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Changes in Educational Wage Premiums in Sweden: 1992-2001

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Department of Economics

Changes in Educational Wage Premiums in Sweden: 1992-2001

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Changes in Educational Wage Premiums in Sweden: 1992-2001[^]

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July 12, 2004

Abstract

This paper investigates changes in educational wage differentials in Sweden between 1992 and 2001 and places them in a longer-term perspective. The university wage premium has increased noticeably between 1992 and 2001 while the gymnasium wage premium has been constant. These results, together with previous research for earlier years, suggest that the university wage premium in Sweden has exhibited a clearly positive trend since the early 1980s while the gymnasium wage premium has been constant since the mid 1970s. Unlike the 1980s, relative supply changes do not appear to have been responsible for the rise in the university wage premium over the most recent decade.

JEL classification: J31

Keywords: Returns to skill, Wage structure, Inequality

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1. Introduction

This paper examines changes in educational wage differentials in Sweden between 1992 and 2001 and places them in a longer-term perspective. The wage differential changes that occurred in Sweden during the 1970s and 1980s are well documented (see Edin and Holmlund, 1995; Palme and Wright, 1998). Educational wage premiums fell rapidly during the 1970s; between 1968 and 1981, the average wage differential between individuals with a 3-year university education and a 3-year gymnasium education (i.e. high school) decreased from 46 percent to 14 percent, and decreases of a similar magnitude occurred between all education attainment levels (Edin and Holmlund, 1995).¹ By comparison, the 1980s did not exhibit the kind of across-the-board changes observed in the 1970s. The return to a university education increased, although the changes were sufficiently modest not to be able to offset the decline during the 1970s. The return to a gymnasium education remained low by comparison.

The international evolution of educational wage differentials during the 1970s and 1980s is also well documented (for a survey, see Katz and Autor, 1999). One clear theme in this literature is that the substantial fall in educational wage differentials during the 1970s was common to most industrialized countries. International patterns diverged, however, during the 1980s. Differentials widened sharply in the U.S. and the U.K., whereas most other industrialized countries experienced only small increases or unchanged differentials. For example, France and Germany's education premiums were fairly constant whereas Canada's rose only marginally.

Changes in educational wage differentials during the 1990s are less well documented internationally. One exception is the United States, where much research effort has been focused. According to Katz and Autor (1999) and Jaeger (2002), educational wage

¹ Calculated as the difference between 15 and 12 years of schooling based on Table 9.2, p. 319 in Edin and Holmlund (1995).

differentials in the U.S. have continued to increase but at a slower pace than during the 1980s.²

More knowledge of country-specific changes in educational wage premiums during the 1990s can shed light on such traditional themes as inequality and incentives for attending higher education but may also illuminate changes in the functioning of the labor market. It can for example be used to investigate the plausibility of the “skill biased technological change” hypothesis. This is done for the U.S. by Card and DiNardo (2002) who investigate how changes in the educational wage premium during the 1980s and 1990s correspond to changes in the spread of computers.

Changes in educational wage differentials in Sweden during the 1990s are also interesting because Sweden was one of the European countries that experienced the most dramatic change in its macroeconomic situation in the early 1990s; the unemployment rate rising from 1.6 percent in 1990 to 8.2 percent by 1993, and extended measures that capture hidden unemployment reaching double-digit levels (see Holmlund, 2003). Subsequently, a marked fall in unemployment occurred only during the last years of the 1990s, with the unemployment rate having returned to 4.0 percent by 2001. Previous Swedish studies have found that differences in wages, annual earnings, and disposable incomes systematically increased between low and high earnings groups during these dramatic changes (Gustafsson and Palmer, 2002; Gustavsson, 2004a).³ However, direct evidence on relative wage changes between education groups is meager.

One previous study has touched on changes in educational wage differentials in Sweden during the 1990s. Based on the Swedish Level of Living Surveys (LNU), Grand, Szulkin and Tåhlin (2001) compare the returns to schooling in 1991 and 2000 by estimating wage equations for these two years. Their estimates indicate that the return to one additional year of

² Another exception is Austria, where Fersterer and Winter-Ebner (2003) find falling returns to education during the 1990s.

³ See Gustafsson and Palmer (2002) for disposable incomes and Gustavsson (2004a) for earnings and wages.

schooling increased from 4.4 to 4.8 percent. However, they do not estimate changes between different education attainment levels, as they only allow for a linear relationship between years of schooling and log wages. They also rely upon a small sample and do not report standard errors, marking it difficult to judge the precision of their estimates.

This paper aims to improve on the research of Grand, Szulkin and Tåhlin (2001) by taking advantage of a substantially larger data set and considering variation across education attainment groups. The paper begins with a description of the data, with the estimation method described in Section 3. Section 4 presents results from Mincerian wage equations for the 1992-2001 period. The results, together with previous research for earlier years, indicate that the university wage premium has exhibited a clearly positive trend since the beginning of the 1980s, with particularly strong increases after the mid 1990s. The gymnasium wage premium, however, has been constant since the mid 1970s. Section 5 discusses how changes in relative supply correspond to the increase in the university wage premium. Unlike the 1970s and 1980s, relative supply changes do not appear to explain much of the movement in the university wage premium during the 1990s. The paper ends with concluding remarks.

2. Data

The data used in this analysis are from the longitudinal database LINDA (Longitudinal Individual Data for Sweden), constructed to be cross-sectionally representative of the Swedish population each year (see Edin and Fredriksson, 2000). The database is large; it contains 3.35 percent of the Swedish population each year.⁴ Information on individual background characteristics stems from various registers maintained by Statistics Sweden.

⁴ This corresponds to 300,000 individuals.

Information on individual wages is collected by Statistics Sweden through employers who are obligated by law to respond.⁵ Between 1992 and 1997, wages are available for individuals in LINDA who were also included in Statistic Sweden's annual study of the structure of wages in Sweden ("lönestrukturstatistiken"). This survey includes all individuals in Sweden employed in the public sector during a specific month, whereas around 50 percent of employees in the private sector employed in a specific month are sampled.⁶ Thus, wages are available for approximately half of the private sector employees in LINDA for the years 1992-1997 and potentially for all public sector employees.

From 1998 onwards, Statistics Sweden substantially improved its coverage of wages in LINDA. Individuals with missing values on the wage variable, either because they were not included in the study of wages in Sweden or classified as a non-response, have been included in an additional survey.⁷ In 1998, this additional survey was carried out only in the private sector but was extended to cover both the private and the public sectors between 1999 and 2001. Comparing the wage sample for 1998 to later samples, there are potentially too few observations from the public sector in 1998.

In Statistics Sweden's survey of the structure of wages in Sweden, the sampling frame for the private sector is made up of firms which are stratified according to industry and number of employees. Random draws are made within each stratum. Larger firms have a higher probability of being sampled. As a result, only 3 percent of firms with less than 10 employees are sampled. For LINDA, this means that wages for employees in small firms are mainly only available from 1998, as the wage information prior to this is only based on Statistics Sweden's survey of wages in Sweden.

⁵ The description of the collection of wages that follows builds on personal correspondence with Sara Andersson at Statistics Sweden.

⁶ The month of collection varies, but it has often been in April, May, September or October.

⁷ The additional surveys for 1998-2001 include individuals who were employed in November or December.

Table 1: Sample means

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
ln(wage), wage in real 2001 SEK	9.706 (.225)	9.684 (.234)	9.711 (.241)	9.707 (.237)	9.760 (.243)	9.800 (.247)	9.836 (.267)	9.867 (.280)	9.901 (.283)	9.920 (.288)
age	42.03 (11.08)	42.33 (10.88)	42.68 (10.68)	42.83 (10.76)	42.97 (10.79)	43.14 (10.77)	42.25 (11.18)	42.13 (11.31)	42.30 (11.37)	42.49 (11.44)
potential experience	23.23 (11.40)	23.52 (11.21)	23.81 (11.01)	23.92 (11.09)	24.04 (11.12)	24.17 (11.12)	23.44 (11.52)	23.29 (11.66)	23.40 (11.73)	23.55 (11.81)
primary education	.192	.190	.178	.172	.168	.163	.179	.174	.165	.159
2-year gymnasium	.367	.361	.357	.359	.355	.348	.345	.333	.329	.322
3-year gymnasium	.111	.115	.120	.121	.125	.132	.164	.178	.184	.193
some university	.151	.154	.158	.159	.161	.161	.145	.143	.143	.142
3-year university	.167	.168	.174	.175	.178	.182	.158	.162	.169	.175
doctorate	.012	.012	.013	.014	.014	.014	.010	.010	.001	.010
female	.593	.592	.580	.580	.569	.564	.487	.484	.493	.493
immigrant	.090	.087	.083	.084	.085	.086	.087	.094	.098	.102
public sector	.742	.686	.689	.683	.668	.643	.444	.416	.422	.413
observations	62,418	63,378	63,688	62,731	65,337	65,992	91,952	102,330	105,633	108,057

Note: Standard deviations are in parentheses.

LINDA contains information about individuals' highest completed education level.⁸ Six educational categories are available: (i) primary (education levels below gymnasium); (ii) 2-year gymnasium (generally vocational education); (iii) 3-year gymnasium (generally required for university entry); (iv) some university (one to two years, including college); (v) university (three years or more, including college); (vi) doctorate (research doctorates, such as Ph.D.).

The samples that I use in the estimation contain individuals aged 20-64 who are neither students nor self-employed.⁹ Table 1 contains descriptive statistics for the included individuals. The increased sample size for the private sector in 1998 is clearly seen as a jump in the values for most variables between 1997 and 1998.

⁸ From 2000, several new sources are used by Statistics Sweden to obtain individuals' highest education level; the educational information in 2000 and 2001 is therefore of better quality than that for the 1992-1999 period. Because LINDA is a panel from which individuals only drop out if they migrate or die, and because it is possible to observe in which year an individual's highest education was completed, the education variable for 2000 can be applied in most cases to individuals in the 1992-1999 samples; see Gustavsson (2004b) for details on this.

⁹ The wage variable contains some extreme values in each year; individuals with wages lower than the 0.5th percentile or above the 99.5th percentile in the wage distribution in a given year are therefore excluded in that year. These exclusions affect the estimated size of educational wage differentials but not their evolution over time.

The sample size for 1998 contains around 10,000 fewer individuals than 1999. This can be partially explained by better macroeconomic conditions in 1999 which drew more individuals into employment. The additional survey of wages for employees in the public sector carried out in 1999 but not in 1998 is also partly responsible (see the discussion above).¹⁰

Wages are defined as monthly wages in full time equivalents and are measured during a specific month each year.¹¹ To get a clearer picture of the evolution of real wages over time, wages in Table 1 are deflated by the consumer price index (“KPI”) and expressed as real 2001 SEK. However, deflated variables are not used in the estimated wage equations. Consistent with the results in SCB (2000), there is a noticeable increase in real wages over time. The standard deviations show that wage dispersion has grown as well.

The average age in the samples generally increases over time which presumably reflects the disproportionate increase in non-employment among youths during the 1990s (see Holmlund, 2003).¹² However, because average age is lower in the private sector and the sample size for the private sector increases in 1998, the average age exhibits a one-off decrease between 1997 and 1998.

Next in Table 1 is potential labor market experience. This variable is calculated as age minus years of schooling minus seven. Years of schooling are imputed from the six education categories in Table 1.¹³ As for average age in the samples, average labor market experience generally increases over time.

The share of individuals in the samples with only primary education decreases over time. This is because most individuals in younger cohorts have a gymnasium education (see

¹⁰ There are also reasons to suspect that there was a larger non-response in the additional survey in 1998. Unfortunately, Statistics Sweden has not documented the non-response rates for 1998 and 1999.

¹¹ The months differ between years; see footnote 6.

¹² The increase in the samples occurs in both the private and the public sector, but it is larger in the public sector.

¹³ Primary education corresponds to 9 years, 2-year gymnasium to 11 years, 3-year gymnasium to 12 years, some university to 13 years, university to 15 years, and a doctorate to 19 years.

Björklund *et al*, 2004). From 1991, virtually all gymnasium educations in Sweden have been 3 years long. Thus, the share of individuals in the sample with only a 2-year gymnasium education decreases over time, while the share with a 3-year gymnasium education increases. University education has also expanded over time, and its share in the samples increases accordingly.

Because of the over-representation of females in the public sector, the share of women in the 1992-1997 samples is well-above 50 percent but decreases markedly in 1998 when more individuals from the private sector are included. Table 1 also shows that the share of individuals employed in the public sector decreases until the mid 1990s. This reflects the marked reduction of public sector employment that occurred during this period (see Holmlund, 2003).

3. Estimation method

To investigate changes in educational wage differentials, Mincerian wage equations are estimated each year between 1992 and 2001. The log of wages is the dependent variable and the explanatory variables are the educational variables in Table 1 (with primary education as the omitted category), potential experience and potential experience squared, and dummy variables for females and immigrants (individuals born outside of Sweden).

It is important to recognize that an assumption of the Mincerian wage equations is separability of wages in education and experience. That is, that the experience profiles are parallel across education levels. Heckman *et al* (2003) find that this assumption is not fulfilled for annual earnings in the United States and, because of this, Mincer equations yield biased results for the rate of return to education in the United States. Appendix B investigates how this assumption affects my estimates of changes in educational wage differentials. Overall, the Mincerian wage equations are found to correctly capture the main patterns in the data at hand.

Least squares with sample weights is used to estimate the wage equations, where the weights recognize that there are too few individuals from the private sector in the 1992-1997 samples. That is, an observation from the private sector is simply weighted with the probability that an observation from the private sector is included in the sample. The corresponding is true for an observation from the public sector. To be consistent, weighted least squares is also applied to the 1998-2001 samples.¹⁴

It would be preferable to use weights that take account of the firm size of a worker's employer, as workers from small firms are under-represented in the 1992-1997 samples. Unfortunately, this is not possible because there is no information in the data on the size of the employers. This means that if the conditional expectation of wages in small firms is systematically different from that in larger firms, the sample composition will cause the estimated wage differentials to be biased. However, the main focus in this paper is on *changes* in educational wage differentials, rather than their absolute values, and there is no clear a priori reason to expect that changes in educational wage premiums should be markedly different in small firms. In Appendix A, I also check how the under-representation of workers from small firms influences estimates based on annual earnings, which are available from tax reports for all individuals in LINDA each year; the obtained results are discussed in the next section.

¹⁴ To obtain the sample weights for the analysis, I use the fact that the LINDA samples are representative of the Swedish population and that information on annual earnings are available for all individuals through tax reports. For a given year, I exclude individuals from the LINDA sample who are students, self-employed, or have zero annual earnings. The remaining individuals 20-64 years old are considered to be a representative sample of the population of interest and are used to obtain the weights

4. Results

Table 2 presents the estimated wage equations for the 1992-2001 samples.¹⁵ All coefficients are significantly different from zero and precisely estimated. The estimate for *2-year gymnasium* is slightly lower for 2001 than for 1992 while the estimate for *3-year gymnasium* can be regarded to be the same for 2001 as for 1992. The estimates for *some university*, *university*, and *doctorate* are noticeable higher in 2001 than in 1992, the increase occurring only after 1997.

An interesting wage differential is that between *university* and *3-year gymnasium* (also referred to as “the university wage premium”), as a 3-year gymnasium is general requirement for entry to university. Table 2 indicates that the log wage differential between these categories increased from 0.182 (20.0 percent) in 1992 to 0.251 (28.5 percent) in 2001.

Based on annual earnings, Appendix A investigates how the under-representation of workers from small firms in the 1992-1997 samples affects the estimates in Table 2. The analysis indicates that the estimated increases between 1997 and 1998 are too large but that the changes between other years are correct. The magnitude of educational wage differentials *per se*, do, however, appear to be affected by the sample composition during 1992-1997, but the bias appears to be small.

As I do not have access to the ideal weights in the estimations, I have checked the robustness of the results in Table 2 with respect to the used weights. None of the main results concerning changes in educational wage differentials (and changes in the other explanatory variables) are altered when the estimations are performed without weights. The same applies when only a dummy variable is included for employees in the public sector; see Table C1 in Appendix C.

¹⁵ I have also performed these estimations with the age interval set to 30-59 each year. The obtained results are very similar to those in Table 2.

Table 2: Estimated wage equations 1992-2001

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.048 (.002)	.052 (.002)	.051 (.003)	.044 (.003)	.037 (.003)	.034 (.003)	.040 (.002)	.044 (.002)	.041 (.002)	.037 (.002)
3-year gymn.	.141 (.004)	.150 (.004)	.150 (.004)	.139 (.004)	.129 (.004)	.120 (.004)	.131 (.003)	.139 (.003)	.133 (.003)	.132 (.003)
some univ.	.179 (.003)	.191 (.003)	.186 (.003)	.178 (.003)	.169 (.003)	.174 (.003)	.202 (.003)	.225 (.003)	.228 (.003)	.236 (.003)
univ.	.323 (.003)	.344 (.003)	.337 (.003)	.326 (.004)	.321 (.004)	.324 (.003)	.345 (.003)	.370 (.003)	.371 (.003)	.383 (.003)
doctorate	.490 (.011)	.531 (.010)	.509 (.010)	.498 (.009)	.499 (.010)	.508 (.010)	.520 (.009)	.583 (.009)	.587 (.009)	.596 (.009)
exp	.016 (.000)	.017 (.000)	.018 (.000)	.018 (.000)	.018 (.000)	.018 (.000)	.019 (.000)	.020 (.000)	.019 (.000)	.020 (.000)
exp ² /100	-.025 (.001)	-.026 (.001)	-.028 (.001)	-.028 (.001)	-.027 (.001)	-.028 (.001)	-.030 (.001)	-.031 (.001)	-.031 (.001)	-.032 (.001)
female	-.164 (.002)	-.166 (.002)	-.170 (.002)	-.163 (.002)	-.177 (.002)	-.180 (.002)	-.172 (.001)	-.163 (.001)	-.174 (.001)	-.173 (.001)
immigrant	-.044 (.003)	-.049 (.003)	-.050 (.003)	-.047 (.003)	-.045 (.003)	-.048 (.003)	-.061 (.003)	-.077 (.003)	-.079 (.002)	-.081 (.002)
constant	9.358 (.004)	9.354 (.004)	9.385 (.004)	9.417 (.004)	9.495 (.004)	9.532 (.004)	9.521 (.003)	9.531 (.003)	9.597 (.003)	9.634 (.003)
obs.	62,418	63,378	63,688	62,731	65,337	65,992	91,952	102,330	105,633	108,057
R ²	.394	.401	.381	.365	.358	.359	.336	.337	.332	.336
σ_e	.183	.197	.198	.198	.202	.204	.218	.228	.232	.236

Note: Weighted least squares is used; see the text for details. White's (1980) heteroskedasticity robust standard errors are in parentheses.

How do the changes in educational wage premiums in Sweden between 1992 and 2001 compare to those during the 1970s and 1980s? Figure 1 contains the evolution between 1968 and 2001 of the log wage differential between *university* and *3-year gymnasium* and between *3-year gymnasium* and *primary*. The estimates for 1968-91 are from Edin and Holmlund (1995). The data that they use are collected through surveys where the individual is the respondent.¹⁶ In LINDA, data on individual background characteristics are based on registers, while wages are collected through surveys with the employer as the respondent. These differences in data collection may affect the magnitude of the estimates. To aid visual

¹⁶ The estimates in Edin and Holmlund (1995) are based on data from the Swedish Market and Nonmarket Activities survey (HUS) and the Swedish Level-of-Living survey (LNU).

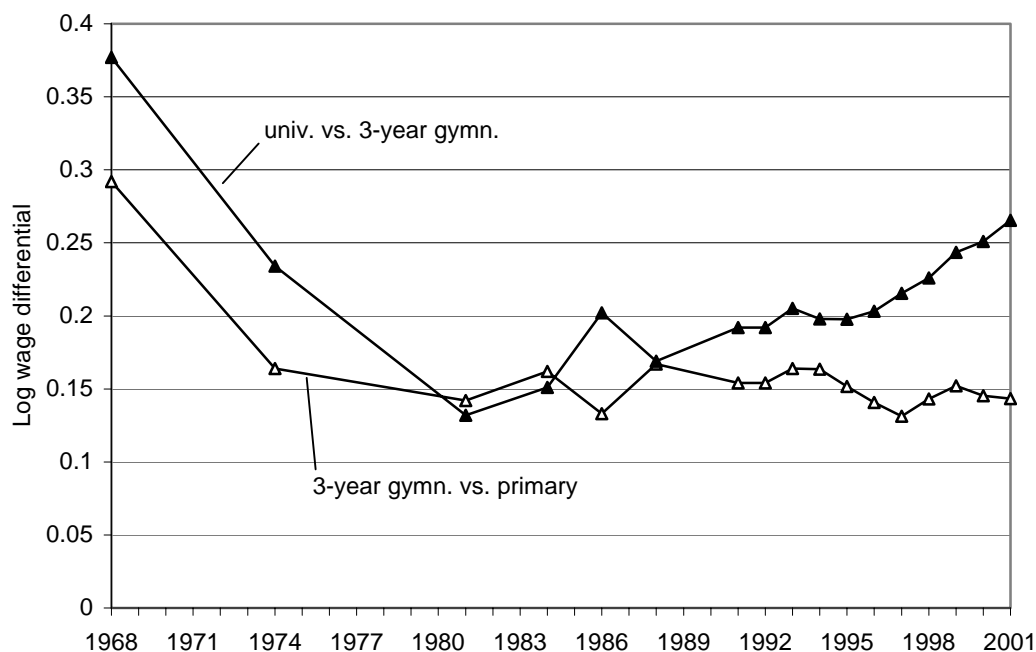


Figure 1: Log wage differentials 1968-2001; *university* versus *3-year gymnasium* and *3-year gymnasium* versus *primary*.

Source: The numbers for 1968-1991 are based on Edin and Holmlund (1995), Table 9.2 (12 and 15 years of education), p. 318. The numbers for 1992-2001 are based on Table 2 and have been re-scaled; see the text for details.

interpretation, my estimates for the 1992-2001 samples in Figure 1 are re-scaled to match as closely as possible the magnitude of their estimates. This is done by assuming that the educational wage differentials in question were constant between 1991 and 1992.

Figure 1 shows that the wage differential between *university* and *3-year gymnasium* fell sharply between 1968 and 1981. It also shows that the increase during the 1980s continued for another decade beyond the end of Edin and Holmlund's sample. Even though the sizes of the year-to-year changes during the 1980s and 1990s are not of the same magnitude as between 1968 and 1981, the steady increase in the wage premium in later decades has offset a significant portion of the 1968-1981 decrease. In fact, the university wage premium is higher in 2001 than in 1974, albeit with reservations about the comparability of the estimates.

Like the university-gymnasium differential, the log wage differential between *3-year gymnasium* and *primary* fell sharply between 1968 and 1974. Unlike the university-gymnasium differential, however, the gymnasium-primary differential exhibits no long-term

trend after the mid 1970s; there are short-term fluctuations but the magnitude of the differential is similar in 2001 to that in the mid 1970s.

The estimated wage equations in Table 2 contain some interesting results concerning changes in other dimensions of the Swedish wage structure between 1992 and 2001. The estimates for the experience variables indicate a markedly steeper experience-wage profile in 2001 than in 1992. There is also a steeper age-wage profile in 2001 than in 1992; see Table C2 in Appendix C for estimations where age is used instead of potential experience in the wage equations. These results point to a clear break to the flattening of the experience-wage and age-wage profiles that occurred during the 1970s and 1980s (see Edin and Holmlund, 1995).

The estimates for the female dummy variable in Table 2 indicate that the male-female wage differential has actually slightly increased between 1992 and 2001. However, the analysis of sample composition bias in Appendix A, based on annual earnings, indicates that the sample composition might affect the estimated changes in Table 2. At the very least, the male-female wage differential has not decreased between 1992 and 2001. Together with the results in Edin and Holmlund (1995), this indicates that the male-female wage differential in Sweden, as captured by a simple dummy variable, has been roughly constant, or perhaps increased slightly, since the mid 1980s.

Table 2 indicates a near doubling of the wage differential between individuals born in Sweden and immigrants; the wage differential increased from 4.5 percent in 1992 to 8.4 percent in 2001. The analysis in Appendix A suggests that sample composition bias partly drive this result, but it cannot explain the whole increase. Of course, grouping all individuals born outside of Sweden into a single variable hides a lot of information and the growing wage

differential may be due to compositional changes among immigrants.¹⁷ To further investigate the cause of these estimates is beyond the scope of this paper.

As captured by the values of the R-square in Table 2, the variables included in the wage equations have become less important over time for explaining variability in wages. There is also a systematic increase in the standard errors of the empirical residuals.¹⁸ These two results could be due to increased within-group inequality, but may also be driven by increased transitory wage variability. The results in Gustavsson (2004a) are more supportive of the former explanation. It is hence possible that increased within-group inequality accounts for a majority of the overall increase in Swedish wage inequality during the 1990s.

Having documented the overall changes in Sweden between 1992 and 2001, I now turn to a more detailed description of changes in educational wage differentials. One interesting question is whether the changes in wage differentials differ across gender. Table 3 contains weighted least squares estimates for 1992 and 2001 for females and males, respectively; results for the full 1992-2001 period are presented in Table C3 and C4 in Appendix C.¹⁹ For both sexes, the estimate for *2-year gymnasium* decreases between 1992 and 2001 while there are minor changes in the estimate for *3-year gymnasium*. The estimates for *some university* and *university* increase for both sexes, but the increases are about twice as large for males. There are, however, no differences in the evolution of the estimate for *doctorate*, as both sexes have experienced a substantial increase between 1992 and 2001. Table 3 also shows

¹⁷ Because of this, I have also estimated wage equations based on samples where all immigrants are excluded. The obtained point estimates are very similar to those in Table 2, and all the conclusions regarding changes over time are unaffected.

¹⁸ Technically, there is no single standard error of the residual. Because I use weights, there is heteroskedasticity in the residuals, and hence, the standard error of the residual differs between individuals. Also, the fall in the estimate between 1997 and 1998 may be due to the inclusion of more individuals from the private sector.

¹⁹ The weights differ to those used in Table 2, as the weights now are obtained separately for females and males. That is, a female from the private sector is weighted with the probability that a female from the private sector is included in the sample, and so forth. However, the results are not sensitive to the choice of weights.

Table 3: Estimated wage equations 1992 and 2001 for males and females

	Females		Males	
	1992	2001	1992	2001
2-year gym.	.037 (.003)	.027 (.002)	.061 (.004)	.044 (.003)
3-year gym.	.093 (.005)	.106 (.003)	.166 (.005)	.150 (.004)
some univ.	.161 (.003)	.198 (.003)	.196 (.005)	.269 (.004)
univ.	.294 (.004)	.336 (.003)	.348 (.006)	.427 (.005)
doctorate	.520 (.021)	.618 (.017)	.489 (.012)	.596 (.011)
exp	.011 (.000)	.016 (.000)	.021 (.001)	.023 (.000)
exp ² /100	-.017 (.001)	-.026 (.001)	-.032 (.001)	-.038 (.001)
immigrant	-.022 (.004)	-.055 (.003)	-.065 (.005)	-.110 (.004)
constant	9.277 (.004)	9.532 (.004)	9.285 (.006)	9.577 (.005)
obs.	36,992	53,239	25,426	54,818
R ²	.349	.300	.305	.288
σ_e	.146	.202	.213	.262

Note: Weighted least squares is used; see the text for details. White's (1980) heteroskedasticity robust standard errors are in parentheses.

that females have had larger absolute increases in the estimates for the experience variables.²⁰

The estimates for *immigrant* are, however, similar for both females and males.

Based on the estimates in Table C3 and C4, Figure 2 graphs the evolution of the university wage premium (i.e. *university* versus *3-year gymnasium*) for females and males, respectively. There are systematic differences; the wage premium for males displays a steady increase while it displays a U-shaped pattern for females.

A large fraction of employment in Sweden takes place in the public sector and a majority of those employed are women.²¹ The most important industries within the public

²⁰ Because female labor force participation in Sweden increased strongly during the 1970s, potential labor market experience is a better approximation for actual experience among younger women. As such, potential experience is a better measure in 2001 than in 1992, and this could cause the larger increase.

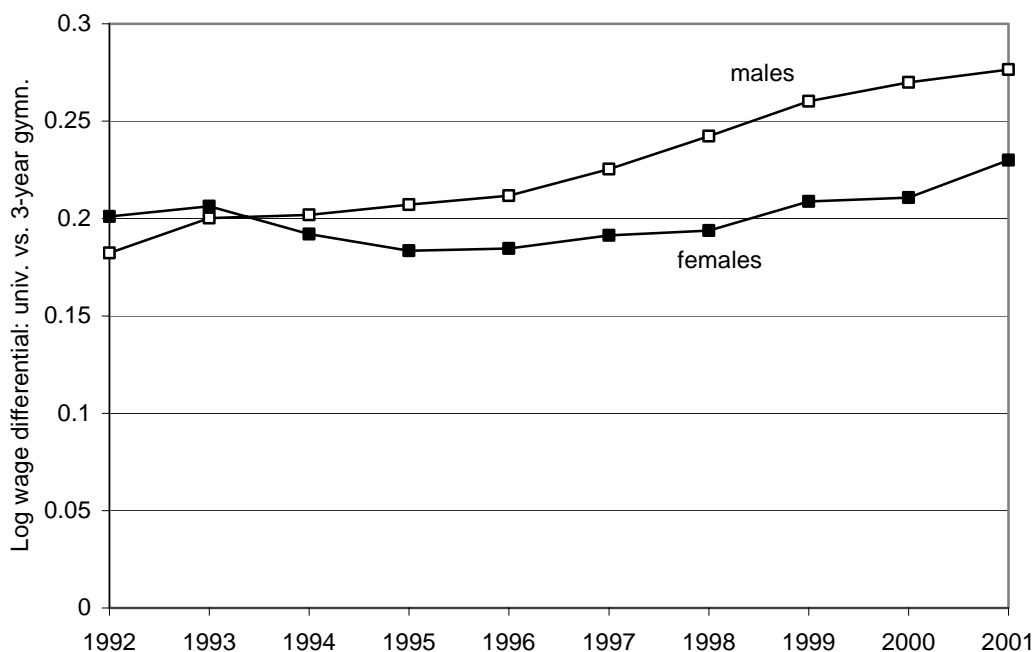


Figure 2: University wage premiums 1992-2001 among females and males.

sector are also generally specific to this sector, with health and social work, education, and public administration being the most important activities. Table C1 in Appendix C indicates that the wage differential between the private and the public sector, as captured by a simple dummy variable, has increased from 9.1 percent in 1992 to 13.4 percent in 2001. It is thus reasonable to ask whether changes in educational wage premiums have been uniform across the public and the private sectors.

Table 4 contains results for 1992 and 2001 for the public and private sectors respectively; results for the full 1992-2001 period are presented in Table C5 and C6 in Appendix C. For both sectors, the estimate for *2-year gymnasium* is lower in 2001 than in 1992, while there are only small changes in the estimate for *3-year gymnasium*. The estimates for *some university* and *university* increase for both sectors but by twice as much in the private sector. There is a marked increase in the estimate for *doctorate* in both sectors, the largest increase occurring in the public sector. For the experience variables and the female

²¹ Table 1 shows that 41 percent of the individuals in the 2001 sample are employed in the public sector. Of these, 69 percent are women.

Table 4: Estimated wage equations 1992 and 2001 for the private and the public sector.

	Private sector		Public sector	
	1992	2001	1992	2001
2-year gym.	.061 (.004)	.048 (.003)	.061 (.002)	.052 (.003)
3-year gym.	.169 (.005)	.153 (.003)	.113 (.003)	.127 (.004)
some univ.	.208 (.006)	.286 (.004)	.196 (.002)	.236 (.003)
univ.	.407 (.008)	.478 (.004)	.331 (.003)	.374 (.003)
doctorate	.563 (.033)	.657 (.019)	.532 (.010)	.652 (.010)
exp	.020 (.001)	.023 (.000)	.014 (.000)	.017 (.000)
exp ² /100	-.031 (.001)	-.038 (.001)	-.021 (.001)	-.026 (.001)
female	-.159 (.003)	-.142 (.002)	-.113 (.002)	-.128 (.002)
immigrant	-.053 (.005)	-.102 (.003)	-.047 (.003)	-.062 (.003)
constant	9.330 (.006)	9.595 (.004)	9.298 (.003)	9.541 (.005)
obs.	16,117	63,380	46,301	44,677
R ²	.375	.332	.483	.443
σ_e	.198	.248	.153	.193

Note: Weighted least squares is used; see the text for details. White's (1980) heteroskedasticity robust standard errors are in parentheses.

dummy variable, there are no striking differences between the two sectors. There is, however, a substantially larger increase in the immigrant-native wage differential in the private sector.

Based on the estimates in Table C5 and C6, Figure 3 graphs the evolution of the university wage premium within the private and the public sectors, respectively. There are striking differences; the private sector exhibits an almost linear increase between 1992 and 2001 compared to relative stability of the premium in the public sector.

The university wage premium has, to a large extent, only increased in the private sector. At the same time, the increase is larger among males than females. To further investigate these differences, I have estimated separate wage equations for females and males within the

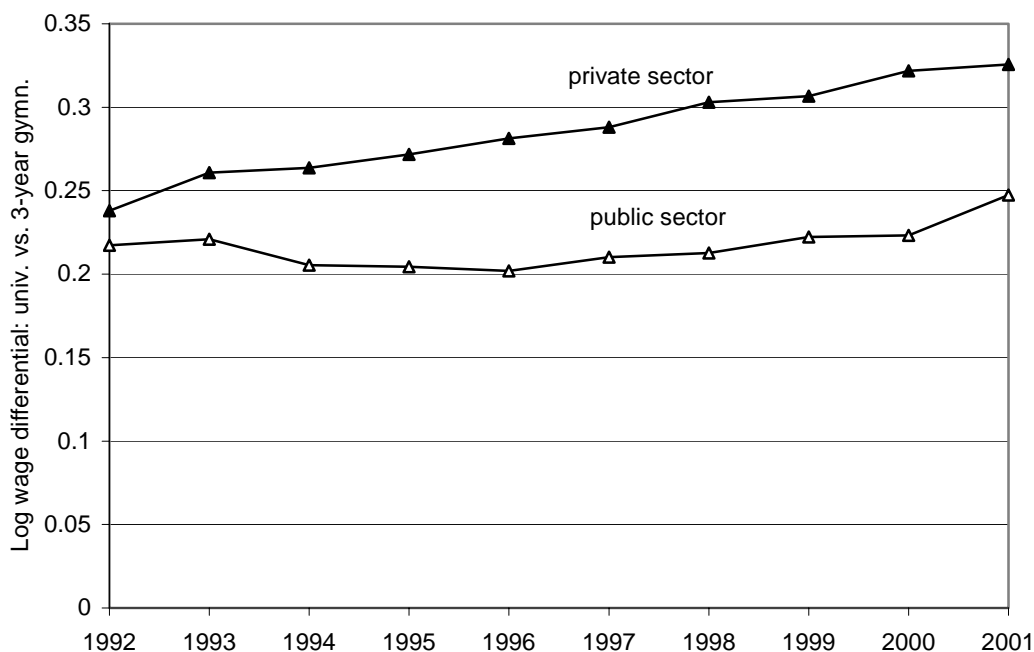


Figure 3: University wage premiums 1992-2001 within the private and public sectors.

public and the private sector; the full sets of results are presented in Table C7-C10 Appendix C. Figure 4 graphs the estimated university wage premiums.

From Figure 4, one can see that the university premium in the public sector has increased among males but displays a U-shaped pattern among females. It is striking that the wage differential among females for the 1992-1995 period converges toward that for males, after which they remain practically identical. The corresponding is not true for differentials between the gymnasium categories and primary education; see Table C7 and C8. In the private sector, the university wage premium has increased for both sexes, with the increase during the second half of the 1990s larger for males.

Overall, the increase in the university wage premium during the second half of the 1990s has been larger among males in both the private and the public sector. This increase potentially drives a majority of the overall increase in the university wage premium observed in the full sample.

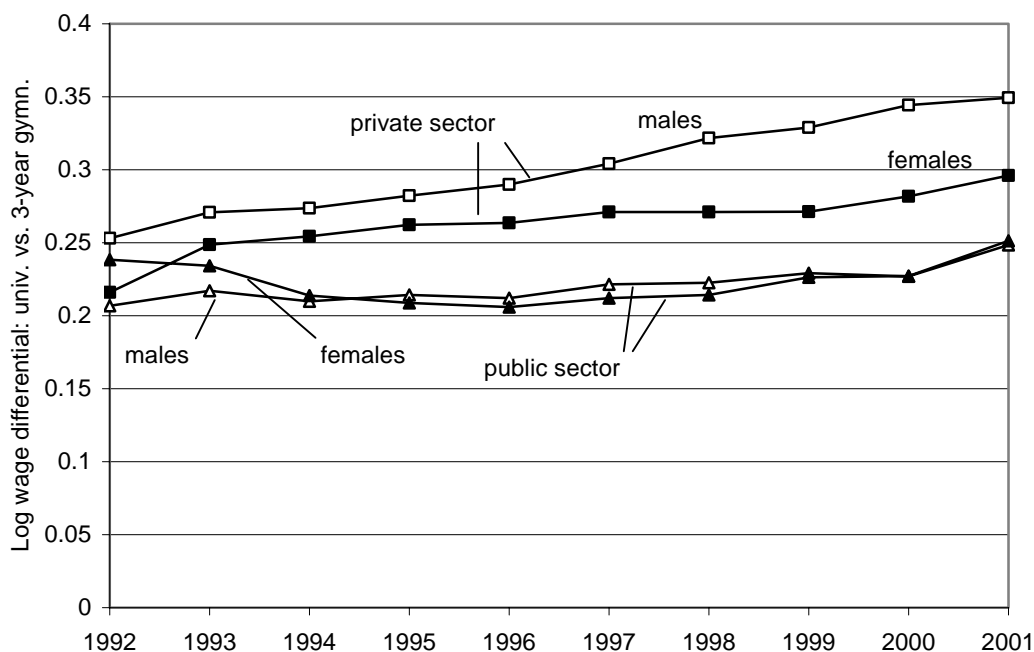


Figure 4: University wage premiums 1992-2001 among females and males within the private and public sector.

5. The university wage premium and changes in supply and demand

The previous section made it clear that there have been substantial movements in the Swedish university wage premium between 1968 and 2001. What is driving these large changes? For the 1970s it may be close at hand to attribute the decrease to the heyday of the wage compressing “solidarity wage policy” pursued by the major blue-collar union, Landsorganisationen (LO). However, Edin and Holmlund (1995) argue that even though this policy may explain part of the pay compression among individuals with lower levels of education, its explanatory power for movements in the university wage premium is seriously incomplete. They instead argue that fluctuations in the relative supply of individuals with a university education better explains changes in the university premium during the 1970s and the 1980s; that is, both before and after the breakdown of the solidarity wage policy in 1983.

How does the increase in the university wage premium during the 1990s correspond to changes in relative supply? To investigate this I follow the approach of Edin and

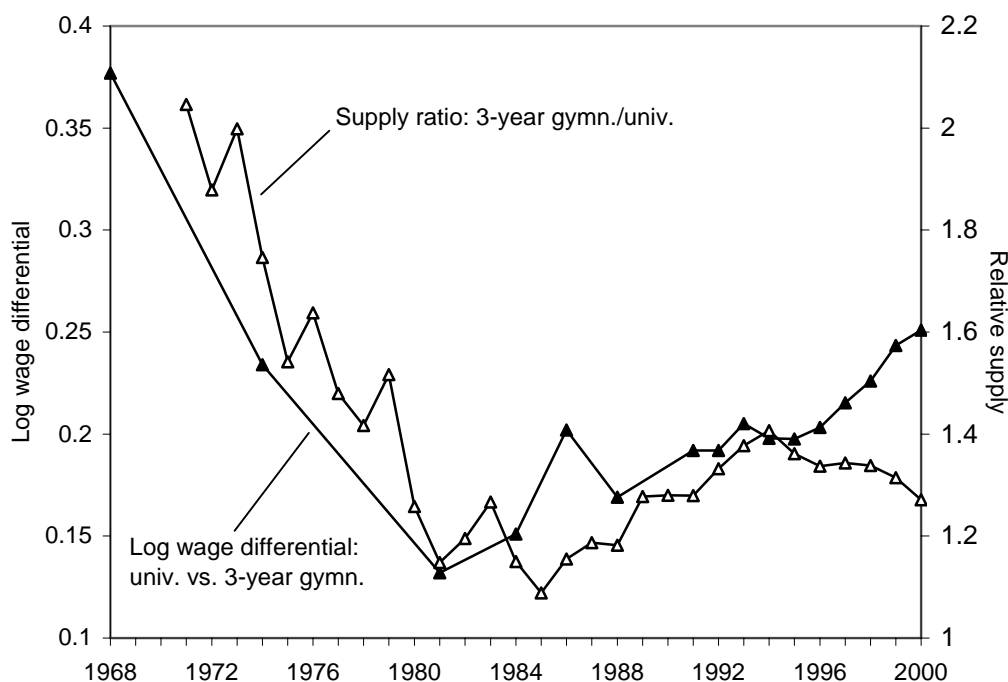


Figure 5: The supply of individuals with 3-years of gymnasium divided by the supply of individuals with at least 3-years of university 1971-2000, and log wage differential between university and 3-year gymnasium 1968-2000.

Source: Supply ratio based on the Labor Force Surveys, Statistics Sweden; see the text for details. Log wage differentials 1968-1991 are based on Edin & Holmlund (1995), Table 9.2 (12 and 15 years of education), p. 318. Log wage differentials 1992-2000 are based on Table 2 and have been re-scaled; see the text for details.

Note: Information on individuals' education is based on interviews 1971-1995 and on registers 1996-2000.

Holmlund (1995) and compare the labor force shares of individuals with 3-years of gymnasium to those with at least 3-years of university. Figure 5 graphs the ratio of the 3-years gymnasium share to the university educated share between 1971 and 2000; 2001 is omitted because of a time-series break in the statistics from then on.²²

Figure 5 also contains the university log wage premium from Figure 1, i.e. the estimate for *university* minus the estimate for *3-year gymnasium*. When the two series in Figure 5 move in the same direction, changes in the university wage premium can be interpreted as being consistent with relative supply changes.

Changes in the university wage premium match closely to changes in relative supply until the mid 1990s. After that, the increase in the university wage premium is actually

²² In 2001 the Labor Force Survey started to use improved sources to obtain individuals' highest education. These new sources are the same as are used for LINDA from 2000 onwards; see footnote 8.

accompanied by a decrease in the supply of individuals with 3-years of gymnasium relative to those with least 3-years of university, i.e. the relative supply of university educated individuals has increased.²³

While the educational composition of the labor force is probably a good approximation to a measure of supply during the 1970s and 1980s, this need not be the case during the 1990s. The deep Swedish recession during the 1990s caused a marked fall in labor force participation which could contaminate the results. Figure C1 in Appendix C contains the same relative supply series as in Figure 5 for the 1990-2000 period together with an equivalent supply series that is based on the educational composition of the Swedish population.²⁴ The magnitudes of the changes in the two series differ, but the main patterns over time are similar and the figure does not alter any conclusions. Hence, it appears like the 1990s in Sweden has seen an accelerated relative demand for university-educated workers.

6. Concluding remarks

This paper investigates changes in educational wage differentials in Sweden between 1992 and 2001 and places them in a longer-term perspective. The university wage premium has increased noticeably between 1992 and 2001 but the gymnasium premium has been stable. These results, together with previous research, suggest that the university wage premium in Sweden has exhibited a clearly positive trend since the early 1980s while the gymnasium wage premium has been constant since the mid 1970s.

Educational wage differentials have evolved differently across men and women, and between the public and the private sectors between 1992 and 2001. In particular, the increase

²³ The same conclusion is also reached if the wage differential between *some university* and *3-year gymnasium* is analyzed in the same way.

²⁴ One problem with using the educational composition of the population during the 1970s and the 1980s is the large variation in female labor force participation during this period.

in the university wage premium has been especially strong among men within the private sector but constant among women employed in the public sector.

Interestingly, the marked increase in the Swedish university wage premium during the second half of the 1990s appears to be inconsistent with changes in the relative supply of highly educated workers. Instead, it appears likely that the 1990s witnessed an accelerated relative demand for highly educated workers.

The performed documentation of changes in wage differentials in Sweden raises several questions suitable for future research. For example, why have there been heterogeneous changes in educational wage differentials across gender and between the public and private sectors? What is driving the estimated increase in the native-immigrant wage differential? Despite a hot topic in the Swedish media during the last decade, why has the wage differential between men and women been constant? Lastly, and a question that has received recent attention, why has the demand for university-educated workers in Sweden increased? Results in Lindquist (2003) suggest that capital-skill complementarity is a chief explanation for the increased demand – at least for changes among employees in the Swedish industry. Savvidou (2003) uses a different empirical framework than Lindquist (2003) and her results also lend some support for the capital-skill complementarity but indicate that this is not the sole explanation. Future research on this, as well as on the other questions cited above, is therefore warranted.

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Appendix A: An indirect test of sample composition bias

As discussed in Section 2, individuals employed in small firms are under-represented in the data before 1998. To indirectly test if this causes the estimates in the main text to be biased, I use the fact that for all individuals in LINDA, information on annual earnings is available through tax reports. Based on all individuals aged 20-64 - except students and the self employed - I estimate OLS earnings equations each year for the 1992-2001 period. I also estimate the same earnings equations with samples that include only individuals who are also present in the wage sample in a specific year. Weighted least squares is used for these restricted samples. To reduce the impact of year-to-year variations in hours worked on changes in the estimates, earnings less than twelve times the lowest monthly wage in the wage sample for a specific year is excluded in that year; in 1992 and 2001 this corresponds to 116,400 and 143,952 SEK, respectively.

Tables A1 and A2 contain the estimates for the representative and restricted samples. The magnitude of the estimates is similar to those obtained for wages, which indicates that annual earnings is a good approximation for wages – at least for the current purpose.

To better compare the estimates for the university categories from the two samples, Figure A1 graphs the estimated coefficients for *some university* and *university*; the patterns for the *doctorate* estimates are very similar to these categories and are therefore excluded. It is clear that the inclusion of more individuals from smaller firms in 1998 cause the estimates based on the restricted sample to increase between 1997 and 1998. But apart from this, the estimates based on the restricted sample correctly capture the changes over time.

Figure A2 graphs the estimates for *2-year gymnasium* and *3-year gymnasium* based on the representative and the restricted samples, respectively. It is clear that the restricted sample produces an artificial increase in the estimates between 1997 and 1998. For *3-year gymnasium*, the estimates based on the restricted sample decrease between 1992 and 1997 but

Table A1: Estimated earnings equations 1992-2001 based on representative samples

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymnasium	.040 (.002)	.040 (.002)	.038 (.002)	.035 (.002)	.035 (.002)	.032 (.002)	.032 (.002)	.032 (.002)	.032 (.002)	.030 (.002)
3-year gymnasium	.163 (.003)	.161 (.003)	.160 (.003)	.152 (.003)	.151 (.003)	.147 (.003)	.144 (.003)	.145 (.003)	.138 (.003)	.136 (.003)
some univ.	.187 (.003)	.188 (.003)	.187 (.003)	.180 (.003)	.186 (.003)	.194 (.003)	.200 (.003)	.218 (.003)	.218 (.003)	.229 (.003)
univ.	.327 (.003)	.333 (.003)	.333 (.003)	.318 (.003)	.326 (.003)	.331 (.003)	.335 (.003)	.357 (.003)	.352 (.003)	.363 (.003)
doctorate	.515 (.011)	.536 (.010)	.522 (.011)	.495 (.011)	.503 (.011)	.520 (.011)	.518 (.011)	.558 (.012)	.575 (.011)	.576 (.012)
exp	.021 (.000)	.021 (.000)	.021 (.000)	.021 (.000)	.021 (.000)	.021 (.000)	.022 (.000)	.023 (.000)	.021 (.000)	.022 (.000)
exp ² /100	-.034 (.001)	-.034 (.001)	-.035 (.001)	-.034 (.001)	-.034 (.001)	-.034 (.001)	-.036 (.001)	-.038 (.001)	-.035 (.001)	-.036 (.001)
female	-.248 (.002)	-.251 (.002)	-.252 (.002)	-.258 (.002)	-.250 (.002)	-.251 (.002)	-.262 (.002)	-.262 (.002)	-.250 (.002)	-.245 (.002)
immigrant	-.037 (.003)	-.036 (.003)	-.035 (.003)	-.037 (.003)	-.044 (.003)	-.043 (.003)	-.048 (.003)	-.058 (.003)	-.060 (.003)	-.062 (.003)
constant	11.871 (.003)	11.881 (.003)	11.908 (.004)	11.947 (.004)	11.990 (.004)	12.020 (.004)	12.035 (.004)	12.040 (.004)	12.116 (.004)	12.152 (.004)
obs.	93,057	90,399	89,118	91,906	92,097	92,103	97,905	103,570	102,266	104,063
R ²	.354	.349	.340	.331	.322	.321	.315	.313	.302	.302
σ_e	.241	.246	.251	.253	.254	.256	.268	.279	.274	.276

Note: White's (1980) heteroskedasticity robust standard errors are in parentheses.

are constant for the representative sample. This could mean that the estimates for 3-year gymnasium should decrease more than Table 2 in the main text suggests. On the other hand, this effect is not visible for any other education category and could therefore be due in some way to the use of annual earnings and/or chance.

Comparing Tables A1 and A2 reveals no marked difference in the evolution of the linear experience variable between the two samples. The absolute value of the squared experience term increases more in the restricted sample. There is, however, no difference between the samples in the change that occur between 1997 and 1998 – when sample composition should have the largest effect – making it difficult to interpret the cause of the differences.

There are some differences between the samples in the evolution of estimates for the female dummy variable. It is roughly constant for the representative sample up to 1997 but

Table A2: Estimated earnings equations 1992-2001 based on restricted samples

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.038 (.003)	.029 (.003)	.027 (.003)	.025 (.003)	.030 (.003)	.028 (.003)	.032 (.003)	.034 (.003)	.033 (.002)	.031 (.003)
3-year gymn.	.157 (.005)	.156 (.004)	.152 (.004)	.148 (.005)	.150 (.004)	.146 (.004)	.147 (.003)	.149 (.003)	.139 (.003)	.138 (.003)
some univ.	.174 (.004)	.170 (.004)	.166 (.004)	.162 (.004)	.173 (.004)	.182 (.004)	.201 (.003)	.224 (.003)	.220 (.003)	.230 (.003)
univ.	.316 (.004)	.317 (.004)	.311 (.004)	.299 (.004)	.311 (.004)	.320 (.004)	.339 (.003)	.362 (.003)	.352 (.003)	.364 (.003)
doctorate	.481 (.013)	.514 (.012)	.487 (.012)	.456 (.012)	.478 (.012)	.495 (.012)	.505 (.011)	.562 (.011)	.573 (.011)	.569 (.012)
exp	.019 (.000)	.020 (.000)	.021 (.000)	.021 (.000)	.020 (.000)	.020 (.000)	.021 (.000)	.022 (.000)	.021 (.000)	.021 (.000)
exp ² /100	-.031 (.001)	-.032 (.001)	-.035 (.001)	-.034 (.001)	-.032 (.001)	-.032 (.001)	-.034 (.001)	-.036 (.001)	-.034 (.001)	-.036 (.001)
female	-.256 (.002)	-.262 (.002)	-.274 (.002)	-.274 (.002)	-.267 (.002)	-.270 (.002)	-.269 (.002)	-.266 (.002)	-.256 (.002)	-.251 (.002)
immigrant	-.027 (.004)	-.035 (.004)	-.031 (.004)	-.032 (.004)	-.035 (.004)	-.036 (.004)	-.043 (.003)	-.058 (.003)	-.056 (.003)	-.058 (.003)
constant	11.904 (.005)	11.925 (.005)	11.961 (.005)	11.998 (.005)	12.046 (.005)	12.080 (.005)	12.053 (.004)	12.052 (.004)	12.128 (.004)	12.162 (.004)
obs.	50,728	52,023	53,025	52,360	55,188	56,491	81,183	91,819	92,373	94,957
R ²	.389	.386	.378	.360	.348	.350	.330	.326	.316	.313
σ_e	.223	.230	.239	.241	.243	.245	.260	.273	.267	.270

Note: Weighted least squares is used; see the text for details. White's (1980) heteroskedasticity robust standard errors are in parentheses.

becomes more negative for the restricted sample. The estimate for the representative sample also becomes more negative between 1997 and 1998, while the estimate for the restricted sample is constant. Hence, the changes in the estimates for the wage sample might be somewhat affected by the sample composition.

For the restricted sample, there is a larger increase in the absolute value of the immigrant dummy variable, and part of the difference occurs between 1997 and 1998. There is, however, also a marked increase in the estimates based on the representative sample; thus there is little reason to doubt that there has indeed been a marked increase in the native-immigrant wage differential during the 1990s.

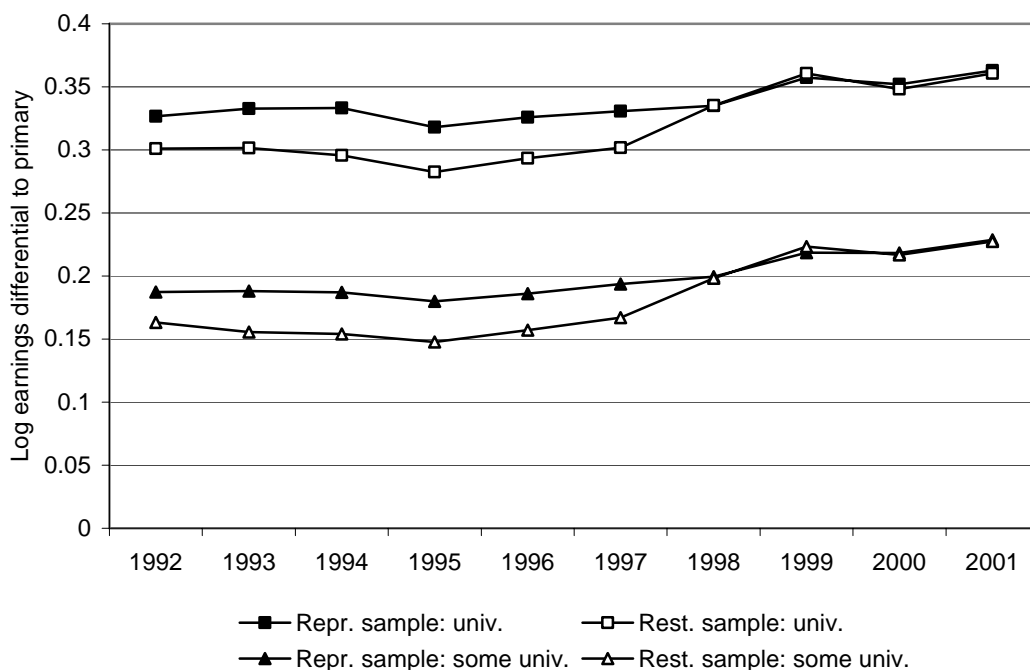


Figure A1: Log earnings differentials 1992-01 based on a representative and a restricted sample; university versus primary and some university versus primary.

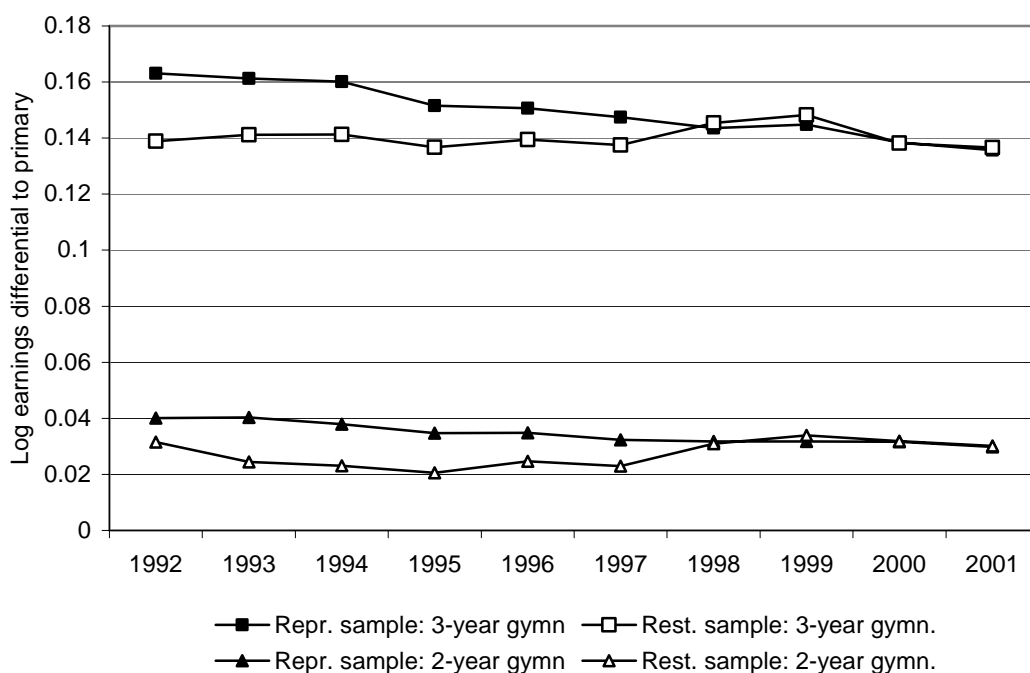


Figure A2: Log earnings differentials 1992-01 based on a representative and a restricted sample; 3-year gymnasium versus primary and 2-year gymnasium versus primary.

I have also performed the above analysis with controls for age instead of potential experience. The results imply that the restricted sample produces too large an increase in the absolute value of the estimates for the two age variables between 1997 and 1998, but that changes between other years are correct.

Appendix B: Nonseparability of wages in education and experience

I have examined my data in various ways, and there are clear signs of nonseparability in Swedish wages between education and experience (this includes estimating separate wage equations for each education category). Hence, an assumption underlying the Mincerian wage equations is unfulfilled. A natural question is then whether this affects the results in the main text concerning *changes* in educational wage differentials. One way to investigate this would be to replicate the analysis of Heckman *et al* (2003) on Swedish data. This is, however, beyond the scope of this paper. To illuminate this question, I instead take advantage of the very large sample size and compare simple mean log wage differentials within 5-year experience intervals in 1992 and 2001.

For females and males, I estimate separate weighted least squares wage equations for each 5-year interval and use the education categories as the only explanatory variables. Immigrants are excluded from the sample (including these does not change any conclusions). Table B1 presents results for selected experience intervals (results for other intervals are available on request).

I first discuss the estimates for females in Table B1. The relevant results for comparison are found in Table 3 in the main text. Consistent with the results in the main text, Table B1 displays a decrease in the estimate for *2-year gymnasium*. For *3-year gymnasium*, the estimate in the main text was roughly constant, while the results in Table B1 display a clear increase. For *university* and *some university*, the results in the main text displayed increased estimates.

Table B1: Estimates for different experience intervals in 1992 and 2001

Years of experience	5-9		15-19		25-29		35-39	
	1992	2001	1992	2001	1992	2001	1992	2001
Females								
2-year gymn.	.035 (.008)	.031 (.013)	.047 (.007)	.026 (.010)	.047 (.005)	.037 (.008)	.045 (.005)	.024 (.006)
3-year gymn.	.040 (.009)	.077 (.011)	.065 (.010)	.109 (.012)	.103 (.010)	.105 (.011)	.080 (.014)	.149 (.011)
some univ.	.108 (.009)	.189 (.014)	.158 (.008)	.199 (.012)	.201 (.007)	.197 (.010)	.211 (.009)	.220 (.008)
univ.	.220 (.011)	.298 (.012)	.294 (.009)	.351 (.013)	.350 (.007)	.327 (.009)	.340 (.008)	.344 (.008)
doctorate	.409 (.040)	.512 (.042)	.593 (.050)	.651 (.047)	.555 (.038)	.718 (.051)	.518 (.032)	.630 (.055)
constant	9.356 (.008)	9.645 (.010)	9.397 (.007)	9.732 (.010)	9.408 (.005)	9.740 (.007)	9.422 (.004)	9.746 (.005)
obs.	3,315	4,545	4,500	6,100	5,162	6,175	3,293	6,232
R ²	.283	.261	.396	.262	.498	.288	.495	.345
σ_e	.120	.195	.137	.220	.136	.203	.119	.182
Males								
2-year gymn.	.066 (.012)	.074 (.012)	.045 (.011)	.056 (.010)	.043 (.009)	.041 (.009)	.052 (.011)	.047 (.008)
3-year gymn.	.082 (.014)	.082 (.010)	.089 (.014)	.178 (.013)	.176 (.012)	.154 (.013)	.229 (.014)	.231 (.010)
some univ.	.131 (.014)	.213 (.014)	.188 (.013)	.298 (.013)	.233 (.012)	.269 (.012)	.256 (.015)	.277 (.013)
univ.	.242 (.016)	.345 (.012)	.345 (.015)	.476 (.015)	.363 (.012)	.443 (.013)	.395 (.016)	.427 (.013)
doctorate	.437 (.040)	.526 (.030)	.583 (.036)	.618 (.028)	.558 (.033)	.681 (.032)	.558 (.042)	.549 (.043)
constant	9.411 (.011)	9.762 (.009)	9.503 (.010)	9.860 (.009)	9.553 (.007)	9.890 (.007)	9.551 (.006)	9.893 (.005)
obs.	2,627	5,254	2,957	6,716	3,409	5,917	2,143	5,925
R ²	.201	.231	.334	.277	.323	.274	.377	.242
σ_e	.177	.229	.206	.261	.212	.279	.204	.270

Note: Weighted least squares is used; see the text for details. White's (1980) heteroskedasticity robust standard errors are in parentheses.

In Table B1 there is similarly a marked increase in both these categories for females with 5-9 and 15-19 years of potential experience but the estimate for *university* decreases for females with 25-29 years of experience and the other estimates are constant. In line with the results in the main text, the estimate for *doctorate* in Table 5 increases markedly for all experience intervals.

Next in Table B1 are the results for males. The results in the main text displayed a fall in the estimate for *2-year gymnasium*, while the changes in Table B1 indicate differences between experience groups. The estimates in the main text for *3-year gymnasium* were roughly constant; Table B1 is consistent with this, except for males with 15-19 year of experience where the estimate is twice as large in 2001 compared to 1992. This is, however, the only experience group where such an increase occurs. The increases for the three higher education categories in Table B1 are consistent with the results in the main text, even though the increases for the two university categories are larger for individuals with 5-9 and 15-19 years of experience.

Overall, the results in Table B1 highlight some heterogeneity in the changes across experience groups. However, except for the changes in the returns to 3-year gymnasium for females, the results in Table B1 do not differ systematically from those obtained from Mincerian wage equations. I therefore conclude that the estimated Mincerian wage equations capture the main patterns in the data.

Appendix C: Additional tables and figures

Table C1: Estimated wage equations 1992-2001, public sector dummy but without weights

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.064 (.002)	.067 (.002)	.070 (.002)	.066 (.002)	.060 (.002)	.057 (.002)	.058 (.002)	.059 (.002)	.058 (.002)	.054 (.002)
3-year gymn.	.135 (.003)	.144 (.003)	.146 (.003)	.140 (.003)	.136 (.003)	.132 (.003)	.143 (.003)	.150 (.003)	.144 (.002)	.144 (.003)
some univ.	.202 (.002)	.207 (.003)	.205 (.003)	.202 (.003)	.198 (.003)	.207 (.003)	.234 (.003)	.253 (.003)	.258 (.003)	.267 (.003)
univ.	.343 (.003)	.358 (.003)	.352 (.003)	.349 (.003)	.349 (.003)	.358 (.003)	.387 (.003)	.407 (.003)	.410 (.003)	.424 (.003)
doctorate	.538 (.009)	.563 (.009)	.551 (.009)	.552 (.009)	.554 (.009)	.570 (.009)	.587 (.009)	.639 (.009)	.652 (.009)	.665 (.009)
exp	.016 (.000)	.016 (.000)	.018 (.000)	.018 (.000)	.018 (.000)	.019 (.000)	.021 (.000)	.021 (.000)	.020 (.000)	.021 (.000)
exp ² /100	-.024 (.000)	-.024 (.001)	-.027 (.001)	-.027 (.001)	-.027 (.001)	-.029 (.001)	-.031 (.001)	-.032 (.001)	-.031 (.001)	-.032 (.000)
female	-.124 (.002)	-.138 (.002)	-.136 (.002)	-.124 (.002)	-.137 (.002)	-.137 (.002)	-.138 (.002)	-.133 (.001)	-.138 (.002)	-.135 (.002)
immigrant	-.049 (.003)	-.050 (.003)	-.053 (.003)	-.052 (.003)	-.051 (.003)	-.056 (.003)	-.065 (.003)	-.081 (.003)	-.083 (.002)	-.085 (.002)
public sector	-.087 (.002)	-.074 (.002)	-.090 (.002)	-.109 (.002)	-.119 (.002)	-.125 (.002)	-.111 (.002)	-.106 (.002)	-.122 (.002)	-.126 (.002)
constant	9.367 (.003)	9.367 (.003)	9.401 (.003)	9.424 (.003)	9.495 (.003)	9.525 (.003)	9.506 (.003)	9.521 (.003)	9.587 (.003)	9.623 (.003)
obs.	62,418	63,378	63,688	62,731	65,337	65,992	91,952	102,330	105,633	108,057
R ²	.450	.447	.421	.424	.424	.428	.372	.367	.371	.376
σ_e	.167	.174	.183	.180	.184	.187	.211	.222	.224	.228

Note: White's (1980) standard errors are in parentheses.

Table C2: Estimated wage equations 1992-2001 with controls for age

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymnasium	.047 (.002)	.051 (.002)	.051 (.002)	.044 (.003)	.038 (.003)	.036 (.002)	.041 (.002)	.045 (.002)	.042 (.002)	.039 (.002)
3-year gymnasium	.134 (.004)	.143 (.004)	.143 (.004)	.133 (.004)	.124 (.004)	.116 (.003)	.126 (.003)	.134 (.003)	.130 (.003)	.130 (.003)
some university	.166 (.003)	.178 (.003)	.174 (.003)	.167 (.003)	.159 (.003)	.165 (.003)	.192 (.003)	.214 (.003)	.220 (.003)	.229 (.003)
university	.303 (.003)	.323 (.003)	.315 (.003)	.306 (.003)	.303 (.003)	.306 (.003)	.324 (.003)	.347 (.003)	.351 (.003)	.365 (.003)
doctorate	.449 (.011)	.487 (.010)	.462 (.010)	.451 (.010)	.456 (.010)	.466 (.010)	.479 (.009)	.542 (.009)	.553 (.009)	.565 (.009)
age	.023 (.001)	.024 (.001)	.027 (.001)	.027 (.001)	.026 (.001)	.027 (.001)	.030 (.001)	.032 (.000)	.031 (.000)	.032 (.000)
age ² /100	-.022 (.001)	-.023 (.001)	-.026 (.001)	-.026 (.001)	-.025 (.001)	-.026 (.001)	-.029 (.001)	-.031 (.001)	-.031 (.001)	-.033 (.001)
female	-.164 (.002)	-.167 (.002)	-.171 (.002)	-.164 (.002)	-.178 (.002)	-.180 (.002)	-.173 (.001)	-.164 (.001)	-.175 (.001)	-.174 (.001)
immigrant	-.042 (.003)	-.048 (.003)	-.049 (.003)	-.046 (.003)	-.044 (.003)	-.047 (.003)	-.060 (.003)	-.077 (.003)	-.078 (.002)	-.081 (.002)
constant	9.030 (.012)	8.998 (.012)	8.993 (.013)	9.024 (.013)	9.112 (.013)	9.127 (.013)	9.071 (.010)	9.049 (.010)	9.123 (.010)	9.138 (.010)
obs.	62,418	63,378	63,688	62,731	65,337	65,992	91,952	102,330	105,633	108,057
R ²	.388	.396	.376	.359	.352	.354	.332	.334	.329	.334
σ_e	.184	.190	.198	.199	.203	.205	.219	.229	.233	.236

Note: Weighted least squares is used; see the text for details. White's (1980) robust standard errors are in parentheses.

Table C3: Estimated wage equations 1992-2001 for females

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymnasium	.037 (.003)	.040 (.003)	.042 (.003)	.031 (.003)	.018 (.003)	.018 (.003)	.027 (.002)	.031 (.002)	.027 (.002)	.027 (.002)
3-year gymnasium	.093 (.005)	.101 (.004)	.101 (.004)	.097 (.005)	.090 (.005)	.086 (.004)	.096 (.004)	.104 (.003)	.106 (.003)	.106 (.003)
some university	.161 (.003)	.160 (.003)	.152 (.004)	.139 (.004)	.129 (.004)	.139 (.004)	.159 (.003)	.181 (.003)	.184 (.003)	.198 (.003)
university	.294 (.004)	.308 (.004)	.293 (.004)	.280 (.004)	.274 (.004)	.277 (.004)	.290 (.003)	.313 (.003)	.317 (.003)	.336 (.003)
doctorate	.520 (.021)	.560 (.018)	.516 (.018)	.527 (.018)	.538 (.018)	.541 (.018)	.515 (.016)	.583 (.015)	.613 (.017)	.618 (.017)
exp	.011 (.000)	.012 (.000)	.013 (.000)	.013 (.000)	.014 (.000)	.014 (.000)	.015 (.000)	.016 (.000)	.016 (.000)	.016 (.000)
exp ² /100	-.017 (.001)	-.019 (.001)	-.021 (.001)	-.021 (.001)	-.022 (.001)	-.023 (.001)	-.025 (.001)	-.025 (.001)	-.026 (.001)	-.026 (.001)
immigrant	-.022 (.004)	-.026 (.003)	-.025 (.004)	-.031 (.004)	-.025 (.004)	-.029 (.004)	-.036 (.003)	-.051 (.003)	-.051 (.003)	-.055 (.003)
constant	9.277 (.004)	9.270 (.004)	9.308 (.005)	9.347 (.005)	9.400 (.005)	9.432 (.005)	9.429 (.004)	9.449 (.004)	9.493 (.004)	9.532 (.004)
obs.	36,992	37,524	36,917	36,405	37,155	37,230	44,817	49,528	52,111	53,239
R ²	.349	.353	.302	.294	.286	.290	.274	.290	.282	.300
σ_e	.146	.151	.162	.163	.169	.172	.184	.192	.200	.202

Note: Weighted least squares is used; see the text for details. White's (1980) standard errors are in parentheses.

Table C4: Estimated wage equations 1992-2001 for males

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.061 (.004)	.064 (.004)	.059 (.004)	.057 (.004)	.053 (.004)	.049 (.004)	.051 (.003)	.055 (.003)	.052 (.003)	.044 (.003)
3-year gymn.	.166 (.005)	.180 (.005)	.178 (.005)	.165 (.005)	.153 (.005)	.143 (.005)	.153 (.004)	.163 (.004)	.153 (.004)	.150 (.004)
some univ.	.196 (.005)	.221 (.005)	.220 (.005)	.216 (.006)	.205 (.005)	.207 (.005)	.241 (.004)	.265 (.004)	.268 (.004)	.269 (.004)
univ.	.348 (.006)	.380 (.005)	.380 (.006)	.372 (.006)	.365 (.006)	.369 (.006)	.396 (.005)	.424 (.004)	.423 (.005)	.427 (.005)
doctorate	.489 (.012)	.531 (.011)	.520 (.011)	.503 (.011)	.498 (.011)	.507 (.011)	.533 (.011)	.594 (.011)	.585 (.011)	.596 (.011)
exp	.021 (.001)	.022 (.001)	.023 (.001)	.023 (.001)	.021 (.001)	.022 (.001)	.022 (.000)	.023 (.000)	.022 (.000)	.023 (.000)
exp ² /100	-.032 (.001)	-.033 (.001)	-.036 (.001)	-.035 (.001)	-.032 (.001)	-.033 (.001)	-.034 (.001)	-.036 (.001)	-.035 (.001)	-.038 (.001)
immigrant	-.065 (.005)	-.075 (.005)	-.076 (.005)	-.064 (.005)	-.067 (.005)	-.069 (.005)	-.086 (.004)	-.105 (.004)	-.108 (.004)	-.110 (.004)
constant	9.285 (.006)	9.274 (.006)	9.298 (.006)	9.336 (.006)	9.427 (.006)	9.465 (.006)	9.459 (.005)	9.465 (.005)	9.541 (.005)	9.577 (.005)
obs.	25,426	25,854	26,771	26,326	28,182	28762	47135	52802	53522	54818
R ²	.305	.322	.314	.307	.283	.283	.287	.299	.287	.288
σ_e	.213	.220	.225	.224	.227	.229	.244	.255	.257	.262

Note: Weighted least squares is used; see the text for details. White's (1980) standard errors are in parentheses.

Table C5: Estimated wage equations 1992-2001 for the private sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.061 (.004)	.064 (.004)	.062 (.004)	.059 (.004)	.052 (.004)	.051 (.004)	.053 (.003)	.054 (.003)	.052 (.003)	.048 (.003)
3-year gymn.	.169 (.005)	.180 (.005)	.174 (.005)	.162 (.005)	.150 (.005)	.140 (.005)	.155 (.003)	.161 (.003)	.153 (.003)	.153 (.003)
some univ.	.208 (.006)	.235 (.006)	.231 (.006)	.232 (.006)	.221 (.006)	.224 (.006)	.267 (.004)	.282 (.004)	.286 (.004)	.286 (.004)
univ.	.407 (.008)	.441 (.007)	.438 (.007)	.434 (.007)	.432 (.007)	.429 (.006)	.457 (.005)	.468 (.005)	.475 (.005)	.478 (.004)
doctorate	.563 (.033)	.634 (.024)	.618 (.021)	.559 (.018)	.565 (.018)	.573 (.018)	.558 (.019)	.638 (.017)	.649 (.019)	.657 (.019)
exp	.020 (.001)	.021 (.000)	.023 (.001)	.023 (.001)	.022 (.000)	.022 (.000)	.023 (.000)	.024 (.000)	.023 (.000)	.023 (.000)
exp ² /100	-.031 (.001)	-.033 (.001)	-.036 (.001)	-.036 (.001)	-.035 (.001)	-.035 (.001)	-.035 (.001)	-.037 (.001)	-.036 (.001)	-.038 (.001)
female	-.159 (.003)	-.156 (.003)	-.159 (.003)	-.142 (.003)	-.146 (.003)	-.152 (.003)	-.141 (.002)	-.139 (.002)	-.144 (.002)	-.142 (.002)
immigrant	-.053 (.005)	-.064 (.005)	-.062 (.005)	-.057 (.005)	-.058 (.005)	-.060 (.005)	-.074 (.004)	-.090 (.003)	-.098 (.003)	-.102 (.003)
constant	9.330 (.006)	9.308 (.006)	9.332 (.006)	9.364 (.006)	9.444 (.006)	9.485 (.006)	9.474 (.004)	9.486 (.004)	9.558 (.004)	9.595 (.004)
obs.	16,117	19,901	19,821	19,863	21,720	23,543	51,096	59,797	61,035	63,380
R ²	.375	.382	.387	.373	.359	.360	.331	.334	.336	.332
σ_e	.198	.206	.210	.209	.210	.212	.232	.243	.242	.248

Note: Weighted least squares is used; see the text for details. White's (1980) standard errors are in parentheses.

Table C6: Estimated wage equations 1992-2001 for the public sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.061 (.002)	.061 (.002)	.071 (.002)	.067 (.002)	.062 (.002)	.058 (.002)	.055 (.003)	.056 (.003)	.053 (.003)	.052 (.003)
3-year gymn.	.113 (.003)	.114 (.003)	.126 (.003)	.123 (.003)	.126 (.003)	.128 (.004)	.122 (.004)	.130 (.004)	.129 (.004)	.127 (.004)
some univ.	.196 (.002)	.191 (.003)	.192 (.003)	.188 (.003)	.187 (.003)	.198 (.003)	.194 (.003)	.212 (.003)	.218 (.003)	.236 (.003)
univ.	.331 (.003)	.335 (.003)	.331 (.003)	.327 (.003)	.327 (.003)	.338 (.003)	.335 (.003)	.352 (.003)	.352 (.003)	.374 (.003)
doctorate	.532 (.010)	.544 (.010)	.537 (.010)	.545 (.010)	.547 (.010)	.567 (.010)	.574 (.010)	.621 (.010)	.633 (.010)	.652 (.010)
exp	.014 (.000)	.014 (.000)	.015 (.000)	.015 (.000)	.016 (.000)	.017 (.000)	.018 (.000)	.018 (.000)	.018 (.000)	.017 (.000)
exp ² /100	-.021 (.001)	-.020 (.001)	-.022 (.001)	-.022 (.001)	-.023 (.001)	-.025 (.001)	-.026 (.001)	-.026 (.001)	-.026 (.001)	-.026 (.001)
female	-.113 (.002)	-.131 (.002)	-.127 (.002)	-.117 (.002)	-.134 (.002)	-.129 (.002)	-.138 (.002)	-.127 (.002)	-.135 (.002)	-.128 (.002)
immigrant	-.047 (.003)	-.041 (.003)	-.049 (.003)	-.049 (.003)	-.048 (.003)	-.053 (.003)	-.054 (.004)	-.070 (.004)	-.062 (.004)	-.062 (.003)
constant	9.298 (.003)	9.327 (.004)	9.348 (.004)	9.351 (.004)	9.408 (.004)	9.425 (.004)	9.449 (.005)	9.474 (.005)	9.512 (.005)	9.541 (.005)
Obs.	46,301	43,477	43,867	42,868	43,617	42,449	40,856	42,533	44,598	44,677
R ²	.483	.494	.439	.448	.446	.454	.441	.431	.424	.443
σ_e	.153	.156	.168	.162	.168	.170	.178	.187	.193	.193

Note: Weighted least squares is used; see the text for details. White's (1980) standard errors are in parentheses.

Table C7: Estimated wage equations 1992-2001 for females in the public sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.055 (.002)	.056 (.002)	.065 (.002)	.058 (.002)	.053 (.002)	.054 (.002)	.050 (.003)	.051 (.003)	.048 (.003)	.050 (.003)
3-year gymn.	.075 (.003)	.083 (.004)	.090 (.004)	.086 (.004)	.091 (.004)	.099 (.004)	.095 (.004)	.098 (.004)	.100 (.004)	.102 (.004)
some univ.	.181 (.003)	.173 (.003)	.170 (.003)	.160 (.003)	.163 (.003)	.180 (.003)	.178 (.003)	.195 (.003)	.201 (.003)	.226 (.003)
univ.	.313 (.003)	.317 (.003)	.304 (.003)	.295 (.003)	.297 (.003)	.311 (.003)	.310 (.003)	.324 (.004)	.327 (.004)	.354 (.004)
doctorate	.510 (.017)	.544 (.017)	.519 (.018)	.523 (.018)	.542 (.018)	.553 (.018)	.551 (.017)	.593 (.016)	.629 (.017)	.653 (.017)
exp	.011 (.000)	.011 (.000)	.012 (.000)	.012 (.000)	.013 (.000)	.014 (.000)	.015 (.000)	.015 (.000)	.015 (.000)	.015 (.000)
exp ² /100	-.016 (.000)	-.016 (.001)	-.017 (.001)	-.018 (.001)	-.019 (.001)	-.021 (.001)	-.023 (.001)	-.022 (.001)	-.022 (.001)	-.022 (.001)
immigrant	-.033 (.003)	-.027 (.003)	-.032 (.003)	-.032 (.003)	-.031 (.003)	-.035 (.003)	-.034 (.004)	-.046 (.004)	-.042 (.004)	-.044 (.003)
constant	9.244 (.003)	9.251 (.004)	9.280 (.004)	9.296 (.004)	9.334 (.004)	9.348 (.004)	9.365 (.005)	9.402 (.005)	9.426 (.005)	9.457 (.005)
obs.	31,463	30,098	29,686	29,191	29,707	28,973	27,560	28,454	30,451	30,874
R ²	.454	.443	.373	.393	.390	.411	.398	.404	.397	.430
σ_e	.128	.133	.145	.137	.142	.143	.150	.156	.162	.163

Note: White's (1980) standard errors are in parentheses.

Table C8: Estimated wage equations 1992-2001 for females in the private sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymm.	.042 (.005)	.049 (.005)	.046 (.005)	.041 (.005)	.031 (.006)	.030 (.005)	.040 (.004)	.042 (.004)	.042 (.004)	.040 (.004)
3-year gymm.	.116 (.008)	.126 (.007)	.116 (.007)	.117 (.007)	.105 (.007)	.092 (.007)	.110 (.005)	.121 (.005)	.122 (.005)	.123 (.005)
some univ.	.183 (.010)	.199 (.009)	.187 (.009)	.196 (.009)	.176 (.009)	.175 (.009)	.213 (.007)	.225 (.006)	.232 (.006)	.231 (.006)
univ.	.332 (.013)	.374 (.011)	.370 (.011)	.380 (.011)	.376 (.011)	.356 (.010)	.382 (.008)	.392 (.007)	.404 (.007)	.419 (.007)
doctorate	.667 (.057)	.699 (.037)	.653 (.036)	.640 (.032)	.644 (.031)	.627 (.032)	.590 (.037)	.685 (.034)	.709 (.036)	.676 (.039)
exp	.013 (.001)	.016 (.001)	.017 (.001)	.017 (.001)	.019 (.001)	.018 (.001)	.019 (.001)	.020 (.001)	.020 (.001)	.020 (.001)
exp ² /100	-.022 (.001)	-.025 (.001)	-.028 (.001)	-.029 (.002)	-.031 (.002)	-.030 (.001)	-.032 (.001)	-.033 (.001)	-.033 (.001)	-.033 (.001)
immigrant	-.018 (.007)	-.034 (.006)	-.029 (.007)	-.043 (.007)	-.037 (.007)	-.042 (.007)	-.051 (.006)	-.066 (.005)	-.072 (.005)	-.077 (.005)
constant	9.275 (.009)	9.244 (.008)	9.285 (.009)	9.325 (.010)	9.378 (.010)	9.425 (.009)	9.413 (.007)	9.430 (.006)	9.481 (.006)	9.526 (.006)
obs.	5,529	7,426	7,231	7,214	7,448	8,257	17,257	21,074	21,660	22,365
R ²	.252	.288	.279	.289	.295	.276	.252	.263	.277	.281
σ_e	.165	.171	.179	.184	.187	.190	.210	.220	.224	.228

Note: White's (1980) standard errors are in parentheses.

Table C9: Estimated wage equations 1992-2001 for males in the private sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymm.	.073 (.005)	.074 (.005)	.072 (.005)	.070 (.005)	.063 (.005)	.063 (.005)	.061 (.003)	.062 (.003)	.059 (.003)	.053 (.003)
3-year gymm.	.184 (.007)	.202 (.007)	.196 (.006)	.177 (.006)	.168 (.006)	.161 (.006)	.172 (.004)	.178 (.004)	.166 (.004)	.167 (.004)
some univ.	.220 (.008)	.253 (.008)	.252 (.007)	.249 (.008)	.241 (.007)	.248 (.007)	.292 (.005)	.311 (.005)	.314 (.005)	.314 (.005)
univ.	.438 (.010)	.473 (.009)	.469 (.009)	.459 (.009)	.458 (.008)	.465 (.008)	.493 (.006)	.507 (.006)	.516 (.006)	.511 (.006)
doctorate	.524 (.038)	.608 (.029)	.612 (.025)	.537 (.020)	.544 (.020)	.555 (.020)	.555 (.021)	.631 (.019)	.633 (.021)	.656 (.021)
exp	.023 (.001)	.024 (.001)	.027 (.001)	.026 (.001)	.024 (.001)	.025 (.001)	.024 (.000)	.026 (.000)	.024 (.000)	.025 (.000)
exp ² /100	-.035 (.001)	-.038 (.001)	-.041 (.001)	-.040 (.001)	-.037 (.001)	-.038 (.001)	-.037 (.001)	-.040 (.001)	-.038 (.001)	-.041 (.001)
immigrant	-.073 (.007)	-.085 (.007)	-.084 (.007)	-.067 (.007)	-.072 (.007)	-.074 (.006)	-.088 (.005)	-.106 (.005)	-.115 (.004)	-.118 (.004)
constant	9.280 (.008)	9.260 (.008)	9.272 (.008)	9.311 (.008)	9.407 (.008)	9.440 (.008)	9.438 (.005)	9.445 (.005)	9.525 (.005)	9.558 (.005)
obs.	10,588	12,475	12,590	12,649	14,272	15,286	33,839	38,723	39,375	41,015
R ²	.321	.331	.345	.341	.321	.324	.314	.321	.321	.316
σ_e	.212	.223	.224	.221	.221	.221	.242	.253	.251	.257

Note: White's (1980) standard errors are in parentheses.

Table C10: Estimated wage equations 1992-2001 for males in the public sector

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
2-year gymn.	.067 (.004)	.067 (.005)	.071 (.005)	.073 (.005)	.072 (.005)	.059 (.006)	.057 (.006)	.060 (.006)	.059 (.006)	.050 (.006)
3-year gymn.	.156 (.006)	.154 (.006)	.169 (.006)	.172 (.006)	.170 (.006)	.165 (.007)	.156 (.007)	.172 (.007)	.170 (.007)	.165 (.007)
some univ.	.226 (.005)	.228 (.006)	.235 (.006)	.242 (.006)	.234 (.006)	.234 (.006)	.226 (.007)	.242 (.007)	.249 (.007)	.257 (.007)
univ.	.362 (.005)	.371 (.006)	.379 (.006)	.387 (.006)	.382 (.006)	.387 (.007)	.378 (.007)	.401 (.007)	.397 (.007)	.413 (.007)
doctorate	.568 (.012)	.575 (.012)	.577 (.012)	.592 (.012)	.585 (.012)	.602 (.012)	.610 (.013)	.659 (.013)	.660 (.013)	.673 (.014)
exp	.022 (.001)	.021 (.001)	.022 (.001)	.022 (.001)	.022 (.001)	.022 (.001)	.023 (.001)	.023 (.001)	.023 (.001)	.023 (.001)
exp ² /100	-.032 (.001)	-.029 (.001)	-.032 (.001)	-.031 (.001)	-.032 (.001)	-.033 (.001)	-.034 (.001)	-.033 (.002)	-.033 (.002)	-.033 (.002)
immigrant	-.078 (.007)	-.075 (.007)	-.086 (.007)	-.085 (.007)	-.084 (.007)	-.094 (.008)	-.099 (.008)	-.115 (.008)	-.108 (.008)	-.105 (.008)
constant	9.175 (.007)	9.206 (.008)	9.228 (.008)	9.229 (.008)	9.284 (.008)	9.320 (.009)	9.347 (.009)	9.375 (.009)	9.415 (.009)	9.452 (.010)
obs.	14,838	13,379	14,181	13,677	13,910	13,476	13,296	14,079	14,147	13,803
R ²	.420	.424	.393	.411	.380	.388	.368	.373	.351	.370
σ_e	.192	.193	.205	.201	.211	.214	.224	.234	.244	.246

Note: White's (1980) standard errors are in parentheses.

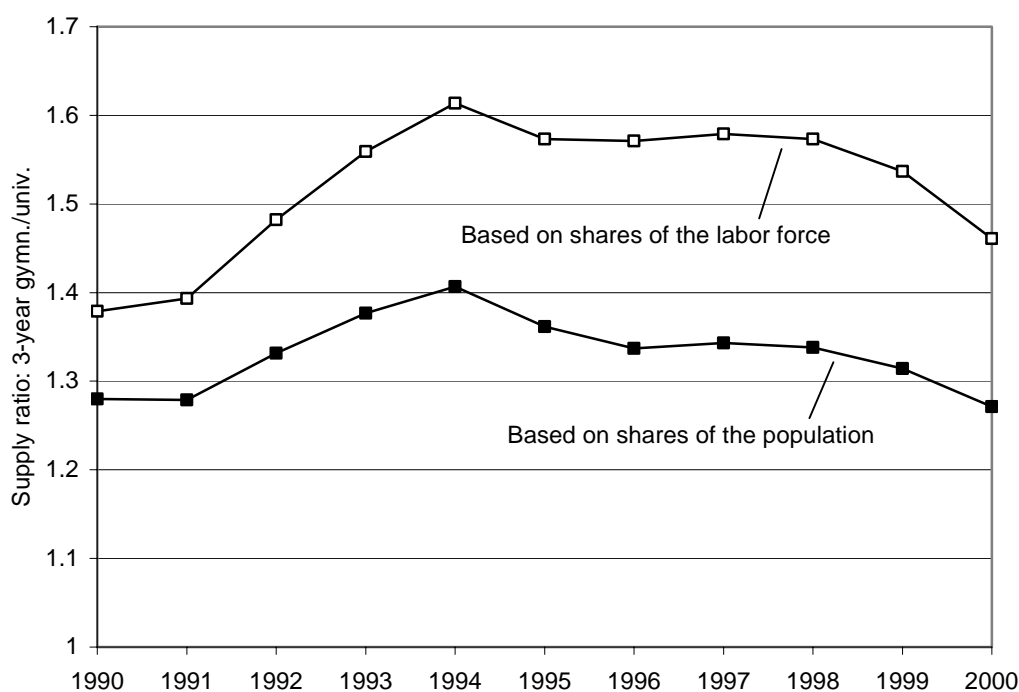


Figure C1: Supply ratio of 3-year gymnasium and university 1990-2000 based on shares of the population and shares of the labor force, respectively; see the text for details.

Source: Labor Force Surveys, Statistics Sweden.

Note: Information on individuals' education is based on interviews 1990-1995 and on registers 1996-2000.

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