercept are included. *, **, *** means significant at the 10%, 5%, and 1% level, respectively. *E*, *TF*, *FS*, *Y*, *R*, *W*, and *U* stand for student enrollment, tuition fees, financial support, per capita income, return to university training, alternative wage, and unemployment, respectively.

3.3 Disaggregated data

The results reported so far were based on a model where the enrollment of students was aggregated over all studies. However, the pattern of adjustment to changes in tuition fees, financial support, and other exogenous variables may be different among studies. In a situation like that, the aggregation may not be valid and the coefficient of adjustment (ϕ) may be biased (see Pesaran and Smith, 1995). To check the robustness of our results, we estimate the error correction model (1) for each study separately. However, we impose that the long-run elasticities

(θ) are the same for each study. *I.e.*, we impose $\theta_i = \theta$, but leave ϕ_i and γ_i unrestricted.[5] This model is called the Pooled Mean-Group (PMG) estimator by Pesaran, Shin and Smith (1999). We estimate the resulting system of eight equations (one for each subject category, with the long-term elasticities restricted between equations) by a two-step Seemingly Unrelated Regressions (SUR) method.[6]

It should be noted that the returns to schooling, the unemployment rate, and perhaps also the alternative wage may differ between the different fields of study and between male and female students (*e.g.* Jacobs (2002) finds substantial differences in the private returns to schooling). We do not have the data to include such cross-sectional differences in the explanatory variables. Caution in the interpretation of the coefficients on the returns to schooling, the alternative wage, and the unemployment rate is therefore warranted.

Results for male students, female students and all students are shown in Table 3.4. The table reports median convergence coefficients and median short-run coefficients; standard errors are not reported.

[5] We also experimented with a dynamic fixed effects model, where we ran regressions on the pooled data set using an error correction structure and imposing homogeneity restrictions across the cross-sections on the convergence coefficient, the long-run coefficients as well as the short-run coefficients. Only the intercept term is allowed to vary across the eight cross-sections in this set-up. We found low convergence coefficients and implausibly large long-run effects. Results are therefore not reported.

[6] In the first step, we run an equally weighted, non-linear least squares optimization. From the residuals of this first step, we compute the variance-covariance matrix of the residuals. Its inverse is used in a SUR estimate, where we perform only one optimization step. This is asymptotically equivalent to iterating to convergence. All calculations have been performed in Eviews.

Table 3.4 **University enrollment, error correction model, pooled mean-group estimation (dependent variable: Δln(*E*))**

	Equation [10] Male students coefficient	standard error	Equation [11] Female students coefficient	standard error	Equation [12] All students coefficient	standard error
Convergence coefficient						
ln(*E*(-1))	-0.073		-0.112		-0.062	
Long-run coefficients						
ln(*TF*(-1))	-0.078	0.298	0.060	0.154	0.093	0.224
ln(*FS*(-1))	0.237**	0.112	0.270***	0.059	0.219***	0.083
ln(*Y*(-1))	-0.458	0.335	0.976***	0.179	-0.064	0.251
ln(*R*(-1))	-0.127	0.667	0.323	0.346	0.391	0.502
ln(*W*(-1))	0.926	0.662	-0.239	0.343	0.876*	0.492
ln(*U*(-1))	-0.027	0.070	0.108***	0.036	-0.008	0.052
Short-run coefficients						
Δln(*TF*)	-0.054		-0.050		0.021	
Δln(*FS*)	0.255		0.261		0.222	
Δln(*Y*)	-0.103		-0.311		-0.422	
Δln(*R*)	0.500		0.342		0.374	
Δln(*W*)	-0.090		-0.699		-0.194	
Δln(*U*)	-0.055		-0.021		-0.046	
N	384		384		384	

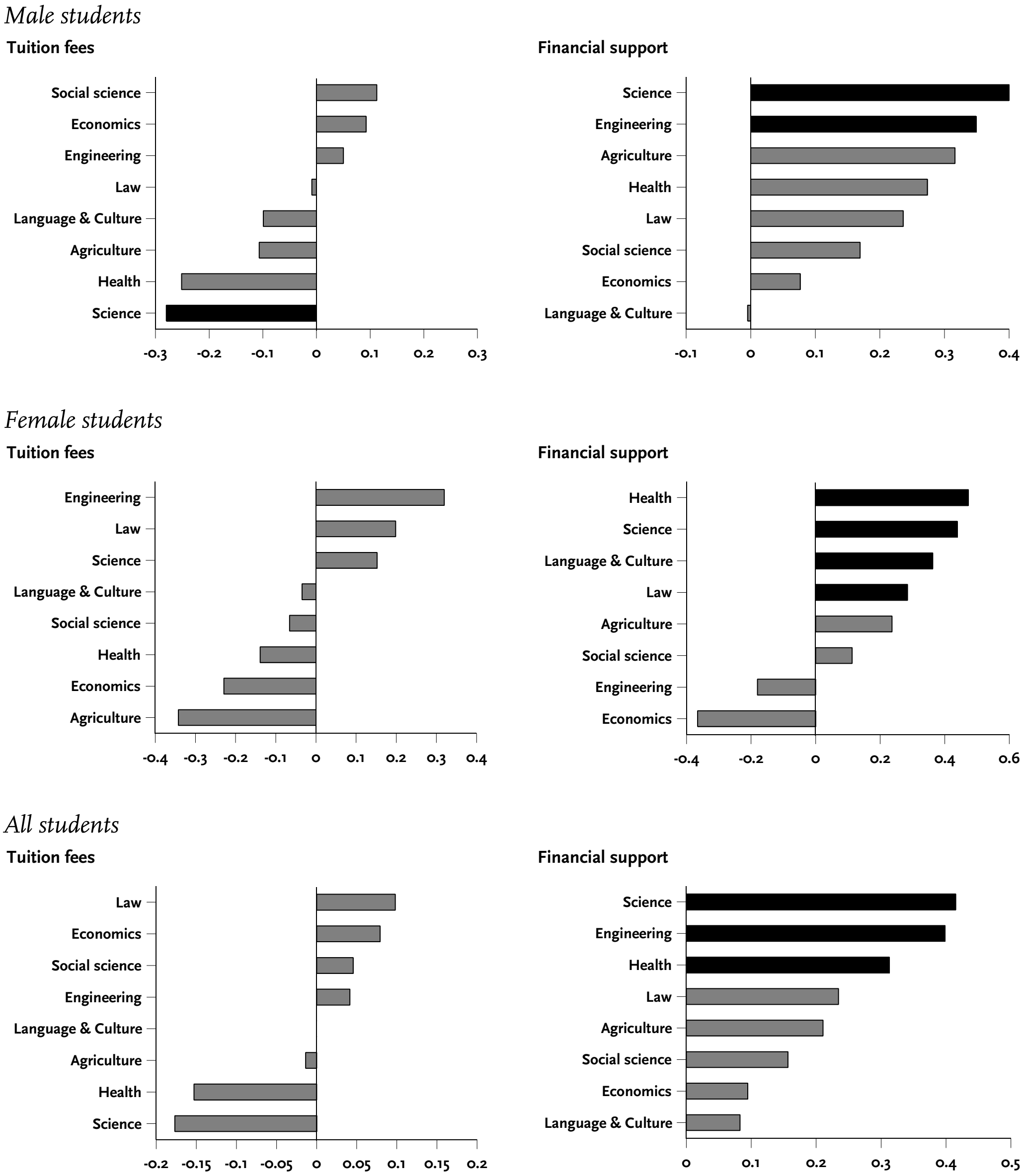
Note: Sample period is 1950-1999. The regression method is PMG combined with SUR. Time dummies for 1972, 1973, 1974, and 1975, and fixed effects are included. *, **, *** means significant at the 10%, 5%, and 1% level, respectively. *E*, *TF*, *FS*, *Y*, *R*, *W*, and *U* stand for student enrollment, tuition fees, financial support, per capita income, return to university training, alternative wage, and unemployment, respectively.

The results from PMG estimations differ to some extent from the results for the aggregated data reported in Tables 3.1-3.3. The convergence coefficient to the long-run equilibrium is much lower in the PMG estimations, and most long-run coefficients turn insignificant. The picture for the short-run elasticities is more in line with the earlier findings.

For male students, the estimates suggest a negative but insignificant long-run effect of tuition fees and unemployment on enrollment. The long-run impact of financial support is positive and significant at the 5% level. Per capita income, the return to schooling, and the alternative wage appear with the "wrong" sign, though the regression coefficients do not significantly differ from zero. With regard to the short-run coefficients, per capita income shows up with the "wrong" sign whereas the other explanatory variables appear with the expected sign. The results of the PMG estimations for female students are more in accordance with the earlier evidence (*cf.* Table 3.2). The long-term impact of tuition fees is again insignificant. Financial support and per capita income appear with positive and significant long-run coefficients. The long-run coefficients on the return to schooling and the alternative wage have the expected sign but are not significantly different from zero. A surprising finding is the statistically and economically significant long-run effect of unemployment on female student enrollment. The estimate suggests an enrollment elasticity with respect to unemployment of about 0.1. With respect to the short-run elasticities, again only per capita income shows up with an unexpected sign. Finally, a surprising result in Eq. [12] is the significant positive long-run effect of the alternative wage, though the coefficient is only significant at the 10% level.

The PMG procedure generates less precise estimations of the long-run effects, and only the financial support variable appears with robustly significant positive coefficients. The short-run effects are less sensitive to the estimation strategy, and are more or less in line with the econometric results presented in Section 3.2. In Figure 3.1 we present the estimated short-run elasticities with respect to tuition fees and financial support for each cross-section for male, female and all students, respectively. An eye-catching result is the responsiveness of male students in science to changes in tuition fees, and the responsiveness of male and female science students and male engineering students to changes in financial support. This brings us at the interesting relationship between education- and technology-policy. When the supply of scientists and engineers is inelastic, government support to R&D-activities basically translates into higher wages for research staff. Romer (2001) has therefore suggested to encourage student enrollment in science and engineering by abolishing tuition fees and providing generous grants to bright students as part of a policy to promote R&D. In light of our empirical results, this idea deserves to be taken seriously in the Dutch context.

Figure 3.1 Short-run elasticities with respect to tuition fees and financial support



Note: The black bars denote coefficients significantly different from zero at the 5% level; the grey bars denote coefficients insignificantly different from zero at the 5% level.

As a final check on the short-run effects, we shut down the error correction mechanism, and estimate the relationship between student enrollment and the economic variables in first-differences. Notice that homogeneity restrictions on the short-run coefficients are imposed, so it is assumed that students in different subjects react to changes in the economic environment in a similar way. Table 3.5 presents the results. The estimations of the short-run elasticities are roughly comparable to those obtained in the PMG estimations.

Table 3.5 University enrollment, first-differences model, pooled estimation (dependent variable: Δln(*E*))

	Equation [13] Male students coefficient	standard error	Equation [14] Female students coefficient	standard error	Equation [15] All students coefficient	standard error
Short-run coefficients						
Δln(*TF*)	-0.046	0.075	0.071	0.088	-0.012	0.076
Δln(*FS*)	0.231***	0.074	0.319***	0.087	0.262***	0.075
Δln(*Y*)	0.260	0.394	-0.106	0.464	-0.014	0.400
Δln(*R*)	0.589***	0.211	0.566**	0.248	0.522**	0.214
Δln(*W*)	-0.311	0.204	-0.627***	0.240	-0.430**	0.207
Δln(*U*)	-0.005	0.037	-0.013	0.044	-0.013	0.038
Δln(*E*(-1))	-0.063**	0.048	-0.175***	0.046	0.030	0.048
Diagnostic statistics						
N	376		376		376	
R^2_{adj}	0.54		0.31		0.54	
D.W.	2.06		2.25		2.09	

Note: Sample period is 1950-1999. The regression method is SUR. Time dummies for 1972, 1973, 1974, and 1975, and fixed effects are included. *, **, *** means significant at the 10%, 5%, and 1% level, respectively. *E*, *TF*, *FS*, *Y*, *R*, *W*, and *U* stand for student enrollment, tuition fees, financial support, per capita income, return to university training, alternative wage, and unemployment, respectively.

4 Conclusions

This paper presents new evidence on an old issue, the economic determinants of the demand for higher education. The few available studies for the Netherlands provide mixed evidence on the importance of economic factors in the spirit of the human capital approach. According to the results presented in this paper, the human capital approach should be taken seriously. Briefly, the main conclusions of this paper are the following:

- The econometric evidence on the aggregated data set may be unreliable when students' responses to changes in the economic environment differ across fields of study. Indeed, our PMG estimations yield less robust results for the long-run effects. The short-run effects show a more stable pattern, and are broadly consistent with the human capital hypothesis;
- The short-run responses of male and female students to changes in the economic variables are more or less comparable, though females seem to be more responsive to changes in the alternative wage. An interesting question for further research is whether behavioral differences can be explained from selection effects in the study choice. For instance, the human capital model may not tell the most plausible story behind the enrollment decision in the category Language & Culture. The estimated reaction to changes in the economic variables may differ when females are over-represented in this category;
- The enrollment elasticity with respect to tuition fees is typically weak and insignificant. This weak effect of tuition fees makes sense from the viewpoint of the human capital model, as this direct cost component is very small when considered against the gain in lifetime income associated with an academic degree;
- The positive contribution of financial support to student enrollment is non-trivial. The enrollment elasticity with respect to financial support is about 0.16-0.32, with robustly significant estimates. The short- and long-run effects are in the same order of magnitude. Unfortunately, the data do not allow an econometric appraisal of the separate contributions of student grants and loans. But the results may nonetheless be useful in the debate on reform of the student support system. Options for reform recently proposed in the Dutch context include the introduction of a student loan system with income-contingent repayment rates (such as the Australian Higher Education Contribution Scheme, *cf.* CPB and CHEPS (2001)), or graduate taxes (Jacobs, 2002);
- The evidence for a positive and robust long-run income elasticity obtained for the aggregated data vanishes when the disaggregated data is analyzed. The PMG estimations imply a positive and significant long-run enrollment elasticity with respect to per capita income only for female students. The short-run income elasticities are mostly insignificant. One should be careful to interpret a positive income elasticity as evidence for a consumption motive for higher education, as it could also reflect the role of credit market problems. As average incomes rise, capital market constraints will be alleviated so that more people can take up academic training

programs. However, when the government adequately responds to credit market imperfections by providing student loans, the estimated elasticity can be interpreted as a consumption effect;

- The short-run enrollment elasticity with respect to the college premium is somewhere between 0.4 and 0.6 (and significant in the investigated cases), and is comparable for male and female students. The long-run elasticities are insignificant;
- The short-run enrollment elasticity with respect to the alternative wage is negative and in most cases significantly different from zero. Most elasticities lie in the range of -0.3 to -0.6. Female students are more responsive to fluctuations in foregone labor market earnings than male students. The estimates of the long-run effects are imprecise;
- The typical impact of unemployment on student enrollment is weak and insignificant.

Winding up, our quantitative assessment of the various influences may be helpful in the design and evaluation of policy for the higher education sector. The empirical results in this paper suggest that economic factors are important in the university enrollment decision. The short-run impacts of the economic variables are broadly consistent with the view that higher education is an investment, though the long-run effects are less convincing.

Data appendix

Student enrollment: Statistics Netherlands provides enrollment data for eight subject categories: Agriculture, Economics, Engineering, Health, Language & Culture, Law, Science, and Social Science. Separate series for male and female students are available. Data over the 1950-'91 period are taken from Statistics Netherlands (1994), and updated until 1999 from CBS Statline. Because of a student boycott, enrollment data for 1972 and 1973 are missing. We construct enrollment data for these two years by linear interpolation. Enrollment rates are obtained by dividing the enrollment series by the number of high-school graduates eligible to register for university programs. The enrollment rate can thus be interpreted as the flow probability of moving from upper secondary school to university training. The number of high-school graduates eligible to register for university programs ("VWO") is taken from "historische reeksen mens en maatschappij" for the 1950-'95 period, and updated until 1999 from CBS Statline.

Tuition fees: Data on tuition fees are provided by CHEPS (Center for Higher Education Policy Studies, University of Twente) and Nota Klein (Klein, 1974).

Financial support: In "Statistiek van de Uitgaven der Overheid voor Onderwijs", Statistics Netherlands provides data on student support for the 1950-'91 period. Total public expenditures on student support include grants ("beurzen"), interest-free advances ("renteloze voorschotten") and two instruments targeted at specific groups ("gedemobiliseerden" and "studenten uit Indonesië en uit West-Indië"). Data for 1986 are missing, due to a radical reform of the student support system (implementation of "wet op de studiefinanciering"). As of 1991 students receive a public transport pass ("OV-studentenkaart"). The figures for 1991 are biased because of an erroneous estimate of the effects of the public transport pass by Statistics Netherlands. Therefore, the following adjustment is carried out. Total public expenditures on the public transport pass (€ 198 million) is multiplied by the ratio of university students eligible for the public transport pass and the total number of entitled students (about 0.25). This yields an amount of € 49 million, which is added to the total amount of grants and loans to university students, € 492 million. So we use a total amount of € 542 million, compared with an amount of € 562 million suggested by Statistics Netherlands. Data for the 1992-'99 period are available from OCenW (1998, 1999, 2000).

Per capita income: For the period 1950-'98 we use data on GDP (market prices) from Statistics Netherlands and RUG (2000), and the number for 1999 is taken from CPB (2001). Population size is available from Statistics Netherlands and RUG (2000) for the entire 1950-'99 period.

Unemployment: Unemployment data over the 1950-'91 period are available from CPB (1998) and the time-series is updated until 1999 from CPB (2001).

Consumer price index: The CPI over the 1950-'98 period is taken from Statistics Netherlands and RUG (2000) and updated for 1999 from CPB (2001).

College premium: As a proxy for the expected returns to higher education we use the percentage difference between the starting salary in the public sector of a university graduate and a high school graduate. We divide this percentage difference by five to obtain an estimate of the annual return to higher education for a five-year training period. High school graduates typically start in bracket 4/0 and university graduates generally start in 10/0 of the "BBRA 1984 schalen" ("Bezoldigingsbesluit Burgerlijke Rijksambtenaren"). BBRA 1984 applies from 1984 onwards. Before 1984 we use bracket 43 ("Adjunct commies / administratief ambtenaar B II") and 103 ("Hoofdcommies / administratief hoofdambtenaar") of BBRA 1948.
Alternative wages: The starting salary in the public sector of a high school graduate is taken as a proxy for a student's alternative wage.

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