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The Wage Premium on Tertiary Education: New Estimates for 21 OECD Countries

by

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Introduction and main findings

The accumulation of human capital through education and training is widely recognised as an important driver of economic growth.[1] Yet, as the decision to continue schooling beyond the secondary level is voluntary, it depends not only on talent and inclination but also on the balance of costs of and benefits from post-secondary education. Therefore, assessing the returns on education is a key input for policymakers who want to bolster a country's endowment with human capital through an increase in educational attainment.[2]

This study focuses on the single most important component of the private return on tertiary education, the gross wage premium. There are at least two additional reasons for paying particular attention to wage *premia*. First, the wage premium earned by existing graduates is easy to observe, so high-school leavers can be assumed to take it into account when deciding for or against enrolment in tertiary education. Second, to the extent that wages reflect marginal labour productivity, estimates of wage *premia* are sometimes used to assess the quality of human capital in an economy with a view to correcting simpler measures based on years of schooling or attainment levels.

The study follows an augmented *Mincerian* wage equation framework with the gross hourly wage as the dependent variable, estimated on individual cross-sections. The latter are obtained from household data covering 2 to 14 survey waves for 21 OECD countries. The time period runs from 1991-2004 for the United Kingdom, from 1994-2004 for the United States and from 1994-2001 for most other countries. The traditional Mincer equation is augmented by a number of labour market-related control variables, *inter alia* for wage earners that are overqualified or under-qualified in their current occupation.

The estimations are country-specific and draw from a common sample of men and women. Over and above the usual gender dummy in the equations, the variables for education and labour-market experience are interacted with the gender dummy, thereby obtaining gender-specific results for the tertiary education wage premium, the wage penalty on not completing upper-secondary education, and the annual labour market experience premium. The results highlight huge cross-country differences. The gross wage premium to tertiary education ranges from 27% for Spanish men to 90% for Hungarian and US degree holders. Cross-country variation remains high even after accounting for the average duration of tertiary studies. The gross wage premium *per annum* of tertiary education is found to lie in an interval from 5.5% for men in Greece and Spain as well as for women in Austria and Italy, to 17% for men and women in Hungary and the United States, and for women in Ireland and Portugal.

The study is structured as follows. The second section provides a brief discussion of methodological issues raised in the microeconomic literature on returns to education in order to highlight the value-added of this contribution and its (data-related) limitations. The third section presents the empirical specification of the *Mincerian* wage equation and in the fourth, data sources and the construction of variables are described. The penultimate section presents the results, discussed in comparison with earlier estimates and the final section concludes.

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Methodological issues related to the estimation of educational wage *premia*

Most studies on returns to education use *Mincerian* equations. The latter relate the log of earnings to the number of completed years of schooling and experience (often as a quadratic term).[3] While Mincer (1958) considers the wage premium to be just a compensation for working in jobs requiring longer education (the net present values of earnings streams net of education costs being identical for all levels of education), Mincer (1974) derives a similar empirical specification from a full human-capital model building on the theoretical work by Becker (1964) and Ben-Porath (1967).[4] The original Mincer equation assumes a linear effect on earnings of each year of education regardless of the attainment level. This study, however, allows for differential effects of upper-secondary and tertiary education.

There are a number of issues to be borne in mind when relating the *Mincerian* schooling coefficient to the causal effect of schooling on earnings (Card, 1999 and Harmon *et al.*, 2003). First, as an investment-decision variable, years of schooling and education attainment should be considered as endogenous, implying a possible bias in OLS estimates of the schooling coefficient. The *endogeneity bias* may arise either from unobserved variation in ability or from unobserved heterogeneity. If those who extend education beyond compulsory schooling have greater ability than others, the estimated Mincer coefficient is biased upwards since part of the productivity differential is actually due to innate abilities or skills acquired outside school (*ability bias*). The ability bias may interact with heterogeneous subjective discount rates that result in under-estimating the true effect of schooling on earnings if the more impatient individuals happen to be the more able ones (*heterogeneity bias*). The total direction of bias in OLS estimates is ambiguous. There is a whole strand of the empirical literature dealing with the endogeneity bias, namely by using instrumental variables (*e.g.* parents' education). This option could not be followed in this study due to lack of data. Nonetheless, the consensus from the empirical literature is that this bias in the estimated *Mincerian* wage premium is likely to be small (*e.g.* see Card, 1999, and Woessmann, 2003).

Second, if there is *measurement error* in the education variable (one year of tertiary schooling representing different stocks of human capital accumulated depending on school quality and individual characteristics), the schooling coefficient will be biased downward.[5]

Third, there is also a potential endogeneity bias related to labour supply effects. Indeed, every new graduate adds to the pool of skilled workers, thereby making skilled labour less scarce and lowering the wage *premium* that triggered the investment decision.[6]

Finally, Heckman *et al.* (2005) point out that using *ex post* estimates of earnings-schooling profiles of existing workers as a decision tool for today's investment decision requires stationarity of earnings across cohorts in the labour market. The latter is rejected for the United States on the basis of 1980 and 1990 Census data. However, in this study, a full-fledged cohort analysis is not feasible due to the limited time coverage of the available household panel data.

Empirical specification

Tertiary-education wage *premia* are obtained by country and year from individual earnings data following the *Mincerian* approach. Estimates are based on household-level data for three educational attainment levels (less than upper-secondary education, completed upper secondary education, completed tertiary education). The estimation is based on hourly wages, which reflect the impact of education on productivity. Monthly or annual wages would in addition capture the effect of individuals' decisions on working hours. Given the only weak (positive) correlation between working time and educational attainment it is reasonable to assume that the choice of hours worked reflects individual preferences rather than education levels. Experience is proxied by the number of years in the labour market rather than age, because this allows better disentangling education from experience effects.

Household-level data allow controlling for a number of individual characteristics that potentially affect earnings but are not directly related to tertiary education. Failing to control for these characteristics may induce statistical bias when estimating the effect of tertiary education. They include gender, marital status, job tenure (in years), the type of work contract and working in the public *versus* the private sector. The estimates also control for the size of the production unit ("plant size") as it is a well-established empirical fact that large firms tend to pay higher wages than small firms. It is true that to the extent that these other labour-market outcomes are dependent on education itself, their inclusion as control variables would be more controversial since they could "blur" the look at the unconditional private returns to education. However, the correlation between attainment and non-wage labour market outcomes is very limited.[7] The final control variables are over- and under-qualification of individuals in their current occupation. The risk that in a given year individuals may work in a job that does not correspond to their educational attainment is not necessarily relevant for their decision to enrol in tertiary education, hence the mismatch between education and occupational status should be treated separately from the education wage premium. Indeed, the available evidence suggests that the majority of over-qualified individuals tend to move up over time into an occupational status corresponding to their educational attainment (Dumont, 2005). On balance, controlling for over- and under-qualification tends to increase the estimated wage *premia*.

The econometric specification is as follows (individual indices are omitted for simplicity):

$$\begin{aligned} Log(hrw) = c &+ \alpha_1 \cdot edu1 + \alpha_2 \cdot edu3 + \alpha_3 \cdot edu1 \cdot woman + \alpha_4 \cdot edu3 \cdot woman + \\ &+ \beta_1 \cdot exper + \beta_2 \cdot woman + \beta_3 \cdot exper \cdot woman + \\ &+ \beta_4 \cdot married + \beta_5 \cdot public + \beta_6 \cdot part\_time + \beta_7 \cdot tenure + \beta_8 \cdot indef\_cont + \\ &+ \delta_1 \cdot Log(plant\_size) + \delta_2 \cdot overqualif + \delta_3 \cdot underqualif + \varepsilon \end{aligned} \quad [1]$$

where:

hrw = gross hourly wage

edu1, *edu3* = dummies for less-than-upper-secondary and tertiary education attainment, respectively

exper = number of years of experience in the labour market

married = dummy for marital status

public = dummy for public sector job

part_time = dummy for part-time worker

tenure = number of years with the same employer

indef_cont = dummy for indefinite-term contract

plant_size = number of employees in the individual's production unit

overqualif, *underqualif* = dummies for over- and under-qualification, respectively.

The equation above is estimated on individual cross-sections (country by country and year by year) rather than a panel mainly for three reasons. First, the *Mincerian* approach is cross-sectional in nature insofar as the variables of the equation usually show little variation over time. A panel approach would require augmenting the model with time-varying variables such as unemployment rates at a disaggregated level (gender/sector/occupation/attainment-specific) that are not readily available in the datasets exploited here. Second, the focus of this study is on the returns to education for countries as a whole rather than changes in individual conditions over time. Third, pooling data over time is sometimes warranted in order to increase the efficiency of the estimation but this argument is not compelling here given the already large size of the country-year samples.

A methodological issue raised in the literature is that the sample of wage earners may be a non-random selection of the overall sample of persons of working age (sample-selection bias, see for example Heckman, 1979 and 1980, and Hoffmann and Kassouf, 2005). This may bias the marginal effect of education on earnings as measured by the *Mincerian* wage regression especially if the probability of employment depends on educational attainment. A two-stage selection model (determining the probability of employment at the first stage and the wage for those employed at the second stage) would be a possibility to avoid this problem but is not followed here because: *i)* it would run counter the focus on "standard" wage earners underlying the sample selection strategy followed here; and *ii)* the empirical extent of the problem is small overall even though selection effects turn out to be stronger for women than for men (see Annex 2 in Strauss and de la Maisonneuve, 2007).

Construction of empirical variables[8]

The data for the estimation of education wage *premia* for 21 OECD countries are taken from six different panel databases: the European Community Household Panel (ECHP), the Consortium of Household Panels for European Socio-Economic Research (CHER), the British Household Panel Survey (BHPS), the US Current Population Survey (CPS), the Cross-National Equivalent File (CNEF), and the Household, Income and Labour Dynamics in Australia Survey (HILDA). Household panel data sources are preferred over labour force surveys, which lack detailed wage data for some countries. The ECHP and the CHER were constructed on a cross-country basis, thereby ensuring consistency of definitions and comparability of values of the variables. Essential information on data sources and sample sizes is given in the Annex A1.

The dependent variable is the log of gross hourly wages.[9] The available information is on current salaries for most countries. It is on labour earnings during the year preceding the interview for Canada, Hungary, Poland, Switzerland, the United Kingdom and the United States. Salaries are first brought to an hourly basis using the number of hours worked in the main job and, where needed, the number of months worked in the previous year. Post-estimation corrections are required for countries that report only net wages (Hungary, Luxembourg, Sweden and Switzerland). The advantage of current monthly salaries is that they are consistent with all other variables, which refer to the time of the household interview. The drawback, however, is the exclusion of the self-employed for lack of current-income data.

Individual income data are originally reported in national currency units. They are converted into purchasing power parity dollars (US$ PPP) from the *OECD Economic Outlook* database in order to make the mean and standard deviation comparable across countries.

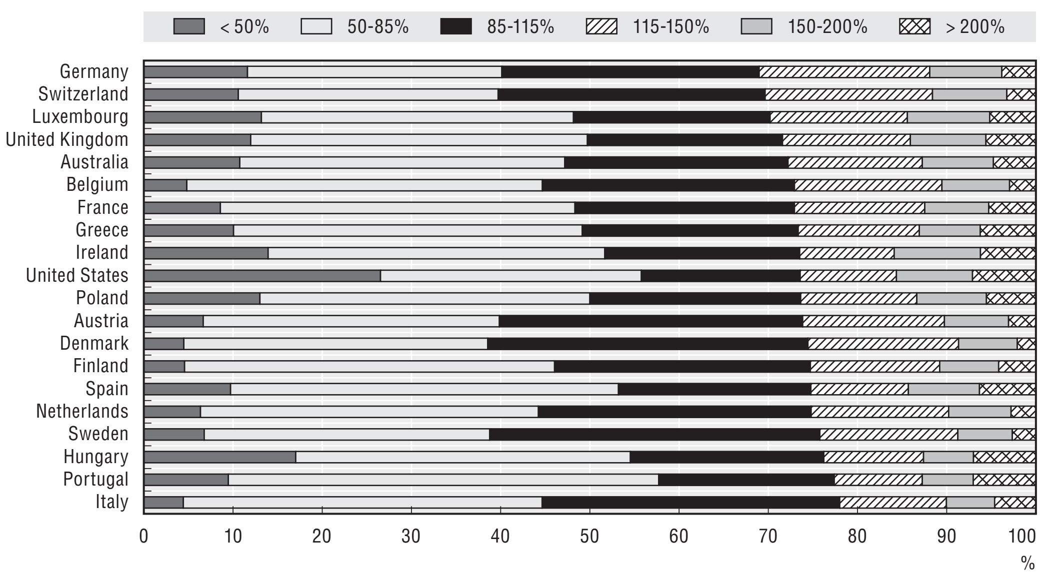
The following additional restrictions are made. First, only employees for whom labour earnings are the main income source are considered. Second, persons working less than 15 hours per week are ruled out[10] as are persons under 16 and over 64. Finally, extreme values are eliminated from the sample (hourly wage below 1 $PPP or above 200 $PPP).

The country- and gender-specific distribution of gross hourly wages is illustrated in Figures 1a and 1b using five income brackets around the country- and gender-specific averages. Regarding the wage distribution for men, some facts are worth pointing out (see Figure 1a):

- The largest share of men (30% or more of the sample) earning more than 115% of the average male wage is found in Germany and Switzerland whereas this share is below 25% in Italy, Portugal, Hungary and Sweden.
- The highest concentration of men in the central bracket (85-115%) is observed for Sweden and Denmark (more than 35% of the sample), the lowest in Portugal (under 20%).
- The share of men in the sample earning less than 85% of the average hourly wage is less than 40% in Austria, Denmark, Sweden and Switzerland but more than 50% in Portugal, Hungary, Spain, Ireland and the United States; the latter are also the countries with the highest share of individuals with an hourly wage in excess of 200% of the average (over 6%), suggesting strong wage dispersion.
- The share of men with hourly wages below half the average is highest in the United States at 26% but below 5% in Italy, Denmark, Finland and Belgium.

Figure 1a. **Wage equation sample distribution 2001:**[1] **Gross**[2] **hourly wage rate of men**

Relative to country average for men – Countries sorted by decreasing frequency of persons earning above 115% of average hourly wage of men

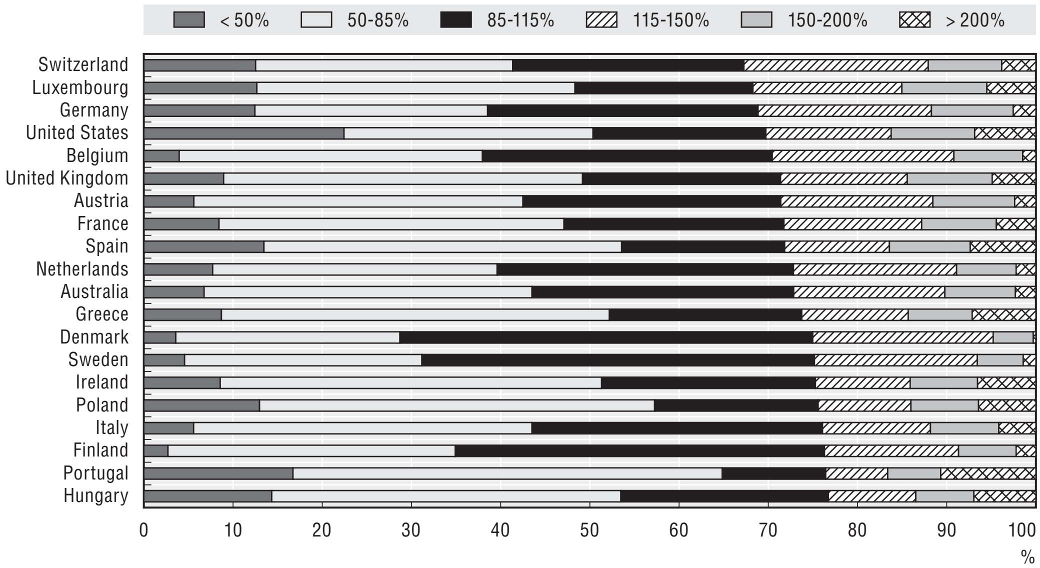


1. Except Hungary (1997); and Poland and Swizerland (2000).
2. Net wage for Hungary, Luxembourg, Sweden, and Switzerland.

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA.

Figure 1b. **Wage equation sample distribution 2001:**[1] **Gross**[2] **hourly wage rate of women**

Relative to country average for women – Countries sorted by decreasing frequency of women earning above 115% of gender-average hourly wage



1. Except Hungary (1997); and Poland and Swizerland (2000).
2. Net wage for Hungary, Luxembourg, Sweden, and Switzerland.

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA.

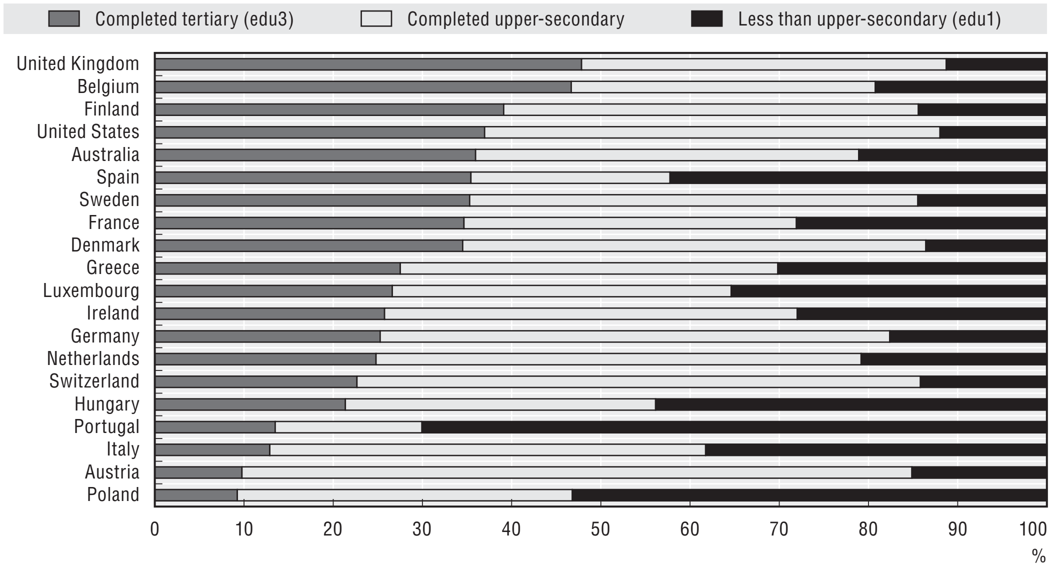
For women, the pattern of gross hourly wages broadly matches that observed for men, with cross-country differences somewhat more pronounced in the central and lower-wage brackets (see Figure 1b).

The most important independent variable is educational attainment. The literature distinguishes the time spent in education from the attainment level, with some positive functional relationship existing between the two (de la Fuente and Jimeno, 2005). However, only the level of educational attainment is consistently available for all 21 countries. The degree of detail varies across databases but for the majority of countries a distinction between only three levels is available: *i)* less than upper secondary education; *ii)* completed upper-secondary education/high school; and *iii)* completed higher/tertiary education. Albeit somewhat rough, this definition of the empirical attainment variable has the advantage of being internationally comparable because it follows the International Standard Classification of Educational Statistics (ISCED, see OECD, 2004).[11, 12]

Most countries do not report the number of years it took individuals to reach their attainment levels. By contrast, the number of years of schooling is available for Australia, Canada, the United Kingdom and the United States. Where necessary (Australia and United Kingdom), a system of correspondence with the above three-tier classification is established using the number of years of education of each individual in the datasets combined with country-wide institutional information on the education system from OECD (2004).

Educational attainment varies widely across countries (Figure 2). Around 45% of the wage-earners in the 2001 samples hold a tertiary degree in Belgium and the United Kingdom. The share is lower but still above one-third in Finland, the United States, Australia, Spain, Sweden, France and Denmark. By contrast, tertiary attainment shares among wage earners

Figure 2. **Wage equation sample distribution 2001:**[1] **Educational attainment**



1. Except Hungary (1997); and Poland and Swizerland (2000).

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA.

cluster around 10% in Portugal, Austria, Italy, and Poland. Tertiary degree holders are overrepresented among wage earners. At the same time, persons with less than upper-secondary degree are underrepresented because participation is more likely the higher the level of educational attainment (Boarini and Strauss, 2009). Over and above differences in attainment structures, countries also vary in their ability to integrate low-skilled persons into the labour market. As a consequence of both influences, the share of wage earners with low attainment is even more dispersed across countries than that of tertiary attainment, ranging from just over 10% in the United Kingdom to 70% in Portugal. The share of workers with completed upper-secondary education is highest in Austria, Switzerland and Germany, where extensive vocational training exists.

To allow for maximum flexibility in the estimation, two dummy variables for educational attainment are created. The first, *edu*1, takes a value of 1 if the individual has not completed upper-secondary education and 0 otherwise. The second, *edu*3, equals 1 for individuals with a degree from tertiary education and zero otherwise. The reference group consists of persons with completed upper-secondary education, for whom both education dummies equal zero.

Another important right-hand-side variable is labour market experience. Human-capital theory discriminates between the productivity effects of formal schooling and those of skills acquired through cumulative work experience. Many empirical studies on returns to education use age (often in a quadratic specification) as a proxy for accumulated labour market experience. Yet age is an imprecise proxy for labour market experience, especially for younger cohorts. This is why a measure of labour market experience (*exper*) is used. *exper* is defined as the difference between the current age and the age at labour market entry. Thus it measures potential rather than actual labour market experience.[13] There is huge cross-country variation in the distribution of labour market experience owing to demographic differences and variation in the effective retirement age. The *Mincerian* equation [1] uses a linear rather than quadratic specification of *exper*. Furthermore, the final specification does not contain an interaction term between educational attainment and experience because this interaction is not supported by the data.

The control variables are as follows. A gender dummy (*woman*) controls for different wage levels between men and women. The gender dummy is interacted with the educational attainment and labour market experience variables to produce gender-specific estimates. For the sake of cross-country comparability of specifications, gender is not interacted with other control variables even though there may be statistically different coefficients for men and women.[14] Marital status also enters the analysis as a dummy variable (*married*) taking the value 1 if the person is formally married and 0 otherwise. The data allow for alternative definitions of living with a partner but this hardly affects the results. Furthermore, job tenure (*tenure*) is calculated as the difference between the year of the interview and that when the person started working with their current employer, plus one. The starting year comes as a discreet variable that is censored in the panel surveys for the majority of countries but with varying "cut-off" years. To ensure cross-country comparability in the definition of this variable *tenure* is capped at 10. Over and above its effect via hours worked, part-time work is captured through a separate dummy variable (*part_time*) since the assumption that each working hour makes the same contribution to weekly earnings (constancy of the hourly wage) may not hold across workers with different time status (part-timer *versus* full-timer). *Part_time* equals one for persons working 15-29 hours per week and zero otherwise. In addition, a variable is created for the type of the employment contract (*indef_cont*), which equals one for persons holding an indefinite-term contract and zero otherwise. Another variable (*public*) is set equal to one for persons working in the public sector and to zero otherwise.

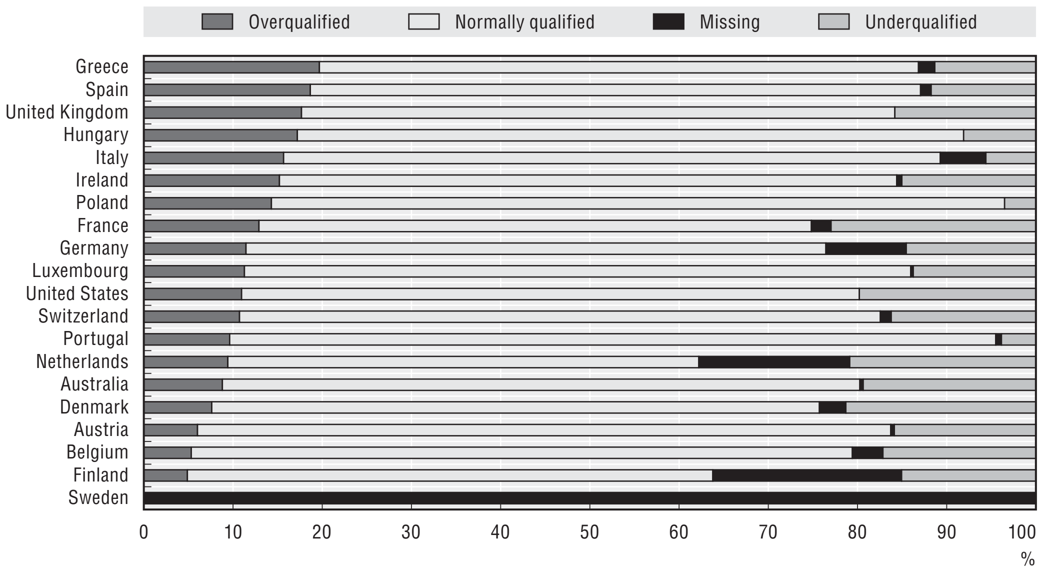
The construction of the remaining two control variables is less straightforward. The first is plant size. Large firms usually pay higher wages than small firms, reflecting either the sharing of profits from market power or higher labour productivity, or both. The ideal information to control for this feature would be firm size, which is not available. A reasonable proxy is the number of persons usually working at the respondent's local production unit (*plant_size*). The variable is provided in a discreet, multinomial form, with plant-size class boundaries varying from country to country. To ensure cross-country comparability the variable is made continuous by assigning each person a random plant size within the limits indicated by the discreet variable, assuming uniform distribution.[15] What enters the equation is the log of that random plant size. Given the way the variable is constructed, cross-country differences in plant-size effects should be interpreted with caution.

Finally, the regression controls for over-qualification and under-qualification in the current job. In fact, hourly wages reflect occupational status as much as educational attainment. Since the two are correlated, the occupational status is not used in the regressions. Rather, the occupational status is accounted for in an indirect way. To illustrate, university graduates working in jobs normally held by high-school degree holders are considered as being over-qualified. Conversely, individuals working in occupations at levels higher than what they can usually access with their attainment levels are considered to be under-qualified.

To assess whether someone has excessive, adequate, or deficient formal education for a job, the occupational levels of all individuals are confronted with their educational attainment levels.[16] The attainment distribution is calculated for each occupation in order to determine the most frequent attainment level for the occupation (the mode), which is defined as the appropriate level. Whenever in a given occupation another attainment level is observed for a share of workers that is within 10 percentage points from the mode, this second-most-frequent attainment level is deemed adequate, too.[17] Individuals with an attainment level above (below) the appropriate level(s) are deemed over-qualified (under-qualified) for their current job. Accordingly, *overqualif* equals one for persons overqualified in their current job (zero otherwise) while *underqualif* equals one for the under-qualified (zero otherwise). The occupation-attainment matrix is calculated separately for each country to account for the diversity of national education and training systems.[18]

Figure 3 depicts the over- and under-qualification pattern for the countries in the sample. The incidence of over-qualification is 12% on average across countries, ranging from 5% in Finland and Belgium to 20% in Greece. The share of persons considered as under-qualified lies in a similar range. It is lowest in Portugal and Poland and highest in Denmark, France and the Netherlands.

Figure 3. **Wage equation sample distribution 2001:**[1]
Over- or under-qualified for current occupation



1. Except Hungary (1997); and Poland and Swizerland (2000).

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA.

Results

For each country and available year, a cross-sectional OLS regression of the *Mincerian* wage equation described in [1] is run. Standard errors are robust, *i.e.* corrected for outliers. Table 1 shows the results for 2001, the year with the maximum number of available countries. Numbers in bold pertain to the wage effect of tertiary education, with those in the upper line representing the wage premium for men and the sum of coefficients in the upper and the lower bolded lines that for women. The same kind of addition of the interacted and the non-interacted coefficients is required to obtain the average female wage "penalty" of not having completed upper-secondary education (*edu*1) and women's average annual wage increment due to labour market experience (*exper*).

In the 21 regressions for the year 2001 (1997 for Hungary; 2000 for Poland and Switzerland) over 95% of the coefficients are significant, virtually all of them at the 1%-level. Moreover, the sign of all non-interacted coefficients is in line with expectations.

Table 1. **Results of the *Mincerian* wage regressions for 21 OECD countries, 2001**[1]

		Australia	Austria	Belgium	Canada	Denmark	Finland	France	Germany	Greece	Hungary[2]	Ireland
edu1	**Estimate**	–0.181***	–0.517***	–0.229***	–0.299***	–0.258***	–0.242***	–0.128***	–0.232***	–0.236***	–0.259***	–0.272***
	Robust std. error	[0.020]	[0.098]	[0.027]	[0.019]	[0.030]	[0.039]	[0.021]	[0.030]	[0.024]	[0.061]	[0.035]
edu1w	**Estimate**	–0.032	0.006	–0.019	–0.083***	–0.038	–0.009	–0.066*	0.045	0.031	–0.058	–0.003
	Robust std. error	[0.028]	[0.040]	[0.047]	[0.028]	[0.043]	[0.041]	[0.034]	[0.039]	[0.035]	[0.071]	[0.045]
edu3	**Estimate**	**0.351*** **	**0.433*** **	**0.334*** **	**0.402*** **	**0.387*** **	**0.424*** **	**0.462*** **	**0.383*** **	**0.303*** **	**0.477*** **	**0.434*** **
	Robust std. error	[0.019]	[0.044]	[0.024]	[0.014]	[0.021]	[0.025]	[0.024]	[0.022]	[0.025]	[0.074]	[0.038]
edu3w	**Estimate**	**–0.057** **	**–0.144*** **	**–0.024**	**0.038** **	**–0.033**	**–0.065** **	**–0.010**	**0.023**	**0.083** **	**–0.011**	**0.086* **
	Robust std. error	[0.023]	[0.054]	[0.032]	[0.018]	[0.028]	[0.031]	[0.036]	[0.030]	[0.035]	[0.091]	[0.050]
exper	**Estimate**	0.007***	0.007***	0.010***	0.007***	0.004***	0.006***	0.007***	0.002***	0.008***	0.006**	0.006***
	Robust std. error	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.002]	[0.001]
experw	**Estimate**	–0.002*	0.000	–0.003*	–0.001*	–0.001	–0.003**	–0.005***	–0.002	–0.002	–0.003	–0.003**
	Robust std. error	[0.001]	[0.001]	[0.002]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.003]	[0.002]
woman	**Estimate**	–0.054**	–0.160***	–0.056	–0.247***	–0.080**	–0.121***	–0.073**	–0.137***	–0.167***	–0.101	–0.136***
	Robust std. error	[0.023]	[0.034]	[0.040]	[0.020]	[0.035]	[0.037]	[0.035]	[0.034]	[0.029]	[0.071]	[0.037]
married	**Estimate**	0.103***	0.037**	0.008	0.157***	0.031**	0.016	0.051***	0.059***	0.106***	0.024	0.120***
	Robust std. error	[0.012]	[0.017]	[0.016]	[0.009]	[0.014]	[0.015]	[0.015]	[0.014]	[0.016]	[0.031]	[0.026]
public	**Estimate**	–0.012	–0.002	–0.023	0.204***	–0.094***	–0.038**	0.058***	0.013	0.109***	–0.052	0.174***
	Robust std. error	[0.013]	[0.017]	[0.015]	[0.010]	[0.015]	[0.015]	[0.016]	[0.015]	[0.018]	[0.032]	[0.023]
part_time	**Estimate**	–0.011	0.060**	0.071***	–0.068***	–0.005	0.100**	0.132***	0.126***	0.293***	0.183**	0.054**
	Robust std. error	[0.015]	[0.030]	[0.026]	[0.012]	[0.033]	[0.047]	[0.031]	[0.025]	[0.036]	[0.080]	[0.027]
tenure	**Estimate**	0.010***	0.015***	0.018***	. .	0.010***	0.010***	0.031***	0.019***	0.022***	0.020***	0.018***
	Robust std. error	[0.002]	[0.003]	[0.003]	. .	[0.002]	[0.003]	[0.003]	[0.002]	[0.003]	[0.005]	[0.004]
indef_cont	**Estimate**	0.046***	0.316***	0.112***	. .	0.184***	0.057**	0.252***	0.296***	0.131***	. .	0.179***
	Robust std. error	[0.013]	[0.036]	[0.030]	. .	[0.030]	[0.027]	[0.028]	[0.024]	[0.020]	. .	[0.033]
plant_size	**Estimate**	0.042***	0.044***	0.049***	. .	0.027***	0.045***	. .	0.062***	0.041***	. .	0.041***
	Robust std. error	[0.003]	[0.004]	[0.004]	. .	[0.004]	[0.004]	. .	[0.003]	[0.004]	. .	[0.006]
overqualif	**Estimate**	–0.233***	–0.003	–0.160***	–0.377***	–0.207***	–0.305***	–0.225***	–0.281***	–0.163***	–0.021	–0.266***
	Robust std. error	[0.020]	[0.050]	[0.043]	[0.023]	[0.027]	[0.036]	[0.020]	[0.025]	[0.021]	[0.053]	[0.034]
underqualif	**Estimate**	0.144***	0.220**	0.131***	0.240***	0.184***	0.236***	0.197***	0.127***	0.086***	0.218***	0.169***
	Robust std. error	[0.017]	[0.094]	[0.023]	[0.013]	[0.019]	[0.032]	[0.018]	[0.025]	[0.023]	[0.063]	[0.028]
Constant	**Estimate**	2.111***	1.713***	1.913***	2.105***	2.320***	2.023***	1.719***	1.670***	1.413***	0.575***	1.869***
	Robust std. error	[0.022]	[0.041]	[0.043]	[0.015]	[0.044]	[0.035]	[0.036]	[0.037]	[0.028]	[0.063]	[0.043]
Observations		5 211	1 965	1 615	25 555	1 585	1 866	2 891	3 688	2 079	879	1 457
Adjusted R-squared		0.31	0.46	0.46	0.17	0.44	0.4	0.42	0.36	0.59	0.33	0.51

Table 1. **Results of the *Mincerian* wage regressions for 21 OECD countries, 2001**[1] *(cont.)*

		Italy	Luxembourg[2]	Netherlands	Poland	Portugal	Spain	Sweden[2]	Switzerland[2]	United Kingdom	United States
edu1	**Estimate**	–0.265***	–0.422***	–0.262***	–0.293***	–0.444***	–0.277***	–0.178***	–0.616***	–0.351***	–0.650***
	Robust std. error	[0.017]	[0.022]	[0.051]	[0.033]	[0.031]	[0.026]	[0.028]	[0.045]	[0.025]	[0.014]
edu1w	**Estimate**	0.003	0.015	0.103**	–0.009	–0.01	0.024	0.054	0.236***	0.019	–0.004
	Robust std. error	[0.024]	[0.042]	[0.045]	[0.040]	[0.034]	[0.032]	[0.039]	[0.062]	[0.031]	[0.018]
edu3	**Estimate**	**0.411*** **	**0.424*** **	**0.348*** **	**0.306*** **	**0.505*** **	**0.234*** **	**0.260*** **	**0.378*** **	**0.502*** **	**0.650*** **
	Robust std. error	[0.027]	[0.025]	[0.029]	[0.071]	[0.043]	[0.024]	[0.023]	[0.033]	[0.016]	[0.010]
edu3w	**Estimate**	**–0.083** **	**–0.023**	**0.030**	**0.307*** **	**0.148*** **	**0.079** **	**–0.046**	**–0.048**	**0.038* **	**–0.011**
	Robust std. error	[0.037]	[0.041]	[0.041]	[0.085]	[0.049]	[0.033]	[0.030]	[0.043]	[0.019]	[0.012]
exper	**Estimate**	0.007***	0.014***	0.006***	0.005***	0.003***	0.006***	0.010***	0.017***	0.007***	0.015***
	Robust std. error	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.001]	[0.000]
experw	**Estimate**	–0.001	–0.006***	–0.003*	0.002	0.003***	0.001	–0.002*	–0.003*	–0.004***	–0.006***
	Robust std. error	[0.001]	[0.002]	[0.002]	[0.002]	[0.001]	[0.001]	[0.001]	[0.002]	[0.001]	[0.000]
woman	**Estimate**	–0.114***	–0.083**	–0.131***	–0.309***	–0.279***	–0.279***	–0.050	–0.143***	–0.122***	–0.186***
	Robust std. error	[0.024]	[0.038]	[0.046]	[0.047]	[0.031]	[0.031]	[0.036]	[0.042]	[0.022]	[0.013]
married	**Estimate**	0.054***	0.089***	0.055***	0.206***	0.066***	0.043***	0.030**	0.025	0.093***	0.166***
	Robust std. error	[0.012]	[0.016]	[0.019]	[0.023]	[0.013]	[0.014]	[0.014]	[0.019]	[0.010]	[0.006]
public	**Estimate**	0.067***	0.252***	–0.008	0.145***	0.204***	0.077***	–0.130***	0.049**	0.063***	–0.083***
	Robust std. error	[0.012]	[0.017]	[0.019]	[0.020]	[0.015]	[0.017]	[0.015]	[0.019]	[0.011]	[0.008]
part_time	**Estimate**	0.280***	0.112***	0.089***	0.183***	0.214***	0.086***	0.105***	0.040	–0.082***	–0.212***
	Robust std. error	[0.020]	[0.036]	[0.027]	[0.043]	[0.043]	[0.028]	[0.028]	[0.032]	[0.014]	[0.012]
tenure	**Estimate**	0.012***	. .	0.019***	. .	0.018***	0.028***	. .	0.003**	0.007***	. .
	Robust std. error	[0.002]	. .	[0.003]	. .	[0.002]	[0.002]	. .	[0.001]	[0.001]	. .
indef_cont	**Estimate**	0.133***	0.276***	0.174***	. .	0.066***	0.062***	0.280***	0.611***	0.170***	. .
	Robust std. error	[0.022]	[0.034]	[0.047]	. .	[0.016]	[0.016]	[0.040]	[0.054]	[0.025]	. .
plant_size	**Estimate**	0.031***	0.041***	0.022***	. .	0.044***	0.054***	0.025***	0.034***	0.041***	0.037***
	Robust std. error	[0.003]	[0.004]	[0.004]	. .	[0.004]	[0.003]	[0.003]	[0.004]	[0.002]	[0.001]
overqualif	**Estimate**	–0.185***	–0.232***	–0.041	–0.234***	–0.300***	–0.234***	. .	–0.132***	–0.404***	–0.455***
	Robust std. error	[0.017]	[0.023]	[0.033]	[0.031]	[0.025]	[0.019]	. .	[0.038]	[0.014]	[0.010]
underqualif	**Estimate**	0.123***	0.264***	0.128***	0.155***	0.357***	0.122***	. .	. .	0.215***	0.306***
	Robust std. error	[0.022]	[0.025]	[0.046]	[0.040]	[0.036]	[0.023]	. .	. .	[0.017]	[0.010]
Constant	**Estimate**	1.927***	1.811***	2.187***	0.955***	1.571***	1.792***	1.470***	1.721***	1.926***	1.888***
	Robust std. error	[0.028]	[0.043]	[0.055]	[0.038]	[0.033]	[0.029]	[0.045]	[0.059]	[0.030]	[0.012]
Observations		3 254	2 213	2 031	2 285	3 859	3 615	2 330	2 260	7 960	49 571
Adjusted R-squared		0.49	0.58	0.31	0.32	0.61	0.5	0.26	0.53	0.38	0.32

Robust standard errors in brackets.
* significant at 10%; ** significant at 5%; *** significant at 1%.
1. Except Hungary (1997); and Poland and Switzerland (2000).
2. Estimation based on *net* hourly wages and, hence, results are not directly comparable with those for other countries.

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA and authors' calculations.

The schooling coefficient of the log-wage equation is usually taken as an approximation of the wage premium associated with tertiary education, exploiting the fact that $ln(1 + x) \approx x$ for small x. This approximation is unproblematic when the right-hand-side variable is the log of a continuous variable, allowing for an elasticity interpretation, nor is it an issue for a discreet right-hand-side variable that can be changed in small increments such as years of schooling since this usually implies small coefficients in the regression. In the case at hand, however, the educational attainment variable of interest (*edu*3) is a binary variable and the change from 0 to 1 represents a major step. Correspondingly, the estimated effect on log wages is substantial – between 0.23 and 0.65 – making the logarithmic approximation unsatisfactory. This is why the estimated coefficients of *edu*3, α_2 for men and ($\alpha_2 + \alpha_4$) for women, are transformed into precise tertiary wage *premia* using the following formulae:

- Male wage premium = $[\exp(\alpha_2)–1].100\%$; and
- Female wage premium= $[\exp(\alpha_2 + \alpha_4)–1].100\%$.

Applying this interpretation to the coefficients of 2001, the tertiary education wage premium for men is highest in the United States at 92% and lowest in Spain at 27%. For women, wage *premia* are highest in Portugal at 92% and lowest in Sweden at 24%[19] (Table 2). Women's tertiary wage *premia* are higher than men's (positive interaction coefficient) in nine of the 21 countries but differences appear to be significant only for Greece, Poland, Portugal and Spain (Figure 4). By contrast, male graduates appear to yield significantly higher wage returns than their female counterparts in Austria, Finland and Italy.

The comparison of the 2001 estimates with those for other years suggests that by and large tertiary wage *premia* are fairly stable over time (see Table 2).[20] If anything, a slight upward tendency is observed from 1994 to 2000 against the backdrop of gradually accelerating economic growth, which might have put more pressure on the demand for high-skilled than for low-skilled workers. Figure 5 gives a graphical illustration of the evolution in seven countries. Wage *premia* on tertiary education are seen to increase in Denmark and Ireland and to a lesser extent also in Germany and in the United States. Notable gender-specific trends with respect to female wage *premia* include steep increases in Greece, Poland and Portugal and a decline in Austria.

In 2001, not having completed upper-secondary education affected log wages very differently, ranging from –0.13 in France to –0.65 in the United States. This implies upper-secondary wage *premia* between 14% and 92% of the average wage of persons without complete upper-secondary education (Table 3). *Premia* appear to be substantially higher for women than for men in Canada and France but lower in the Netherlands and in Switzerland. Over and above the substantial cross-country variation, upper-secondary wage *premia* are subject to stronger fluctuations over time. The coefficient of variation averages 0.20 for the 42 country-gender pairs, which is 1½ times that for tertiary education. Thus persons without completed secondary education appear to be more exposed to business cycle shocks than lower-educated persons.[21]

Table 2. **Gross wage *premia* on tertiary education for men and women in 21 OECD countries, 1991-2005**

Average percentage changes from wage of upper-secondary degree holders[1, 2]

		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	Multi-period average	Cross-period coefficient of variation
Australia	Men	. .	. .	. .	. .	. .	. .	. .	. .	. .	. .	42.1	40.6	41.3	. .	. .	**41.4**	0.02
	Women	. .	. .	. .	. .	. .	. .	. .	. .	. .	. .	34.2	40.0	39.2	. .	. .	**37.8**	0.08
	Difference m-w	. .	. .	. .	. .	. .	. .	. .	. .	. .	. .	*7.9*	*0.7*	*2.2*	. .	. .	*3.6*	
Austria	Men	. .	. .	. .	. .	63.4	61.7	51.2	54.8	58.9	61.8	53.8	. .	. .	. .	. .	**58.0**	0.08
	Women	. .	. .	. .	. .	74.4	76.0	54.7	48.9	49.8	39.1	33.3	. .	. .	. .	. .	**53.7**	0.30
	Difference m-w	. .	. .	. .	. .	*–11.0*	*–14.3*	*–3.5*	*5.9*	*9.2*	*22.7*	*20.5*	. .	. .	. .	. .	*4.2*	
Belgium	Men	. .	. .	. .	36.9	34.9	35.9	36.7	43.2	36.1	36.1	40.2	. .	. .	. .	. .	**37.5**	0.07
	Women	. .	. .	. .	24.9	28.8	24.4	30.0	31.4	45.0	34.2	36.3	. .	. .	. .	. .	**31.9**	0.21
	Difference m-w	. .	. .	. .	*12.0*	*6.1*	*11.6*	*6.7*	*11.8*	*–8.8*	*1.9*	*3.8*	. .	. .	. .	. .	*5.6*	
Canada	Men	. .	. .	27.7	26.3	29.3	40.2	31.3	38.9	38.7	51.9	49.5	47.1	. .	. .	. .	**38.1**	0.24
	Women	. .	. .	39.3	37.6	34.4	41.4	36.8	43.3	42.8	58.7	55.4	55.2	. .	. .	. .	**44.5**	0.20
	Difference m-w	. .	. .	*–11.6*	*–11.3*	*–5.1*	*–1.2*	*–5.4*	*–4.4*	*–4.1*	*–6.8*	*–5.8*	*–8.1*	. .	. .	. .	*–6.4*	
Denmark	Men	. .	. .	. .	24.1	31.6	30.2	31.2	38.5	39.8	40.5	47.6	. .	. .	. .	. .	**35.4**	0.21
	Women	. .	. .	. .	25.2	29.2	30.8	25.3	38.4	35.9	37.2	42.6	. .	. .	. .	. .	**33.1**	0.19
	Difference m-w	. .	. .	. .	*–1.1*	*2.4*	*–0.6*	*5.9*	*0.1*	*3.8*	*3.3*	*5.0*	. .	. .	. .	. .	*2.4*	
Finland	Men	. .	. .	. .	. .	. .	53.6	56.4	47.4	53.5	54.5	52.6	. .	. .	. .	. .	**53.0**	0.06
	Women	. .	. .	. .	. .	. .	39.8	38.2	36.0	41.3	38.9	43.1	. .	. .	. .	. .	**39.5**	0.06
	Difference m-w	. .	. .	. .	. .	. .	*13.8*	*18.2*	*11.5*	*12.2*	*15.7*	*9.6*	. .	. .	. .	. .	*13.5*	
France	Men	. .	. .	. .	67.9	65.5	66.3	70.3	72.6	64.8	71.1	58.8	. .	. .	. .	. .	**67.2**	0.07
	Women	. .	. .	. .	56.7	55.9	56.4	63.8	63.7	57.7	57.6	57.2	. .	. .	. .	. .	**58.6**	0.05
	Difference m-w	. .	. .	. .	*11.2*	*9.6*	*9.9*	*6.5*	*8.9*	*7.0*	*13.6*	*1.6*	. .	. .	. .	. .	*8.5*	
Germany	Men	. .	. .	. .	41.9	42.4	39.9	40.7	46.3	53.5	50.5	46.3	. .	. .	. .	. .	**45.2**	0.11
	Women	. .	. .	. .	35.2	42.8	37.4	41.1	40.8	47.1	42.2	49.6	. .	. .	. .	. .	**42.0**	0.11
	Difference m-w	. .	. .	. .	*6.8*	*–0.4*	*2.5*	*–0.3*	*5.5*	*6.4*	*8.3*	*–3.3*	. .	. .	. .	. .	*3.2*	
Greece	Men	. .	. .	. .	30.6	30.1	32.6	30.2	29.2	35.4	35.7	35.3	. .	. .	. .	. .	**32.4**	0.08
	Women	. .	. .	. .	26.6	23.4	27.5	33.4	50.7	45.2	45.5	47.2	. .	. .	. .	. .	**37.4**	0.29
	Difference m-w	. .	. .	. .	*3.9*	*6.7*	*5.1*	*–3.1*	*–21.4*	*–9.8*	*–9.8*	*–11.9*	. .	. .	. .	. .	*–5.0*	
Hungary[3]	Men	. .	66.2	58.6	61.2	41.1	55.5	61.1	. .	. .	. .	. .	. .	. .	. .	. .	**57.3**	0.15
	Women	. .	44.1	58.9	66.4	58.4	49.9	59.3	. .	. .	. .	. .	. .	. .	. .	. .	**56.2**	0.14
	Difference m-w	. .	*22.1*	*–0.4*	*–5.2*	*–17.3*	*5.7*	*1.8*	. .	. .	. .	. .	. .	. .	. .	. .	*1.1*	
Ireland	Men	. .	. .	. .	31.8	37.3	40.3	60.5	53.3	51.7	52.8	54.3	. .	. .	. .	. .	**47.8**	0.21
	Women	. .	. .	. .	40.5	47.8	51.1	76.0	58.7	71.7	59.6	68.4	. .	. .	. .	. .	**59.2**	0.21
	Difference m-w	. .	. .	. .	*–8.7*	*–10.5*	*–10.9*	*–15.5*	*–5.3*	*–20.0*	*–6.8*	*–14.1*	. .	. .	. .	. .	*–11.5*	
Italy	Men	. .	. .	. .	42.3	39.8	43.1	44.6	51.8	50.4	49.6	50.9	. .	. .	. .	. .	**46.6**	0.10
	Women	. .	. .	. .	39.2	37.5	35.5	35.1	38.3	40.4	43.1	38.8	. .	. .	. .	. .	**38.5**	0.07
	Difference m-w	. .	. .	. .	*3.1*	*2.3*	*7.6*	*9.5*	*13.5*	*10.0*	*6.4*	*12.1*	. .	. .	. .	. .	*8.1*	

Table 2. **Gross wage *premia* on tertiary education for men and women in 21 OECD countries, 1991-2005** *(cont.)*

Average percentage changes from wage of upper-secondary degree holders[1, 2]

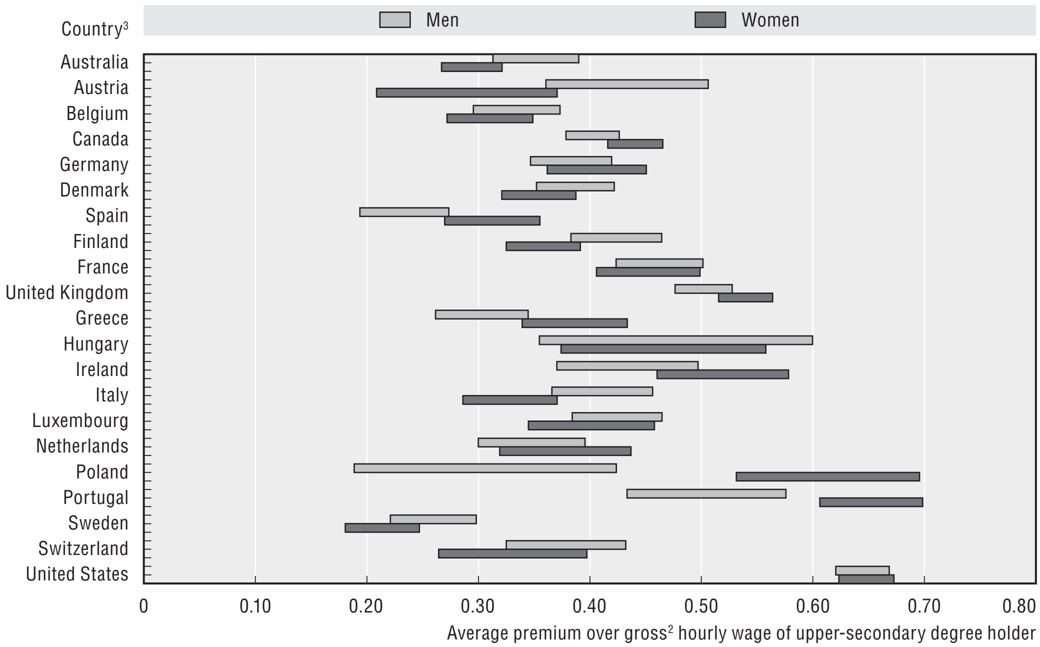
		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	Multi-period average	Cross-period coefficient of variation
Luxembourg[3]	Men	. .	. .	. .	. .	62.7	56.2	51.4	64.5	62.5	67.4	52.6	. .	. .	. .	. .	**59.6**	0.10
	Women	. .	. .	. .	. .	67.6	51.7	50.4	59.0	53.8	56.4	49.3	. .	. .	. .	. .	**55.5**	0.11
	Difference m-w	. .	. .	. .	. .	*–4.9*	*4.5*	*0.9*	*5.5*	*8.6*	*11.1*	*3.3*	. .	. .	. .	. .	*4.1*	
Netherlands	Men	. .	. .	. .	48.9	49.3	48.6	43.0	40.3	33.4	38.5	41.7	. .	. .	. .	. .	**43.0**	0.13
	Women	. .	. .	. .	36.2	41.0	41.7	32.0	31.1	30.9	30.3	45.9	. .	. .	. .	. .	**36.1**	0.17
	Difference m-w	. .	. .	. .	*12.6*	*8.3*	*6.9*	*11.0*	*9.2*	*2.5*	*8.2*	*–4.2*	. .	. .	. .	. .	*6.8*	
Poland	Men	. .	. .	. .	. .	. .	. .	40.2	48.5	52.9	35.8	. .	. .	. .	. .	. .	**44.4**	0.18
	Women	. .	. .	. .	. .	. .	. .	55.9	65.8	80.1	84.7	. .	. .	. .	. .	. .	**71.6**	0.18
	Difference m-w	. .	. .	. .	. .	. .	. .	*–15.7*	*–17.2*	*–27.1*	*–48.9*	. .	. .	. .	. .	. .	*–27.2*	
Portugal	Men	. .	. .	. .	67.0	87.2	104.6	80.2	87.6	75.8	87.0	65.8	. .	. .	. .	. .	**81.9**	0.15
	Women	. .	. .	. .	68.6	77.1	90.2	77.2	95.7	82.2	113.6	91.8	. .	. .	. .	. .	**87.1**	0.16
	Difference m-w	. .	. .	. .	*–1.6*	*10.1*	*14.4*	*3.1*	*–8.1*	*–6.3*	*–26.6*	*–26.0*	. .	. .	. .	. .	*–5.1*	
Spain	Men	. .	. .	. .	27.8	29.8	29.8	23.4	22.2	18.2	15.6	26.9	. .	. .	. .	. .	**24.2**	0.22
	Women	. .	. .	. .	34.5	36.8	34.6	29.8	25.2	25.2	25.0	36.5	. .	. .	. .	. .	**31.0**	0.17
	Difference m-w	. .	. .	. .	*–6.7*	*–7.0*	*–4.8*	*–6.4*	*–3.0*	*–7.0*	*–9.4*	*–9.6*	. .	. .	. .	. .	*–6.7*	
Sweden[3]	Men	. .	. .	. .	. .	. .	. .	27.2	32.1	32.9	32.1	29.6	. .	. .	. .	. .	**30.8**	0.08
	Women	. .	. .	. .	. .	. .	. .	25.5	24.4	18.5	22.1	23.7	. .	. .	. .	. .	**22.9**	0.12
	Difference m-w	. .	. .	. .	. .	. .	. .	*1.7*	*7.7*	*14.3*	*9.9*	*5.9*	. .	. .	. .	. .	*7.9*	
Switzerland[3]	Men	. .	. .	. .	. .	. .	. .	. .	. .	50.7	46.0	. .	. .	. .	. .	. .	**48.4**	0.07
	Women	. .	. .	. .	. .	. .	. .	. .	. .	40.9	39.2	. .	. .	. .	. .	. .	**40.1**	0.03
	Difference m-w	. .	. .	. .	. .	. .	. .	. .	. .	*9.8*	*6.8*	. .	. .	. .	. .	. .	*8.3*	
United Kingdom	Men	69.7	68.5	71.5	69.2	62.8	68.3	65.9	66.4	68.9	64.5	65.2	62.9	58.5	52.1	. .	**65.3**	0.08
	Women	72.3	84.7	79.7	74.5	74.5	71.1	74.4	75.1	71.7	65.2	71.5	69.3	61.9	53.9	. .	**71.4**	0.10
	Difference m-w	*–2.6*	*–16.2*	*–8.2*	*–5.2*	*–11.6*	*–2.8*	*–8.5*	*–8.7*	*–2.8*	*–0.7*	*–6.3*	*–6.4*	*–3.3*	*–1.8*	. .	*–6.1*	
United States	Men	. .	. .	. .	80.9	80.1	84.1	80.3	76.9	90.5	89.8	91.6	100.8	92.2	95.4	94.6	**88.1**	0.08
	Women	. .	. .	. .	82.6	84.1	87.6	84.9	82.5	89.0	87.9	89.4	94.2	84.0	90.6	89.9	**87.2**	0.04
	Difference m-w	. .	. .	. .	*–1.8*	*–4.0*	*–3.6*	*–4.6*	*–5.6*	*1.5*	*1.9*	*2.2*	*6.6*	*8.2*	*4.8*	*4.6*	*0.9*	

. .: Means "not available".
1. Effect of changing variable *edu*3 from 0 to 1, leaving other variables unchanged. Given the point estimates for men (α_2) and women ($\alpha_2 + \alpha_4$), *premia* equal $[\exp(\alpha_2)–1]*100\%$ and $[\exp(\alpha_2 + \alpha_4)–1]*100\%$, respectively.
2. t-values of the underlying point estimates not reported. All coefficients for men significant at the 1% level. Wage *premia* for women are significantly different from zero but not all are significantly different from male wage *premia* (see Figure 4).
3. Estimation based on *net* hourly wages and, hence, results not directly comparable with those for other countries.

Source: Authors' calculations.

Figure 4. **Male-female differences in tertiary-education coefficients**

90% confidence intervals of point estimates, 2001[1]



1. Except Hungary (1997); and Poland and Swizerland (2000).
2. Net wage for Hungary, Luxembourg, Poland, and Switzerland.
3. Upper bar: men; lower bar: women.

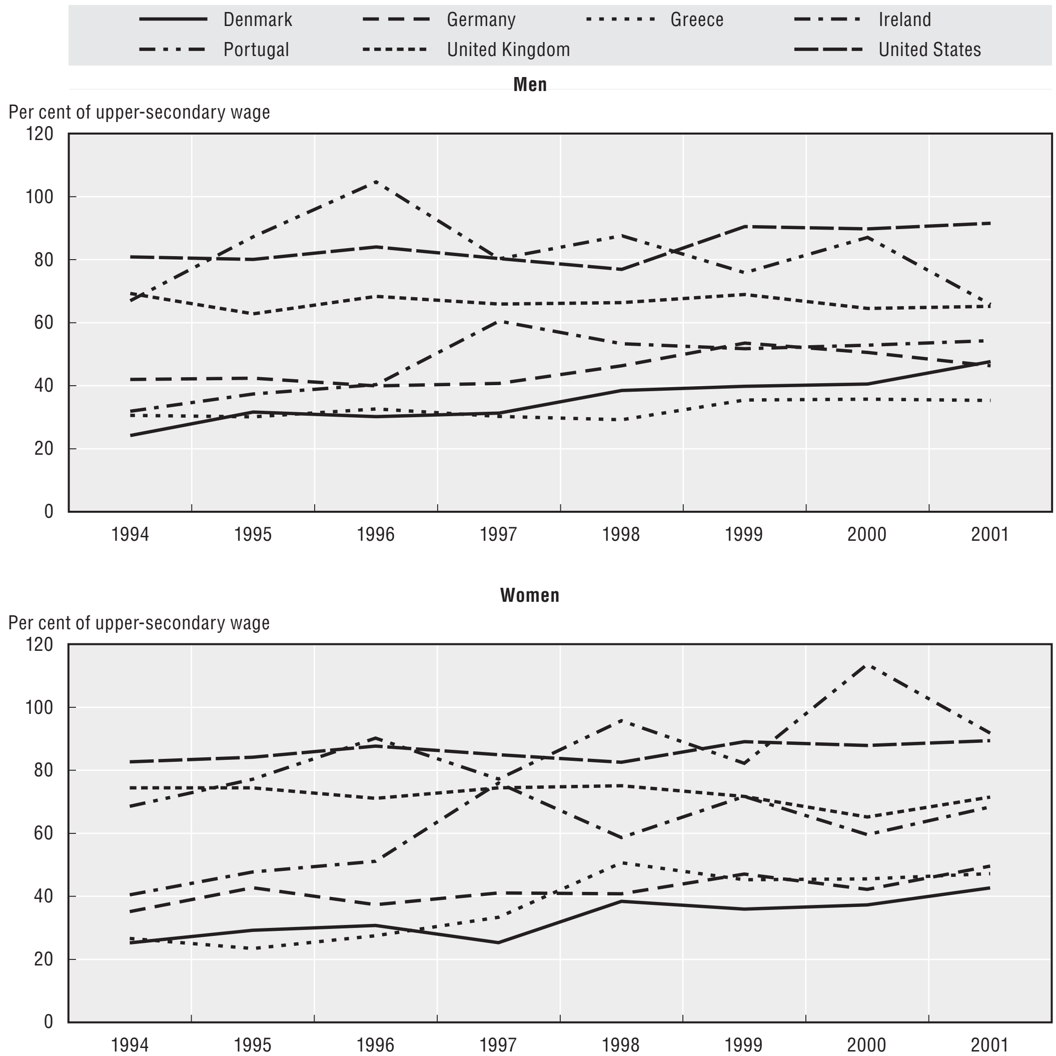
Source: Authors' calculations.

Concerning the variables other than educational attainment, the following points are worth noting with reference to 2001 estimates:[22]

- Skill accumulation is rewarded throughout but the relative importance of general labour-market relative to job-specific skills varies across countries:
  - ❖ The experience premium ranges from 0.23% *per annum* in Germany to 1.69% in Switzerland, appears to be lower for women than for men (except for Poland and Portugal), and turns out to be fairly stable over time;
  - ❖ the effect of an additional year of job tenure (0.3% to 3.1%) tends to be negatively correlated with the experience premium;
- The estimated gender pay gap (all education levels) averages 15% in 2001 and is highest in Poland (36% of the male wage) and lowest in Sweden (5%).
- Married persons tend to earn significantly more than non-married persons in 17 countries (up to 23% in Poland) but not in Belgium, Finland, Hungary and Switzerland.
- Working for the public sector entails positive wage effects in the majority of countries, with the premium exceeding 20% in Canada, Luxembourg and Portugal, but a penalty in the Nordic countries and the United States.[23]
- The wage effect of working part-time is heterogeneous across countries: while positive in the majority of countries (as high as 30% in Greece and Italy), it tends to be negative in the English-speaking countries (nearly –20% in the United States).

Figure 5. **Evolution of gross wage *premia* for selected countries**

1994-2001



Source: Authors' calculations.

- An indefinite-term contract improves a worker's wage by between 6% and 84% on average.
- A person working in a plant twice the size of another person's plant earns between 2% (Netherlands) and 6% (Germany) more.
- Being over-qualified reduces the hourly pay by between 12% (Switzerland) and 37% (United States) of the upper-secondary degree holder's wage whereas being under-qualified raises wages by 9% to 43% compared with persons with the same attainment level working in lower occupations; the point estimators suggest that on average: *i)* working in "too low" an occupation does not fully take away the education wage premium; and *ii)* making it to "too high" an occupation does not fully substitute for education.

Table 3. **Gross wage *premia* on upper-secondary education for men and women in 21 OECD countries, 1991-2005**

Average percentage changes from wage of persons with less than upper-secondary degree[1, 2]

		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	Multi-period average	Cross-period coefficient of variation
Australia	Men	..	..	..	..	..	..	..	..	..	..	19.9	21.1	21.0	..	..	**20.7**	0.03
	Women	..	..	..	..	..	..	..	..	..	..	23.8	17.6	24.2	..	..	**21.9**	0.17
	Difference m-w	..	..	..	..	..	..	..	..	..	..	*–3.9*	*3.5*	*–3.2*	..	..	*–1.2*	
Austria	Men	..	..	..	..	28.1	26.6	32.5	33.3	43.1	31.0	68.2	..	..	..	..	**37.5**	0.39
	Women	..	..	..	..	29.5	33.6	35.6	33.1	40.8	33.0	67.2	..	..	..	..	**39.0**	0.33
	Difference m-w	..	..	..	..	*–1.5*	*–7.0*	*–3.1*	*0.2*	*2.3*	*–2.0*	*1.0*	..	..	..	..	*–1.5*	
Belgium	Men	..	..	..	11.7	12.6	12.5	17.8	19.5	18.7	20.5	25.5	..	..	..	..	**17.4**	0.28
	Women	..	..	..	26.6	28.8	27.3	25.9	27.6	16.5	6.7	28.7	..	..	..	..	**23.5**	0.33
	Difference m-w	..	..	..	*–14.9*	*–16.2*	*–14.8*	*–8.1*	*–8.1*	*2.2*	*13.9*	*–3.2*	..	..	..	..	*–6.1*	
Canada	Men	..	..	33.5	32.5	22.2	25.5	23.8	26.8	29.3	27.9	34.9	24.0	..	..	..	**28.0**	0.16
	Women	..	..	29.0	33.3	29.7	35.5	29.1	40.6	40.3	44.6	46.6	39.5	..	..	..	**36.8**	0.18
	Difference m-w	..	..	*4.5*	*–0.8*	*–7.6*	*–10.0*	*–5.3*	*–13.8*	*–11.0*	*–16.7*	*–11.7*	*–15.6*	..	..	..	*–8.8*	
Denmark	Men	..	..	..	32.7	24.0	23.0	27.8	36.0	36.3	33.7	29.3	..	..	..	..	**30.3**	0.17
	Women	..	..	..	27.8	21.8	21.5	28.9	34.1	42.0	48.2	33.9	..	..	..	..	**32.3**	0.29
	Difference m-w	..	..	..	*5.0*	*2.2*	*1.4*	*–1.1*	*1.9*	*–5.7*	*–14.5*	*–4.6*	..	..	..	..	*–1.9*	
Finland	Men	..	..	..	..	..	12.7	16.8	16.1	28.9	36.2	27.2	..	..	..	..	**23.0**	0.40
	Women	..	..	..	..	..	15.0	16.5	14.4	22.3	37.3	28.5	..	..	..	..	**22.3**	0.41
	Difference m-w	..	..	..	..	..	*–2.3*	*0.3*	*1.7*	*6.6*	*–1.1*	*–1.4*	..	..	..	..	*0.7*	
France	Men	..	..	..	30.3	28.8	25.0	20.7	23.7	16.0	10.7	13.6	..	..	..	..	**21.1**	0.34
	Women	..	..	..	36.1	32.0	30.3	29.7	28.6	21.8	22.8	21.4	..	..	..	..	**27.8**	0.19
	Difference m-w	..	..	..	*–5.8*	*–3.2*	*–5.3*	*–9.0*	*–4.9*	*–5.8*	*–12.0*	*–7.8*	..	..	..	..	*–6.7*	
Germany	Men	..	..	..	9.6	13.1	15.5	18.7	21.0	23.5	27.6	26.7	..	..	..	..	**19.5**	0.33
	Women	..	..	..	17.3	20.6	27.6	30.3	28.5	23.4	29.0	21.1	..	..	..	..	**24.7**	0.19
	Difference m-w	..	..	..	*–7.7*	*–7.5*	*–12.1*	*–11.5*	*–7.5*	*0.1*	*–1.5*	*5.6*	..	..	..	..	*–5.3*	
Greece	Men	..	..	..	31.4	34.1	26.0	22.4	27.5	28.1	26.8	26.6	..	..	..	..	**27.9**	0.13
	Women	..	..	..	32.0	39.3	32.7	30.5	26.5	34.0	26.6	22.6	..	..	..	..	**30.5**	0.17
	Difference m-w	..	..	..	*–0.6*	*–5.3*	*–6.7*	*–8.0*	*1.0*	*–5.9*	*0.2*	*4.0*	..	..	..	..	*–2.7*	
Hungary[3]	Men	..	36.2	41.5	38.3	54.3	31.6	29.6	..	..	..	..	..	..	..	..	**38.6**	0.23
	Women	..	47.5	41.3	46.5	54.2	43.9	37.4	..	..	..	..	..	..	..	..	**45.1**	0.13
	Difference m-w	..	*–11.3*	*0.2*	*–8.1*	*0.1*	*–12.3*	*–7.8*	..	..	..	..	..	..	..	..	*–6.5*	
Ireland	Men	..	..	..	29.9	32.5	26.4	30.4	25.8	25.9	22.3	31.4	..	..	..	..	**28.1**	0.13
	Women	..	..	..	35.5	39.8	35.3	42.9	38.1	35.0	33.5	31.4	..	..	..	..	**36.4**	0.10
	Difference m-w	..	..	..	*–5.6*	*–7.3*	*–8.9*	*–12.4*	*–12.4*	*–9.1*	*–11.2*	*0.0*	..	..	..	..	*–8.4*	
Italy	Men	..	..	..	34.6	28.3	28.5	27.5	25.4	24.4	25.8	30.6	..	..	..	..	**28.1**	0.12
	Women	..	..	..	39.3	32.2	30.5	33.6	32.5	29.4	25.2	30.2	..	..	..	..	**31.6**	0.13
	Difference m-w	..	..	..	*–4.6*	*–4.0*	*–2.0*	*–6.1*	*–7.1*	*–5.0*	*0.6*	*0.4*	..	..	..	..	*–3.5*	

Table 3. **Gross wage *premia* on upper-secondary education for men and women in 21 OECD countries, 1991-2005** *(cont.)*

Average percentage changes from wage of persons with less than upper-secondary degree[1, 2]

		1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	Multi-period average	Cross-period coefficient of variation
Luxembourg[3]	Men	. .	. .	. .	. .	66.4	50.6	49.4	43.7	50.0	52.0	52.3	. .	. .	. .	. .	**52.1**	0.13
	Women	. .	. .	. .	. .	72.6	68.7	52.2	39.1	47.9	57.6	50.4	. .	. .	. .	. .	**55.5**	0.21
	Difference m-w	. .	. .	. .	. .	*–6.2*	*–18.1*	*–2.8*	*4.6*	*2.1*	*–5.6*	*1.9*	. .	. .	. .	. .	*–3.4*	
Netherlands	Men	. .	. .	. .	40.4	29.0	33.1	29.9	28.2	24.3	26.0	30.1	. .	. .	. .	. .	**30.1**	0.16
	Women	. .	. .	. .	38.0	32.1	28.1	30.2	24.4	21.7	21.2	17.4	. .	. .	. .	. .	**26.7**	0.25
	Difference m-w	. .	. .	. .	*2.4*	*–3.1*	*5.0*	*–0.3*	*3.8*	*2.6*	*4.8*	*12.7*	. .	. .	. .	. .	*3.5*	
Poland	Men	. .	. .	. .	. .	. .	. .	35.0	31.2	36.1	34.1	. .	. .	. .	. .	. .	**34.1**	0.06
	Women	. .	. .	. .	. .	. .	. .	30.7	30.1	34.3	35.3	. .	. .	. .	. .	. .	**32.6**	0.08
	Difference m-w	. .	. .	. .	. .	. .	. .	*4.4*	*1.1*	*1.7*	*–1.2*	. .	. .	. .	. .	. .	*1.5*	
Portugal	Men	. .	. .	. .	77.8	74.8	61.3	72.6	41.9	54.7	44.8	55.9	. .	. .	. .	. .	**60.5**	0.23
	Women	. .	. .	. .	98.5	92.8	86.0	84.7	53.3	59.0	48.2	57.8	. .	. .	. .	. .	**72.5**	0.27
	Difference m-w	. .	. .	. .	*–20.7*	*–18.0*	*–24.8*	*–12.1*	*–11.5*	*–4.3*	*–3.4*	*–1.8*	. .	. .	. .	. .	*–12.1*	
Spain	Men	. .	. .	. .	42.8	43.9	45.6	49.3	49.5	42.3	41.4	31.4	. .	. .	. .	. .	**43.3**	0.13
	Women	. .	. .	. .	47.3	47.4	39.9	47.3	50.2	45.5	40.6	28.6	. .	. .	. .	. .	**43.3**	0.16
	Difference m-w	. .	. .	. .	*–4.5*	*–3.5*	*5.7*	*2.1*	*–0.7*	*–3.2*	*0.8*	*2.8*	. .	. .	. .	. .	*–0.1*	
Sweden[3]	Men	. .	. .	. .	. .	. .	. .	14.3	14.8	16.5	20.0	19.3	. .	. .	. .	. .	**17.0**	0.15
	Women	. .	. .	. .	. .	. .	. .	11.3	8.8	14.7	8.9	13.3	. .	. .	. .	. .	**11.4**	0.23
	Difference m-w	. .	. .	. .	. .	. .	. .	*3.0*	*5.9*	*1.8*	*11.1*	*6.0*	. .	. .	. .	. .	*5.6*	
Switzerland[3]	Men	. .	. .	. .	. .	. .	. .	. .	. .	106.7	85.2	. .	. .	. .	. .	. .	**96.0**	0.16
	Women	. .	. .	. .	. .	. .	. .	. .	. .	71.6	46.3	. .	. .	. .	. .	. .	**58.9**	0.30
	Difference m-w	. .	. .	. .	. .	. .	. .	. .	. .	*35.1*	*38.9*	. .	. .	. .	. .	. .	*37.0*	
United Kingdom	Men	32.2	48.9	49.4	56.7	47.8	52.5	47.8	43.5	44.7	45.9	42.0	39.5	43.5	43.4	. .	**45.6**	0.13
	Women	38.8	47.3	53.1	55.2	53.0	56.1	47.6	46.3	45.2	42.4	39.3	41.1	42.2	43.7	. .	**46.5**	0.13
	Difference m-w	*–6.6*	*1.6*	*–3.6*	*1.5*	*–5.2*	*–3.5*	*0.2*	*–2.9*	*–0.5*	*3.4*	*2.7*	*–1.6*	*1.3*	*–0.3*	. .	*–1.0*	
United States	Men	. .	. .	. .	95.9	90.1	85.6	95.2	86.7	92.5	94.8	91.5	89.9	82.1	85.5	84.1	**89.5**	0.05
	Women	. .	. .	. .	82.8	88.9	78.2	83.7	87.8	83.4	92.2	92.3	91.1	76.3	79.1	82.8	**84.9**	0.07
	Difference m-w	. .	. .	. .	*13.1*	*1.2*	*7.4*	*11.5*	*–1.1*	*9.2*	*2.6*	*–0.9*	*–1.2*	*5.8*	*6.4*	*1.4*	*4.6*	

. .: Means "not available".

1. Effect of changing variable *edu1* from 0 to 1, leaving other variables unchanged. As the Mincer equations use the average upper-secondary degree holder as the reference person, this requires transforming the initial (negative) Mincer coefficient for men, α_1, into $[1/\exp(\alpha_1)–1]*100\%$ and that for women, $(\alpha_1 + \alpha_3)$, into $[1/\exp(\alpha_1 + \alpha_3)–1]*100\%$. Example: $\alpha_1 = –0.30$ would translate into an upper-secondary wage premium of 35% as $[1/\exp(–0.30)]–1 = 0.35$.
2. t-values of the underlying point estimates not reported. All coefficients for men are significant at the 1% level except for Belgium in 2000 (5% level). Wage *premia* for women are also significantly different from zero but not all of them are significantly different from male *premia*.
3. Estimation based on net hourly wages and, hence, results not directly comparable with those for other countries.

Source: Authors' calculations.

The gross hourly wage premium per annum of tertiary education

Two adjustments are made for the results to be fully comparable across countries. First, a correction is required to assess the gross wage premium for the four countries for which only net wages are available. Second, all gross wage *premia* upon completion of tertiary education are transformed into wage *premia per annum* of tertiary education to account for cross-country differences in the duration of tertiary studies.

The correction of net attainment *premia* follows de la Fuente and Jimeno (2005). The average income tax rate *t* of the average earner in 2001 is taken from OECD (2005a).[24] It equals 33.3% for Hungary, 27.8% for Luxembourg, 32.4% for Sweden, and 22% for Switzerland. The precise wage *premia* derived from the estimated *edu*3 coefficients in the log net wage equations are divided by $(1 - t)$ to obtain gross tertiary wage *premia*. This brings Hungary to the top of the country ranking. Moreover, Sweden is no longer at the bottom.

To express the results as gross wage *premia per annum* of tertiary education, the country-specific average tertiary study duration *d* is taken from Table B1.3b of OECD (2005b) and applied to the wage *premia* reported in Table 2.[25] This is done by taking (1 + wage premium) to the power of $(1/d)$, *i.e.* assuming a constant percentage increase in the hourly wage for each year of tertiary education. Figure 6 summarises the results for 2001, illustrating both the attainment-specific wage *premia* and the wage *premia per annum* of tertiary education. The latter enter the calculation of the private internal rates of return to tertiary education in Boarini and Strauss (2009).

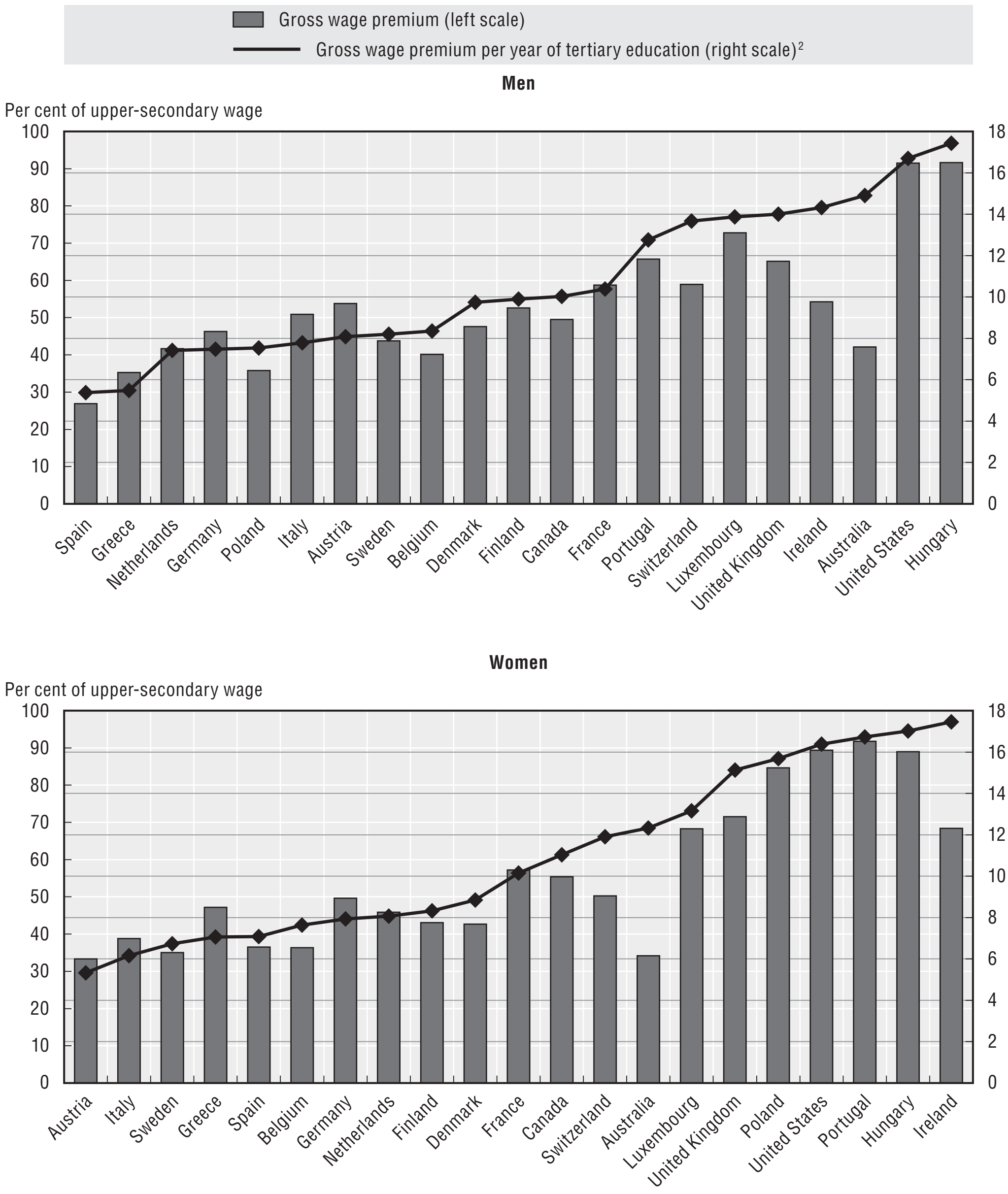
The country average of the gross hourly wage premium *per annum* is 10.6% for men and 11% for women. At 10% for each gender, the median is lower than the average, implying that the majority of countries is below the average. The median countries are Finland (men) and France (women). Wage *premia* are comprised between 5½% to 6% (men in Greece and Spain, women in Austria and Italy) and 16% to 18% in Hungary and the United States (both men and women) and for women in Ireland and Portugal. Women in Poland and the United Kingdom come close to these top *premia*.

Intermediate values of gross hourly wage *premia per annum* of tertiary education are observed for the majority of countries. For male wage *premia*, countries fall into three groups. The first with *premia* distinctly lower than the median includes Austria, Belgium, Germany, Italy, the Netherlands, Poland and Sweden. The second has *premia* at 10% (Canada, Denmark, Finland and France). The remaining countries (third group) outperform both the median and the average. For women, intermediate *premia* fall into two country groups, one with median or higher wage *premia* (Australia, Canada, France, Luxembourg and Switzerland), the other with *premia* from 7% to 9%.

Controlling for the duration of tertiary studies significantly changes the position of some countries. Australia and Ireland, where studies tend to be shorter than the OECD average, are now among those with the highest *premia*. Switzerland and the United Kingdom also improve their relative position. At the same time, countries characterised by long study duration such as Austria, Germany, Greece and Italy, fall behind. The same is true, albeit to a lesser extent, for France and the Netherlands where the average study duration is about half a year longer than the OECD average. The negative Pearson rank correlation between study duration and the wage premium *per annum* of tertiary education at –0.70 for men and –0.65 for women suggests decreasing returns to additional years of tertiary education beyond the first higher-education degree. Countries with long average study duration even fail to produce higher tertiary attainment *premia* (not controlling for duration) than those with

Figure 6. **Gross wage *premia***

2001[1]



1. Except Hungary (1997); and Poland and Swizerland (2000).
2. The total wage premium associated with a tertiary education level is converted to an annual basis by taking into account the duration of tertiary studies.

Source: Authors' calculations.

shorter programmes, with the coefficient of rank correlation between the 2001 tertiary attainment *premia* (see Table 2) and study duration being insignificant at best (–0.20 for men and –0.26 for women). These results imply that countries with long study duration may have scope for strengthening the overall incentive to invest in tertiary human capital through curricular reform, *e.g.* by streamlining and better co-ordinating study programmes, reducing slack in student timetables, and strengthening incentives for studying faster.

Comparison with other estimates in the literature

Given the vast amount of country-specific empirical studies, the focus here is on a limited number of empirical studies, which like this study have attempted to yield comparable results across countries through the use of cross-country data sources and a unified framework of sample selection and econometric specification.[26] Despite the similarities the comparability across studies is limited by differences in time and country coverage, data sources and methods. Bearing this *caveat* in mind, the following comparisons suggest that the results of this study are broadly in line with those from earlier studies.

Psacharopoulos (1994) and Psacharopoulos and Patrinos (2004) have the broadest country coverage. Where they present university-specific results these pertain to returns to education rather than wage *premia* and mostly refer to sample periods prior to those covered in this study. However, they also present results from traditional *Mincerian* wage equations that include years of schooling (all levels) and labour market experience. The schooling coefficients average 0.072 for the 16 countries for which the year reported (early 1990s to mid-1990s, see Table A2 in Psacharopoulos and Patrinos, 2004) is not too far from the results reported here, compared with 0.100 in this study (gender-country average of *edu*3 coefficients of 1996 or closest available year, divided by 4.21 years, the OECD average duration of tertiary studies).

The *Mincerian* years-of-schooling coefficients (all educational levels) based on data of the mid-1990s collected in Asplund and Pereira (1999) and also reported in Harmon *et al.* (2003, Table 2) average 0.075 for men and 0.083 for women (specification using potential experience) for the 14 countries that their samples and those of this study have in common. This compares with a somewhat higher 0.087 in this study (gender-country average of *edu*3 coefficients of 1996 or closest available year, divided by 4.6 years, the average study duration in the relevant country group).

Blöndal *et al.* (2002) compute private internal rates of return to tertiary education at the end of the 1990s for ten countries, eight of which are also in the country sample reported here. Methodologically, their "narrow rate" comes closest to the wage premium *per annum* of tertiary education reported in Figure 6. The average gross wage premium *per annum* of education in this group is 11.9% for men and 11% for women and compares to the 2001 average premium of 10.5% for each gender in this study. The results are relatively close to each other given marked differences in data sources, available control variables and methodology (annual rather than hourly earnings; returns not estimated through regressions but taken from aggregate data on average gender-age-specific earnings of tertiary- and upper-secondary degree holders).

De la Fuente and Jimeno (2005) use data from labour-force surveys, a quadratic specification of potential experience and a smaller set of control variables than this study. Their uncorrected[27] OLS estimate for the EU15 countries averages 0.08, comparable with the 2001 estimate of 0.085 presented here (gender-country average of *edu*3 coefficients, divided by the average study duration of 4.7 years). Correlation analysis confirms that the new results presented in this study are broadly in line with comparable cross-country studies in the earlier literature. Based on the same sample of countries and using the same reported year (or its closest neighbours) for each cross-study comparison, the correlation coefficients between our tertiary wage *premia* and other studies are 0.91 with respect to Blöndal *et al.* (2002); 0.70 with respect to Asplund and Pereira (1999); 0.6 with respect to

Psacharopoulos and Patrinos (2004). Finally, the correlation coefficient between our Mincer coefficients and those reported in the *Handbook of the Economics of Education* (Peracchi, 2006) amounts to 0.68.

Conclusion

The gross hourly wage premium is the single most important driver of private returns to tertiary education. This study has presented cross-section estimations based on a unified framework for 21 OECD countries from the 1990s to the early 2000s using international (and a few national) household surveys to maximise international comparability. One of the main advantages of the estimates presented here have been the use of a richer set of control variables and the extension of a single framework to a larger number of OECD countries than had usually been the case. The results of the "augmented" *Mincerian* wage equations point to an average gross hourly wage premium on completed tertiary education of 55% in 2001 (country-gender average), translating into a premium of almost 11% *per annum* of tertiary education. Wage *premia* display little variation over time but huge cross-country variation, ranging from 27% for men in Spain to 90% for Hungary and the United States. At 6 %, the premium *per annum* of tertiary education is lowest in Greece and Spain while reaching 14%-18% in most Anglo-Saxon countries as well as in Portugal and Hungary.

As discussed in Oliveira Martins *et al.* (2009, in this issue), it is remarkable that wage *premia* have held up well and even tended to increase over time despite the ongoing expansion of tertiary education. This indicates that the demand for high-skilled workers has increased at least as much as their supply. The assessment of wage *premia* on tertiary education should be based on whether they are conducive to the right skill mix in the economy. Very low wage *premia* are either indicative of low quality of tertiary education or insufficient wage differentiation, leading to low demand for tertiary education and under-provision of skills. But very high wage *premia* are not all good either since they may reflect an inability of the education system to provide graduates despite strong signals from the labour markets.

Finally, it should be borne in mind that the financial return expressed in the wage *premia* presented in this study are not the only benefits attached to higher education. As stated in the Stiglitz commission report, education is one of the great pillars of well-being. It matters not only for financial returns but also for the quality of life. Even though clear causality is difficult to establish, "better-educated people typically have better health status, lower unemployment, more social connections, and greater engagement in civic and political life" (Stiglitz Commission, 2009, p. 46).

Notes

1. See Sianesi and Van Reenen (2003) for a survey of empirical studies on macroeconomic returns to education.
2. The other large area of policymaking in this respect is university access policies: individuals may be constrained either by a lack of necessary educational credentials (*e.g.* in countries rationing access to upper-secondary attainment) or by a lack of liquidity. See Oliveira Martins *et al.* (2009) for a joint empirical analysis of demand and supply-side determinants of investment in tertiary education and how policies affect them.
3. See Psacharopoulos and Patrinos (2004) for a survey of the empirical literature on Mincer equations.

4. For more details see Heckman *et al.* (2005) where these two interpretations of the Mincer specification are discussed as the "Compensating Differences Model" and the "Accounting-Identity Model".
5. In the presence of measurement error, instrumenting the missing ability variable with family background is likely to aggravate the measurement error bias (Mellander and Sandgren Massih, 2008).
6. The size of this general-equilibrium effect is somewhat controversial. While Heckman *et al.* (1999) find the graduate-wage-depressing labour-supply effect of graduation to be large enough to undo discounted net lifetime income gains, Lee (2005) finds an only mild reduction in these gains from the labour-supply effect.
7. The correlation coefficient between attainment and obtaining an indefinite-term contract is significant only in about half the countries and is rarely larger than 0.10. Attainment is more systematically correlated with working in the public sector, with the correlation coefficient averaging some 0.2. Finally, the correlation between attainment and firm size is fairly systematic, too, albeit less strong (coefficient of 0.11 on average). A careful country-by-country tabulation of all bilateral correlations among independent variables can be found in the working paper version of this study (Strauss and de la Maisonneuve, 2007).
8. For an in-depth description of the data sources, data issues and the construction of variables see Strauss and de la Maisonneuve (2007).
9. What matters for the return to tertiary education is the additional net wage at tertiary level compared with that at upper-secondary level. This requires the marginal income tax rate of an upper-secondary degree holder to be applied on the gross wage premium (Boarini and Strauss, 2009).
10. As the intention is to use wage *premia* of today's workers to gauge the profitability of tertiary education for today's students contingent on normal labour market participation, dropping workers with less than 15 hours per week is appropriate based on the assumption that their decision to work few hours is voluntary.
11. For France and the Netherlands the attainment variable is corrected for errors in the raw data (see Strauss and de la Maisonneuve, 2007).
12. Returns to tertiary education depend not only on study completion but also on the type of degree and the field of study (*e.g.* see Borland 2002 and Stark 2006 on empirical results for Australia and Canada, respectively). However, information on study fields is missing in most of the datasets analysed here.
13. Only for Australia is the actual number of years worked directly available. For all other countries there is some deviation from the actual labour market experience. The error is expected to be small for men but larger for women. Thus it is expected that the female experience premium falls short of its male counterpart.
14. F-tests conclude that the joint hypothesis of the effects of all control variables jointly being the same for men as for women is rejected for each country. However, the subset of variables responsible for the rejection varies from country to country.
15. For example, persons in the ECHP whose answer corresponds to the plant-size class "between 20 and 49 persons" are randomly assigned an integer value from 20 to 49.
16. Dumont (2005) discusses the pros and cons of alternative assessment methods for over-qualification.
17. While the margin of 10% is arbitrary, not correcting at all does not seem useful either since simply taking the mode of the attainment distribution would lead to implausibly high shares of occupation-attainment mismatches.
18. Since educational requirements tend to increase over time, many older workers may appear to be under-qualified according to the simple metric used here but they compensate for this by longer experience in their jobs or on the labour market more generally. There is indeed positive correlation between experience and *underqualif* in the data, potentially reducing the impact of the experience variable. Conversely, the correlation between *exper* and *overqualif* is negative as expected – most overqualified persons tend to be young – but the correlation coefficient is much smaller in absolute terms and not significant for a number of countries.
19. However, estimations for Sweden are based on net wages. When correcting for the income taxes, the wage premium is lowest in Austria at 33%.
20. Exceptions to this rule with respect to the wage effect of tertiary education include women in Austria (much lower *premia* in 2000 and 2001 than before), Belgium (one-off jump in 1999), Greece

(one-off jump in 1998), Ireland (jump in 1997), Luxembourg (1995 premium much higher than for all other years), Portugal (one-off jump in 2000). Results for men are very stable apart from drops in *premia* in 2001 for both Poland and Portugal. A sudden increase of about 15 percentage points is observed for both genders in Canada in 2000.

21. That said, labour market institutions such as minimum wages and collective bargaining agreements shield lower-skilled persons to some extent from large swings in wage *premia*. However, the cyclical vulnerability of the low-skilled shows very clearly in the employment probability (see Boarini and Strauss, 2009, in this issue).
22. The precise wage effect of a 0-1 change in any dummy variable reported in this paragraph is obtained in the same way as the wage effect of completing tertiary education.
23. The interpretation of this result is that employees of the public sector do not appear to "pay the price" for their much higher job security in most European countries, at least until 2001.
24. Admittedly, correcting with the same average tax rate for all individuals is a rough correction. To the extent that persons with higher incomes have higher average tax rates than persons with lower income, the gross *premia* obtained in this way might slightly understate the true wage *premia*.
25. Missing study-duration data for six countries (Belgium, Canada, Luxembourg, Poland, Portugal, and the United States) are replaced with the simple OECD country mean (4.21 years). For six countries, study duration refers to the year 2001/02, for the others to 1994/95.
26. Country-specific studies can exploit different micro data sources that are not necessarily comparable with the survey data used in this study. A case in point is Ciccone *et al.* (2004) who report a net hourly wage premium of 9% for Italy, in contrast with the 7% for the gross wage premium reported in this study.
27. The final gross wage *premia per annum* of education in de la Fuente and Jimeno (2005) contain an *ad hoc* correction factor of 0.9 accounting for the likely net endogeneity bias (see Card, 1999).

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ANNEX A1

Data Sources, Sample Size and Descriptive Statistics of the Gross Hourly Wage Rate

Table A1.1 shows the international panel data bases used in this study, their original national sources, their sizes as well as the time coverage and sample sizes of the analysis in this study.

Table A1.1. **Micro data bases and sample size**

	Panel database[1]	Original national source[1]	Number of waves used	Starting year	Latest available year	Number of individuals in 2001	Size of basic sample[2] in 2001[3] *Mincerian* wage equations
Australia	HILDA	HILDA	3	2001	2003	19 914	5 211
Austria	ECHP	ECHP	7	1995	2001	5 605	2 262
Belgium	ECHP	ECHP	8	1994	2001	4 299	1 896
Canada[4]	CNEF	SLID	10	1993	2002	..	25 555
Denmark	ECHP	ECHP	8	1994	2001	3 789	2 098
Finland	ECHP	ECHP	6	1996	2001	5 637	2 513
France	ECHP	ECHP	8	1994	2001	10 119	3 892
Germany	ECHP	GSOEP	8	1994	2001	10 624	5 003
Greece	ECHP	ECHP	8	1994	2001	9 419	2 287
Hungary	CHER	HHS, HHBS	6	1992	1997	3 626	889
Ireland	ECHP	ECHP	8	1994	2001	4 022	1 557
Italy	ECHP	ECHP	8	1994	2001	13 392	4 174
Luxembourg	ECHP	PSELL	7	1995	2001	4 916	2 356
Netherlands	ECHP	ECHP	8	1994	2001	8 608	2 670
Poland	CHER	HBS	4	1997	2000	7 747	2 286
Portugal	ECHP	ECHP	8	1994	2001	10 915	4 146
Spain	ECHP	ECHP	8	1994	2001	11 964	3 939
Sweden	ECHP	NSLC	5	1997	2001	9 291	4 551
Switzerland	CHER	SHP	2	1999	2000	6 835	2 988
United Kingdom	BHPS	BHPS	14	1991	2004	18 867	8 078
United States	CPS	CPS	12	1994	2005	128 821	49 592

1. The sources are:
   HILDA: Household, Income and Labour Dynamics in Australia (see Watson, 2005).
   ECHP: European Community Household Panel (see Eurostat, 2003).
   CNEF: Cross-National Equivalent File.
   SLID: Survey of Labour and Income Dynamics.
   GSOEP German Socio-Economic Panel.
   CHER: Consortium of Household Panels for European Socio-Economic Research (see Birch *et al.*, 2003).
   HHS: Hungarian Household Survey.
   HHBS: Hungarian Household Budget Survey.
   PSELL: Panel socio-économique “Liewen zu Lëtzebuerg”.
   HBS: Household Budgets Survey.
   NSLC: National Survey on Living Conditions.
   SHP: Swiss Household Panel.
   PSID: Panel Study of Income Dynamics.
2. The basic sample is defined as the number of individuals with non-missing values for gender, educational attainment, and wage and for which the latter conforms to the definition of the variable given below.
3. Except Hungary (1997); Poland and Switzerland (2000).
4. For Canada, due to confidentiality, the number of individuals was not known.

Table A1.2 shows the sample size and descriptive statistics of the gross hourly wage rate, the left-hand-side variable used in the study, country by country for the year 2001 without distinguishing by gender. The average hourly wage is found to be highest in Switzerland and lowest in Hungary and Poland. Wage dispersion is found to be largest in the United States, Canada, Hungary and Portugal, and lowest in Denmark.

Table A1.2. **Descriptive statistics of gross hourly wage rate[1] for 21 OECD countries, 2001[2]**

In US$ PPP

	Mean	Standard deviation	Coeff. of variation	Minimum	Maximum	10th percentile	90th percentile	Ratio 90th/ 10th percentile
Australia	14.5	7.9	0.54	1.0	129.7	7.5	22.6	3.0
Austria	11.8	5.2	0.44	1.2	43.5	6.7	18.2	2.7
Belgium	15.2	6.6	0.44	1.8	84.4	8.9	22.7	2.6
Canada[3]	16.5	14.0	0.84	..	..	..	..	..
Germany	12.5	6.1	0.49	1.0	77.6	5.6	19.7	3.5
Denmark	17.3	6.0	0.35	2.4	69.1	11.2	24.6	2.2
Spain	10.3	6.4	0.62	1.4	78.6	5.0	18.3	3.7
Finland	13.0	6.1	0.47	1.7	95.5	8.0	19.9	2.5
France	12.8	7.2	0.56	1.1	139.7	6.7	20.7	3.1
United Kingdom	14.9	8.7	0.59	1.1	166.0	7.1	25.0	3.5
Greece	8.3	4.7	0.57	1.2	55.0	4.2	14.4	3.5
Ireland	13.4	8.1	0.60	1.5	94.6	6.5	23.5	3.6
Italy	11.1	5.4	0.49	1.2	72.8	6.6	17.0	2.6
Luxembourg	*14.6*	*8.5*	0.58	*1.1*	*88.0*	*6.8*	*24.9*	3.7
Netherlands	17.5	10.1	0.58	1.1	187.0	9.6	26.5	2.8
Portugal	6.6	5.3	0.81	1.0	68.7	3.1	12.2	3.9
Sweden	*8.3*	*4.1*	0.49	*1.0*	*120.0*	*5.0*	*11.8*	2.4
Hungary	*2.5*	*2.1*	0.83	*0.5*	*24.7*	*1.1*	*4.2*	4.0
Poland	3.7	2.6	0.70	0.5	31.2	1.7	6.3	3.6
Switzerland	*18.2*	*10.0*	0.55	*1.0*	*139.1*	*7.8*	*29.1*	3.7
United States	16.2	16.1	0.99	1.0	188.2	4.8	28.9	6.0

1. Hungary, Luxembourg, Sweden and Switzerland: net wage.
2. Except Hungary (1997); and Poland and Switzerland (2000).
3. For Canada, due to confidentiality, not all descriptive statistics are available.

Source: ECHP, CHER, BHPS, CPS, CNEF and HILDA.