

Revisiting the returns to schooling in the Portuguese economy

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April 2017

Abstract

This paper provides an overview of the evolution of the private returns to schooling in the Portuguese economy along the 1986-2013 period. We estimate the returns separately for men and women, at the mean and along the conditional wage distribution. Returns to schooling are found to be high, particularly for women, and to increase along the distribution. Our results point that the magnitude of the returns increased throughout the 1986-2013 period, but particularly in the 1990s. We also provide estimates of the relative wage premium associated with specific levels of educational attainment. The returns are highest for tertiary education. In the first decades under analysis, relative wage premia associated with the 9th grade stand above those estimated for secondary education, whereas in the most recent period these differences are negligible. All in all, our results suggest that education remains a profitable investment for individual agents. This is a valuable piece of information also for policymakers, who should take it into account together with the social returns to education when designing policies and incentive schemes. (JEL: I26, J31, C21)

Introduction

As formalized in Becker (1962), the assessment of the private returns to schooling provides a key piece of information for the individual decision determining the optimal level of investment in formal education. Regardless of the potential social returns to education, information on private returns is also relevant for policymakers, guiding them in the design of programs and incentive schemes to promote individual investment in education.

Individual returns to schooling are typically measured as the proportional increment in earnings resulting from an additional year of education on the basis of the so-called Mincerian wage equation (Mincer (1974)). There is a wide strand of empirical literature shedding light on the magnitudes and

Acknowledgements: We thank comments and suggestions by Nuno Alves, Cláudia Braz, Isabel H. Correia, Andy Garin, Miguel Gouveia, Pedro Portugal, Pedro Raposo and participants in a seminar held at Banco de Portugal. The views herein expressed are our own and do not necessarily coincide with those of Banco de Portugal or the Eurosystem.

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explanatory factors of returns to schooling in both advanced and emerging economies. Card (1999) provides a comprehensive review of existing literature on returns to schooling. Cross-country estimates are presented in, for instance, Psacharopoulos (1994), Martins and Pereira (2004), Psacharopoulos and Patrinos (2004) and Montenegro and Patrinos (2014).

International comparisons show that returns obtained for Portugal rank high among other European Union countries. Their evolution over the last decades may however have changed reflecting important reshufflings in the educational composition of the labour force that may have affected the way the market values education and specific schooling levels. In particular, female participation increased considerably in the last decades and women are increasingly more educated than men. More generally, the overall supply of workers completing tertiary and secondary education sharply increased, reflecting higher individual investment in university degrees, particularly since mid-1990s. At the same time, there was a strong reduction in the percentage of individuals with less than 9 years of education, reflecting the enactment of specific legislation extending compulsory schooling.¹

This paper aims to complement the existing evidence on returns to schooling in the Portuguese economy and provide an overview of how they have changed since the late 1980s. In particular, we use *Quadros de Pessoal* data (QP henceforth) spanning the 1986-2013 period to estimate the returns to schooling separately for men and women, at the mean and along the conditional wage distribution. We also provide estimates of the relative wage premium associated with specific levels of educational attainment. The main goal of this paper is to provide a comprehensive description of the evolution of returns to schooling in this period, without claiming a causal relationship between schooling and earnings.

In broad terms, our results may be summarized as follows: the returns to schooling are found to be high, particularly in the case of women, and to increase along the distribution. The returns are highest for tertiary education. In the first decades under analysis, relative returns to the 9th grade stand above those estimated for secondary education. In the most recent period these differences are negligible, in line with the typical evolution in advanced economies (Montenegro and Patrinos (2014)). The detailed analysis undertaken in this paper allows the pinpointing of exceptions to these general findings.

1. Note that in spite of the remarkable changes underwent by the Portuguese economy in recent decades, the share of workers with at least secondary education remains the lowest in the European Union. According to Labour Force Survey data made available by the Eurostat, in 2016, 46.8 per cent of employed individuals in Portugal had completed either secondary or tertiary educational attainment. This figure compares with 73.4 per cent in the European Union as a whole.

Our results are broadly in line with previous studies on the estimation of the private returns to schooling in the Portuguese economy. Vieira (1999), using QP data for the 1982-1992 period, found evidence of returns to schooling of approximately 7 per cent at the mean of the wage distribution. Acknowledging that conventional estimates based on Mincerian equations are hampered by the so-called “ability bias”, Vieira (1999) attempts to circumvent it by estimating the returns to schooling using instrumental variables (IV). Specifically, the author uses changes to compulsory schooling legislation as an exogenous source of variation in educational attainment. This results in lower - albeit still positive - returns to education. Sousa *et al.* (2015) also focus only on returns at the mean of the distribution. Using QP data spanning the 1986-2009 period and a standard Mincer equation, they found returns of 10.0 per cent in the case of men and close to 10.5 per cent for women in the last year under analysis. Sousa *et al.* (2015) also use IV, presenting results based on three different instruments: changes to compulsory education, quarter of birth and the average education by region in the year the individual first entered school. In this case estimates of returns to schooling are higher than those obtained using OLS, but show a similar evolution over time.

There are other studies that assess the returns at different points of the conditional wage distribution - not only at the mean. Machado and Mata (1998), using QP data for the 1982-1994 period, found returns ranging from 4 to 11 per cent, respectively at its lower and upper part (and around 7-8 per cent at the mean). Similar evidence is provided in Hartog *et al.* (2001). In the latter case, however, the authors consider a richer set of covariates in the regressions, which yields slightly lower returns than in Machado and Mata (1998). Martins and Pereira (2004) also provide estimates of returns to schooling at different points of the distribution. Using the 1995 wave of QP, they find increasing returns along the distribution (of 6.5 and 14.5 per cent, respectively at the bottom and at the top of the distribution).

Alves *et al.* (2010) and Portugal (2004) provide estimates of the returns to tertiary education. In both cases, the authors find positive and significant returns benefiting individuals with university degrees (relative to non-university educated counterparts). Alves *et al.* (2010) provide estimates of the tertiary education wage premium at different points of the distribution and on the basis of QP data for 1982, 1995 and 2006. In the latter year, they find returns ranging from approximately 45 per cent to almost 100 per cent, respectively at lower and upper quantiles of the distribution.

It is worth highlighting that we do not resort to IV or control function methods for estimating the private returns to schooling. Estimates based on these methods are highly dependent on the sub-sample whose schooling attainment is affected by the change in the instrument chosen for the analysis. Different instruments yield different estimates of the returns to schooling and lead to different interpretations (Imbens and Angrist (1994)). Moreover, we are interested in providing a broad picture of the evolution of returns along the

1986-2013 period and, as shown in Sousa *et al.* (2015), relying in IV estimates does not change the overall evolution.

Finally, note also that our paper focuses only on the *private* (or individual) returns to education and does not address the *social* returns to education. The latter stem from the existence of positive externalities (from higher labour productivity, to lowering crime rates, improving overall health standards, decreasing mortality rates, or promoting better citizenship and voting decisions - refer to Lochner (2011) for a comprehensive review). Assessing these effects are out of the scope of this article.

The paper is organized as follows: Section 2 presents the data source, also providing a comprehensive analysis of descriptive statistics. Section 3 describes the theoretical framework underlying the estimations presented in the article, whereas Section 4 sheds light on the strategy used to implement the analysis. Section 5 lists the key results of the article. Finally, Section 6 presents the main conclusions and discusses topics relevant in terms of education policy.

Data description

Data are drawn from *Quadros de Pessoal*, a matched employer-employee dataset including a personal identification number that allows the tracing of individuals across time. The information is based on a compulsory survey conducted annually by the Ministry of Social Security. Data covers every establishment paying wages in the Portuguese private sector: general government, military staff, self-employed and household employees are thus excluded. The questionnaire covers attributes of workers and firms. Regarding the former, it includes information on gender, age, education, occupation, industry, tenure and earnings, among other dimensions. For the purpose of our analysis, we use data covering the 1986-2013 period (except 1990 and 2001 for which data are not available). We focus on a sub-sample made of full- and part-time employees aged between 16 and 65 years.

We define wages as the sum of every work-related category of income (including base salary, overtime pay, and other regular payments). Hourly wages are adjusted for the whole amount of working hours, both normal and supplementary. Real wages are computed on the basis of each year's Consumer Price Index (taking 1986 as the base-year). In QP, individual educational attainment corresponds to a categorical variable reporting the highest level completed.² An additional variable providing information on the minimum number of school years required to complete the highest educational level reported was also created.

2. More precisely, we consider the mode of the highest level of education reported throughout the panel.

Table 1 briefly describes selected QP waves used for the analysis. It provides evidence of a remarkable increase in the average schooling, from 5.6 to 9.9 years, respectively in 1986 and 2013. This reflects a significant drop in the share of employees reporting lower educational levels and a strong increase in the percentage with either secondary or tertiary education (Figure 1).

This evolution was particularly noticeable in the case of female employees. They are, on average, more educated than men throughout the entire period under analysis and this disparity widened in the last decade. In particular, the percentage of female employees with an university degree increased from 2.3 per cent in the 1986 wave of QP to 22.0 per cent in 2013 (panel D of Figure 1).

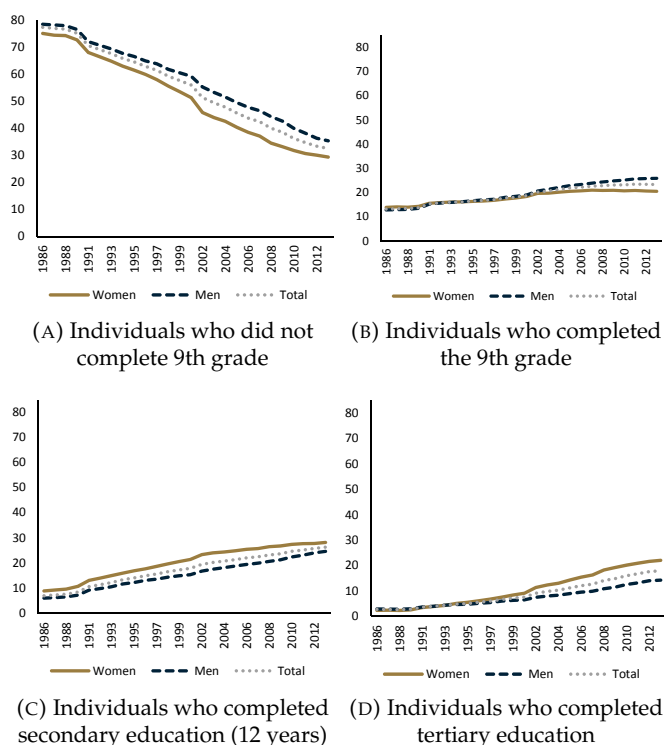


FIGURE 1: Employees by level of educational attainment (percentage).

Source: Authors' calculations based on *Quadros de Pessoal*.

In spite of their better educational endowments, QP data show that, on average, women earn consistently lower wages than male employees over the whole period (Figure 2). Nonetheless, although the two genders have experienced similar real wage increases in the first part of the 1986-2013 period, women's earnings have been growing more sharply than men's since 2000 (Figure 3). As emphasised in Cardoso *et al.* (2016), this evolution may be explained by a composition effect stemming from the higher educational

		1986	1991	1996	2000	2005	2010	2013
Women	Education (years)	5.73 [3.42]	6.54 [3.57]	7.37 [3.84]	8.16 [4.05]	9.17 [4.26]	10.09 [4.38]	10.39 [4.38]
	Age (years)	32.34 [10.2]	32.24 [10.26]	34.18 [10.15]	35.08 [10.15]	36.56 [10.17]	38.06 [10.31]	39.37 [10.21]
	Tenure (years)	8.55 [7.15]	7.34 [7.83]	7.77 [7.98]	7.21 [7.97]	7.21 [7.66]	7.61 [7.92]	8.53 [8.2]
	No. obs.	327,634	467,428	584,109	714,836	836,568	923,898	901,793
	% of total	33.1	36.7	39.9	41.5	42.2	45.2	47.3
Men	Education (years)	5.50 [3.33]	6.17 [3.5]	6.89 [3.7]	7.40 [3.81]	8.18 [3.96]	9.03 [4.08]	9.45 [4.08]
	Age (years)	36.30 [11.65]	36.35 [11.74]	36.88 [11.32]	37.27 [11.16]	37.79 [10.84]	39.01 [10.68]	39.92 [10.5]
	Tenure (years)	9.56 [8.05]	8.93 [8.71]	8.84 [8.62]	8.11 [8.52]	7.73 [8.07]	8.14 [8.27]	8.89 [8.54]
	No. obs.	662,723	806,480	880,628	1,009,561	1,144,560	1,118,236	1,003,012
	% of total	66.9	63.3	60.1	58.5	57.8	54.8	52.7
Total	Education (years)	5.6 [3.36]	6.31 [3.53]	7.08 [3.76]	7.71 [3.93]	8.60 [4.12]	9.51 [4.25]	9.9 [4.25]
	Age (years)	34.99 [11.35]	34.84 [11.39]	35.80 [10.95]	36.36 [10.8]	37.27 [10.58]	38.58 [10.53]	39.66 [10.37]
	Tenure (years)	9.23 [3.36]	8.34 [3.53]	8.42 [3.76]	7.73 [3.93]	7.51 [4.12]	7.90 [4.25]	8.72 [4.25]
	No. obs.	990,357	1,273,908	1,464,737	1,724,397	1,981,128	2,042,134	1,904,805

TABLE 1. Descriptive statistics.

Source: Authors' calculations based on *Quadros de Pessoal*.

Notes: Unless otherwise specified, the table reports averages (and standard-deviations in brackets). Variable "tenure" corresponds to the number of years working in the current firm.

level of the women joining the labour market. Indeed, on average, wages for university- educated women, who represent an increasing share of our

sample, grew more than wages of males with the same educational level (Figure 4).

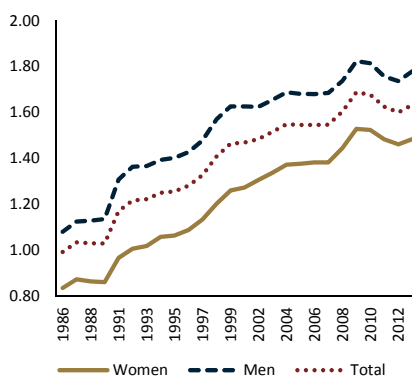


FIGURE 2: Average real wage per hour (in euro).

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The chart depicts the average real wage per hour worked in each wave of QP (deflated using CPI, 1986 base year).

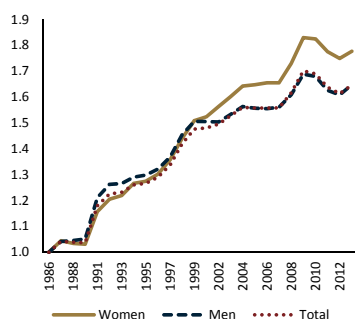


FIGURE 3: Real wage growth (index 1986=1).

Source: Authors' calculations based on *Quadros de Pessoal*.

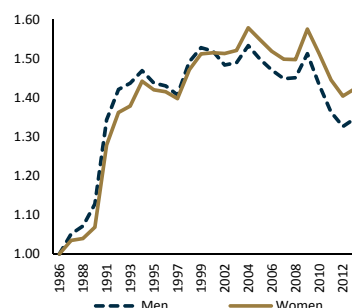


FIGURE 4: Real wage growth for workers with tertiary education (index 1986=1).

Source: Authors' calculations based on *Quadros de Pessoal*.

Figure 5 depicts the average real wage by educational level along the 1986-2013 period. As expected, wages increase with education but differences between workers with tertiary education and their less educated counterparts are particularly significant. This differential widened up to 1995, remained relatively constant up to mid-2000s and, more recently, it has been shrinking.

Focusing on workers with lower educational level, Figure 5 also points out that while in the late 1980s wages of individuals who did not complete the 9th grade were considerably below those referring to workers who did, this difference almost disappears in more recent QP waves. Conversely, whereas

in the beginning of the period average real wages were similar among workers with the 9th grade and those having completed secondary schooling, the gap between the two groups has been widening since the 1990s.

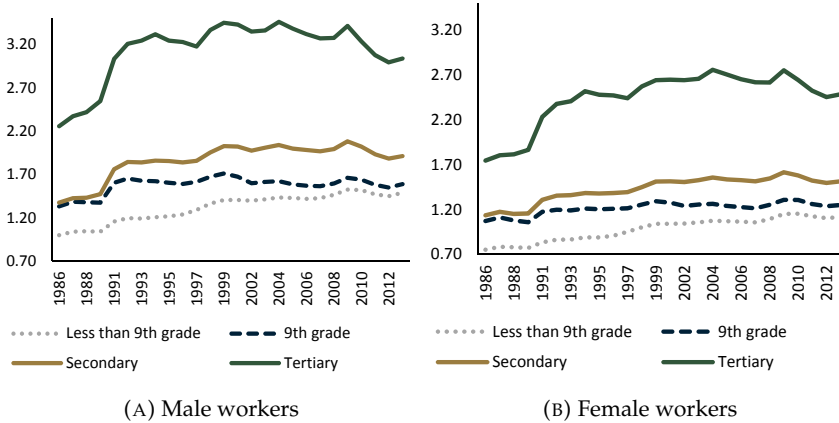


FIGURE 5: Average hourly wage by educational level.

Source: Authors' calculations based on *Quadros de Pessoal*.

For both genders the distribution of wages has shifted to the right and become slightly less compressed than in the late 1980s (details depicted in Appendix A). This increase in wage inequality was particularly noticeable at the upper tail of the distribution and only up to mid-1990s, having remained stable since then. The wage distribution is more compressed in the case of less educated individuals.

The educational *composition* of individuals across the distribution changed considerable along the 1986-2013 period. While in 1986 the share of workers with tertiary (or even secondary) educational attainment with below-median wages was low, it increases when focusing on the 2013 wave of QP (Figure 6). Although this may reflect a wide range of aspects and changes in the composition of private employment or in the productive structure of the economy, it can also be interpreted as a symptom of an over-education phenomenon.

Theoretical framework

Becker (1962) pioneered in applying utility theory to investment in education. In his framework the proportional pecuniary returns associated with educational attainment are a key component of the individual decision on whether and by how much to invest in human capital. In particular, individuals select the optimal number of years of schooling so as to maximise

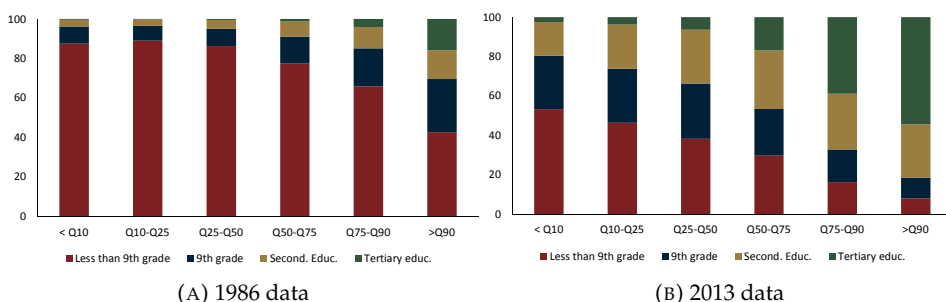


FIGURE 6: Educational composition of the wage distribution.

Source: Authors' calculations based on *Quadros de Pessoal*.

the discounted present value of future earnings net of the cost of schooling. This corresponds to an optimization problem whose solution is such that individuals would continue to invest in additional education up to the point where the marginal benefits match the marginal costs.

Mincer (1974) provided an empirical approximation to the marginal benefits' side of the individuals' optimization problem. In particular, the so-called traditional Mincerian wage equation corresponds to:

$$\ln y_i = \alpha + \beta S_i + \lambda_1 Exp_i + \lambda_2 Exp_i^2 + \varepsilon_i \quad (1)$$

where β corresponds to the pecuniary return from an additional year of formal education and Exp refers to individuals' experience in the labour market. As the bulk of literature on the relationship between education and earnings, we rely on Mincer's framework for the estimation of the returns to schooling.

Card (1999) made it clear that the decision on how much to invest in schooling is very much conditional on individual preferences and it is subject to individual heterogeneity, both in terms of the marginal returns to schooling (determined by, for instance, differences in individual ability) and in terms of its costs (accounting for differences in rates of substitution between schooling and future earnings on the basis of, *eg*, access to funds or personal taste). This implies a heterogeneous effects framework in which the way covariates affect wages varies across individuals.

Empirical strategy

To estimate the returns to schooling in the Portuguese economy we adopt the standard approach relying on Mincerian wage regressions such as the one in Equation (1). The regressions are run separately for each wave of QP, assuming

a cross-sectional set-up. We adopt a homogeneous effect framework, in the sense that we assume that the impact of schooling on wages of is the same for all individuals: $\beta_i = \beta_1 = \dots = \beta_N$, for all $i = 1, \dots, N$.

Moreover, we take individuals' age as a proxy for overall labour market experience and, in line with Mincer (1974), we also consider a second order polynomial. Our regressions include a set of other covariates: the individual's tenure in the current firm (also as a second order polynomial), the logarithm of the current firm's size and, when pooling data jointly for men and women, gender dummies (equal to one for male employees). We allow for some additional flexibility by running the wage regressions separately for men and women using the same set of covariates, except for the gender dummy.³

Returns to an additional year of schooling

Our baseline specification corresponds to

$$\ln y_i = \alpha + \beta S_i + \lambda_1 age_i + \lambda_2 age_i^2 + \mathbf{x}_i^T \gamma + \varepsilon_i \quad (2)$$

where y_i corresponds to individuals' real hourly wage (deflated using CPI taking 1986 as the base year) and S_i represents the minimum number of schooling years required to complete the highest level reported by the individual. As such, coefficient β , our parameter of interest, represents the per cent increase in hourly wage resulting from an additional year of schooling estimated using Ordinary Least Squares (OLS). Vector \mathbf{x}_i groups the set of observable characteristics aforementioned and estimates for the parameters in vector γ measure the respective marginal impact on y_i . Finally, the marginal impact of age is given by $\hat{\lambda}_1 + 2 * \hat{\lambda}_2 age_i$, where age_i refers to the worker's age.

In spite of adopting a homogeneous effect set-up, we allow for some heterogeneity in the returns by letting them change depending on the individuals' placement along the conditional distribution of wages. In particular, we also run our baseline specification within the Quantile Regression (QR) framework proposed in Koenker and Bassett (1978). This allows our covariates to affect the shape and tail behaviour of the conditional wage distribution and implies assuming

$$\ln y_i = \alpha_\theta + \beta_\theta S_i + \lambda_{1,\theta} age_i + \lambda_{2,\theta} age_i^2 + \mathbf{x}_i^T \gamma_\theta + \varepsilon_{\theta,i}, \quad (3)$$

where θ represents different quantiles of the conditional distribution of hourly wages: $\theta = \{0.1, 0.25, 0.5, 0.75, 0.9\}$. Therefore, β_θ corresponds to the return

3. For the purpose of assessing robustness of the estimates, we also ran regressions including industry and region controls. This brings down the magnitude of the coefficient associated with schooling attainment, but the evolution of returns over time is same (Appendix B)

to an additional year of schooling at the θ -th quantile of distribution of the logarithm of hourly wages conditional on the individuals' observed attributes. By assessing the returns to schooling at these different quantiles, we complement the evidence provided by OLS, which refers to the mean of the wage distribution.

Returns to specific education levels

In addition to the baseline specification, we also consider an alternative specification in which the highest completed level of schooling is included on the basis of dummy variables:

$$\ln y_i = \alpha + \sum_{j=2}^4 \beta_j E_{j,i} + \lambda_1 age_i + \lambda_2 age_i^2 + \mathbf{x}_i^T \gamma + \varepsilon_i, \quad (4)$$

where $E_j, j = \{1, 2, 3, 4\}$ are indicator variables that equal one for individuals reporting each of the following levels of schooling attainment: 1) less than the 9th grade; 2) 9th grade; 3) secondary education; and 4) tertiary education. The first category is omitted in the regressions. In this case, each $\beta_j, j > 1$ corresponds to the wage premium benefiting individuals holding schooling level j vis-à-vis comparable counterparts with less than the 9th grade ($j = 1$). We also implement this alternative specification within a QR framework:

$$\ln y_i = \alpha_\theta + \sum_{j=2}^4 \beta_{\theta,j} E_{j,i} + \lambda_{1,\theta} age_i + \lambda_{2,\theta} age_i^2 + \mathbf{x}_i^T \gamma_\theta + \varepsilon_{\theta,i}, \quad (5)$$

As pointed out by Card (2001), estimates of returns to schooling based on Mincerian wage equations may be hampered by two sources of bias. In the first place, there may be mismeasurements in terms of the individual schooling, in which case estimates of β would be downward biased. Although the possibility of measurement errors cannot be ruled-out, as we are using an administrative data source we are confident that erroneous cases are negligible in our sample.

The second source of bias arises from fact that we are not controlling for the whole set of individual-specific attributes that affect wages ("ability bias"). These factors - some of which are not observable - are included in error term ε_i . If they are also correlated with schooling attainment, generating endogeneity, the estimator would also be inconsistent. Since the standard Mincerian equation does not account for the impact of individual innate ability on wages and educational level, $\hat{\beta}$ would be upward biased. As we are interested in providing a broad overview of returns to schooling along the 1986-2013 period and not on analysing causal relationships, we do not apply methods such as IV or control function for circumventing these issues.

Moreover, estimates based on these methods are highly dependent on the subsample whose schooling attainment is affected by the change in the elected instruments (Imbens and Angrist (1994)).

Findings

Returns to an additional year of schooling

Figure 7 presents the estimates of returns to schooling obtained from running OLS regressions with specification (2) in each available wave of QP.⁴

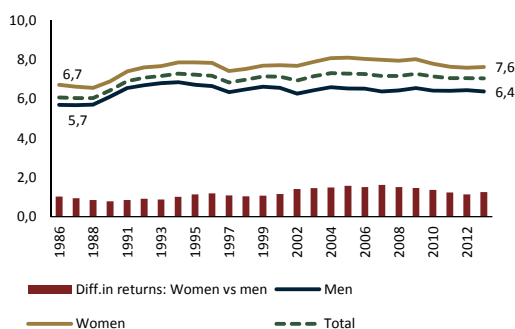


FIGURE 7: OLS-based returns to an additional year of schooling.

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The chart depicts the coefficient of S_i estimated on the basis of specification (2) using OLS. Coefficients are significant at the 1% level.

It provides evidence of positive returns to an additional year of schooling both in the case of men and women. In the latter case, the estimated returns are slightly higher, over the whole period in analysis: in 2013, an additional year of schooling is estimated to yield, on average, a 7.6 per cent increase in females' hourly wage, whereas for men the estimated increment stands at 6.4 per cent. The gender gap in the returns is statistically significant along the entire period and has remained relatively stable since 1986. Over time, there has been a slight increase in the returns for both genders. This increase was

4. All coefficients are significant. Table B.1 in Appendix B presents the full set of results of non-gender specific regressions for selected years. Results obtained for the whole 1986-2013 period, also split for male and female workers, are available upon request. Table B.1 also presents the results of the estimation of specification (2) controlling for industry and region effects. Introducing these additional covariates yields a decrease in the magnitude of the returns to schooling, but the overall picture does not change.

particularly marked along the 1990s and in more recent decades the returns remained relatively constant, albeit with a minor drop as of 2009.⁵

Results presented in Figure 7 refer to the estimates of returns to an additional year of schooling *at the mean* of the conditional wage distribution. Such evidence may hide important differences at different points of the distribution. By relying on the QR framework we are able to estimate the returns to schooling at different quantiles. Results obtained with this methodology are summarized in Figure 8.

Figure 8 indeed shows that the magnitude of the estimates for the returns to schooling changes considerably along the distribution. For instance, estimates obtained on the basis of OLS using the 2013 wave of QP and pooling data for both men and women point that an additional year of schooling implies a mean 6.4 per cent increase on wages. This figure masks the fact that, applying the same procedure to the same data but using the QR framework, one additional year of schooling yields a 3.1 per cent wage increase at the 1st decile of the distribution and a 8.8 per cent impact at the 9th.

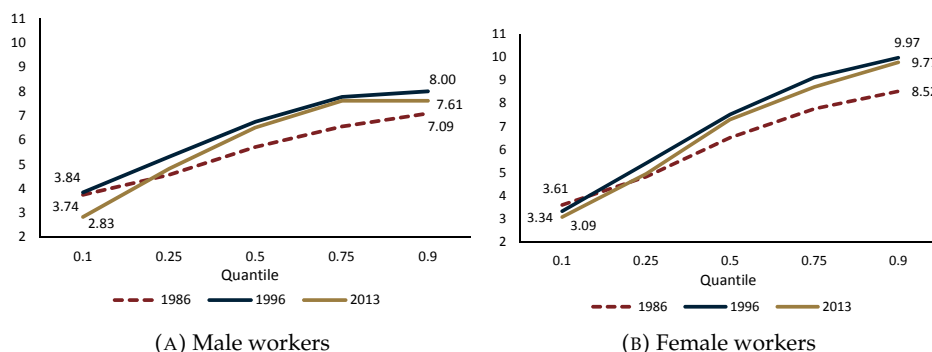


FIGURE 8: Returns to schooling across the wage distribution (per cent).

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The charts depict the per cent wage increment from an additional year of schooling, obtained on the basis of specification (3). For greater detail refer to Appendix B.

In fact, estimates of returns to schooling increase along the wage distribution. This feature holds for both genders, but it is particularly noticeable in the case of women. Evidence presented in Figure 8 also clarifies that the increase in returns to schooling along 1986-2013 period holds only for individuals placed above the 25th quantile of the wage distribution. Returns estimated at its lower tail using the 2013 QP wave stand below those obtained

5. Significance tests show that this drop, although small, is statistically significant.

using the 1986 data. Moreover, up to 2003, evidence of higher returns for women also holds only above the 1st decile of the distribution.⁶

The comparison between the returns to schooling estimated at different points of the distribution provides a measure of their dispersion. Such comparison, illustrated in Figure 9, shows in the first place that, across the whole distribution, returns are more dispersed among women than among men. In both cases, inequality in returns widened along the 1986-2013 period, but it was particularly noticeable in the case of female employees and in the early 1990s. This evolution seems to be largely driven by developments at the lower part of the conditional wage distribution, since at the upper quantiles inequality in returns has remained relatively stable. Moreover, among high earners variability in the returns is lower than at the lower part of the wage distribution.

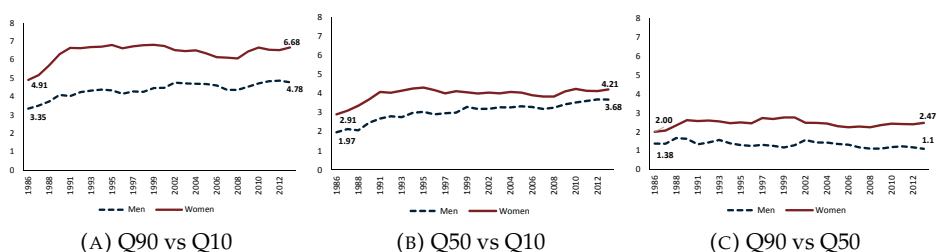


FIGURE 9: Dispersion in returns to schooling (percentage points).

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The charts depict the difference between the returns to schooling estimated for different points of the distribution.

Returns to specific education levels

On the basis of specifications (4) and (5) it is possible to assess the average wage premium associated with specific levels of education. In this case, coefficients β_j , $j = \{2, 3, 4\}$, represent the wage gain from completing schooling level j relative to individuals who have not completed the 9th grade (corresponding to education level $j = 1$, the omitted category). As we are interested in the wage gain relative to the schooling level immediately

6. Note that the differences in the returns estimated on the basis of QR for the 1986 and 2013 waves of QP are found to be statistically significant.

before, we plot in Figure 10 the difference in the coefficients estimated using specification (4) as follows:

$$r_{9th} = \beta_2$$

$$r_{secondary} = \beta_3 - \beta_2$$

$$r_{tertiary} = \beta_4 - \beta_3$$

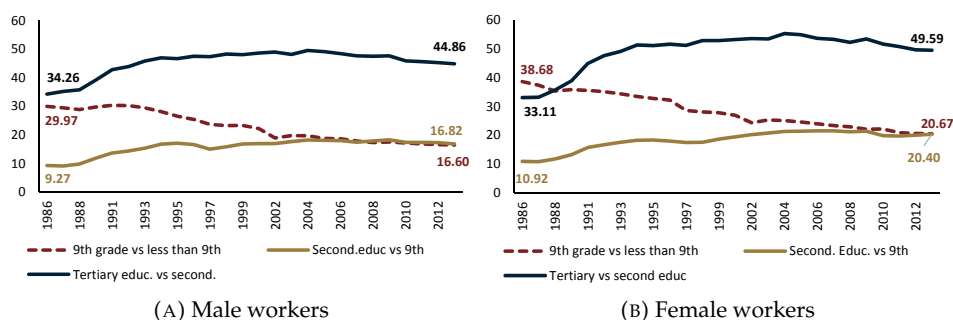


FIGURE 10: Returns to schooling at the mean of the wage distribution by educational attainment (per cent).

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The charts depict the per cent wage premium associated with each educational level relative to level immediately before. For greater detail refer to Appendix B.

In the first place, Figure 10 confirms that women benefit from larger returns to education than men, except as regards the relative premium associated with tertiary education in the first years of the sample. It also shows that the increase over time in overall returns to schooling documented in the previous subsection is largely driven by the evolution of the premium associated with tertiary education. Indeed, in the beginning of the 1986-2013 period, completing an university degree is estimated to yield male workers a 34.3 per cent premium vis-à-vis completing secondary education, while for women such figure stands at 33.1 per cent. Results obtained using the 2013 wave of QP imply that men holding an university degree enjoy a 44.9 per cent wage premium relative to comparable workers who complete only secondary education. For women, such figure stands at 49.6 per cent.

Regarding secondary education, there is evidence that the gain relative to completing only the 9th grade increased along the 1986-2013, but it is still considerably below the one referring to tertiary education: 16.8 per cent for men and 20.4 per cent for women. The increase in the premia estimated for secondary and tertiary education occurred against a background of an expansion in the pool of workers holding these schooling levels, suggesting that it may have been demand-driven.

The premium for tertiary education increased markedly in the first half of the 1990s - especially in the case of women - and then remained stable

up to 2009, when there is a minor decline in its magnitude. Evidence from regressions focusing specifically on university-educated individuals aged between 25 and 45 suggests that the younger workers benefit from lower returns to schooling and have experienced a slightly larger drop in returns than the overall sample (Figure 11).

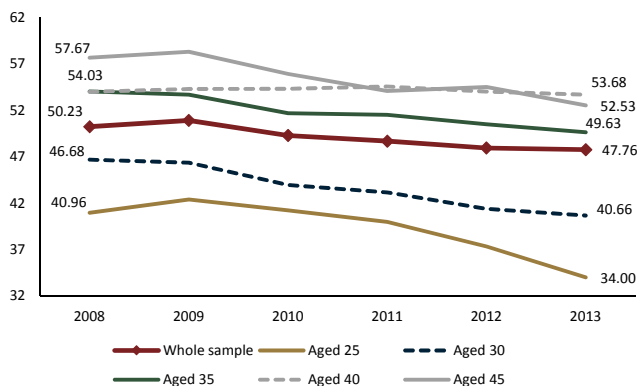


FIGURE 11: Returns to tertiary education in the post-2009 period for younger cohorts (per cent).

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: The chart depicts the relative wage increment from obtaining an university degree vis-à-vis completing only secondary education. Figures are obtained pooling data for both men and women.

Finally, it is worth highlighting the sharp decline in the premium associated with completing the 9th grade. In the late-1980s, it was very close to the relative wage gain enjoyed by individuals holding an university degree (in the case of women it was in fact higher). Since then, our estimates suggest a decline and the magnitude estimated on the basis of 2013 data corresponds to approximately half the figures obtained with the 1986 QP wave. This evolution is in line with the pattern typically found for advanced economies (Montenegro and Patrinos (2014)).

The drop in the returns to the 9th grade has been accompanied by an increase in the share of employees reporting it as the highest level completed (and a sharp decline in those holding less than the 9th grade). This evolution, plotted in Figure 12, may be related to the fact that compulsory schooling was extended to the 9th grade in 1986. The measure applied only to individuals born as of 1980, which would be showing up in QP data as of 1996.

However, even older individuals who were still attaining school in 1986 may have anticipated that the market would start valuing completion of the 9th grade and decided to study longer - either to complete just the 9th grade or further levels as a differentiation factor. This would result in a decline in the

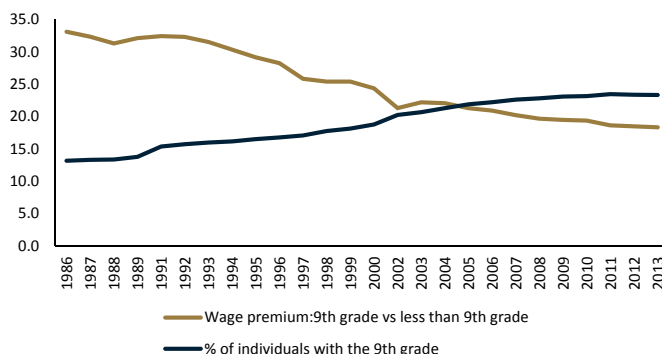
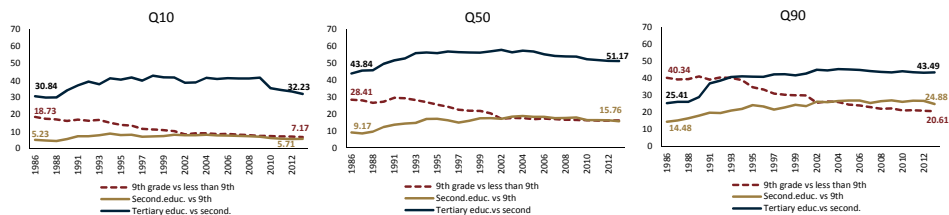


FIGURE 12: Returns to completing the 9th grade *vs* share of individuals with the 9th grade.

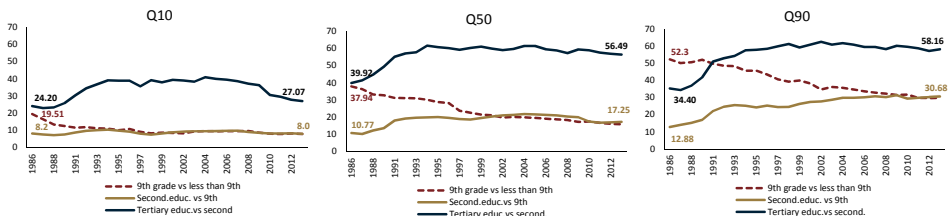
Source: Authors' calculations based on *Quadros de Pessoal*.

share of individuals with less than the 9th grade even before the first cohorts affected by the legal change join the labour force. Although it is not possible to establish a causality link between this decline and the drop in the returns to the 9th grade, *ceteris paribus*, an expansion in the pool of workers who have completed the latter level would in principle result in such an evolution. This suggests that the drop in the premium for completing the 9th grade was supply-driven. Additionally, this evolution may have been reinforced by selection effects. In particular, it is arguable that the individuals who drop-out after completing the 9th grade in the recent period differ from those who did it some years or decades ago, in terms of characteristics that may result in lower returns to schooling (eg, younger 9th grade drop-outs may be expected to have, on average, poorer innate ability endowments, or less favourable family backgrounds).

Evidence obtained from QR estimates of specification (5), presented in Figure 13, shows that, for women, the relative premium from completing tertiary education increases along the wage distribution. In the case of male employees, such evidence holds only below the 9th decile. Figure 13 also suggests that the rise in the relative return to university degrees occurred throughout the wage distribution, but it is more noticeable at the upper quantiles and in the case of women. Regarding the already mentioned drop in the magnitude of the returns as of 2009, it appears to result from developments at the lower tail of the conditional distribution. Finally, regarding the premia estimated for completing secondary education and the 9th grade, their average evolutions are driven by results in the upper quantiles, as below the median of the wage distribution they have remained broadly constant.



(A) Male workers



(B) Female workers

FIGURE 13: Returns to schooling across the wage distribution by educational attainment (per cent).

Source: Authors' calculations based on *Quadros de Pessoal*.

Discussion and concluding remarks

This article provided evidence of the existence of private returns to schooling in the Portuguese economy. The returns increased in the late-1980s and the 1990s, especially as regards tertiary education. This occurred in parallel with an expansion of the pool of workers holding university degrees, suggesting it was surpassed by a rise in the demand for skilled labour. During the late-1990s and early-2000s, the returns remained relatively constant, largely reflecting the stabilization of the wage premium for tertiary education. In the post-2009 period, however, our results point to a minor decrease in the magnitude of the returns, both in the case of tertiary and secondary education.

In spite of these changes along the 1986-2013 period, the overall picture does not change: the returns to schooling are found to be higher in the case of women and to increase along the wage distribution and with educational attainment: formal education appears to be more valued for women and high-pay and high-skill jobs. Several factors may be put forward as possible explanations of the evolution of returns to schooling just described.

A phenomenon of over-education, which we have indeed documented, is the first one: over-education, measured as a non-negligible share of highly-educated workers in blue-collar occupations, results in their placement in the bottom quantiles of the wage distribution. This translates into low returns

to schooling for these individuals, increasing dispersion within the same educational level and contributing to explain the pattern of increasing returns along the distribution.

The effects of over-education may be a reflection of qualitative aspects of schooling: while the estimation of the returns takes only into account the *quantity* of educational attainment, it disregards factors such as school quality or the different valuation attributed to different areas of study. Attending poor-quality schools or investing in a field of study that receives low valuation in the labour market would, in principle, result in low-pay jobs and in positions requiring low skills.

We cannot rule-out the possibility that the developments hereby described are affected by the fact that individual differences in ability (or other unobserved attributes) are not being controlled for. In particular, it is expectable that differences in individual ability play a bigger role in explaining the dispersion in returns among more skilled workers. For low-educated individuals, by contrast, differences shall be relatively washed-down. Not controlling for these differences would result in an overestimation of returns to schooling in the upper quantiles of the distribution and reinforce the effects of over-education and low school quality.

In spite of focusing only on the *private* returns to education, our results unveil important messages for individuals and policymakers alike: in Portugal, education remains a profitable investment for individual agents and policymakers must take this into account when designing policies and incentive schemes.

The returns are highest for tertiary education and it is likely that individuals will continue to invest in education and, in particular, in university degrees. Compulsory schooling has recently been extended to 12 years, encompassing secondary education. This may also provide incentives for individuals that would otherwise leave school to go further and complete a tertiary educational level to differentiate themselves from the holders of secondary schooling. These factors would in principle result in the expansion of the student population in the next decades but are likely to be counteracted by demographic trends.

Against a background of tight budgetary constraints, the challenge for policymakers relies in ensuring the quality of public school system while providing low-income households conditions to access tertiary level education. Moreover, this cannot be done at the expense of low quality pre-school or elementary education, as investments in lower schooling levels increase the returns to subsequent ones⁷. These tensions may require a reshuffling in terms of the funding sources of public expenditure on education policy. A common suggestion relies on increasing the share of costs supported

7. Refer to Heckman and Cunha (2007).

by the individuals in tertiary education. This line of reasoning is based on the idea that *social* returns to schooling are relatively lower for tertiary levels, whereas *private* returns are high - an evidence supported by our empirical findings. Examples of measures aimed at increasing individual participation in financing include mere increases in tuition or the recently higher education reform in the UK encompassing the setting-up of loan scheme that is contingent on graduates' future earnings. Resorting to this sort of measures may create additional leeway to reinforce support to low-income households.

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Appendix A: Details on the empirical distribution of the logarithm of hourly wages

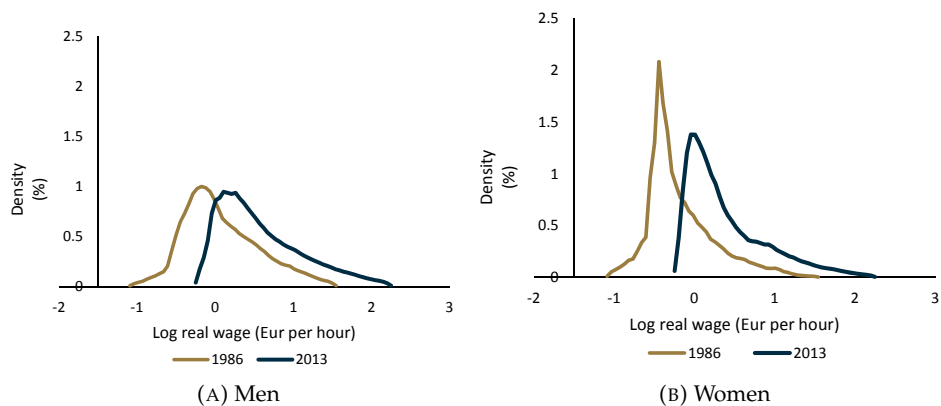


FIGURE A.1: Distribution of real wages in 1986 and 2013.

Source: Authors' calculations based on *Quadros de Pessoal*.

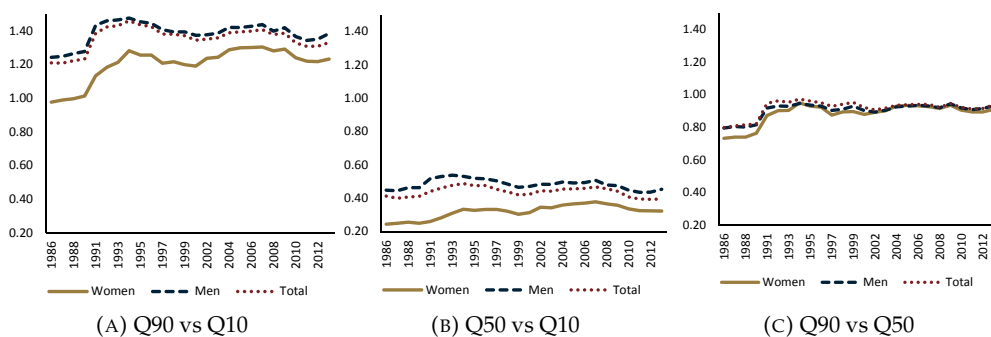
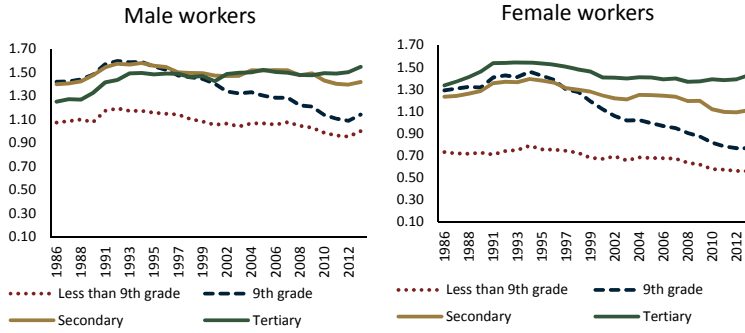


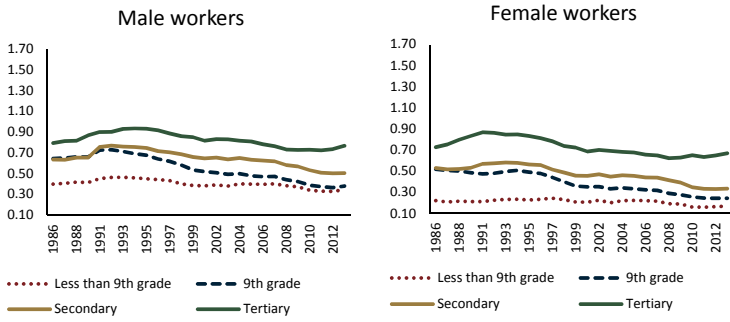
FIGURE A.2: Real wage dispersion.

Source: Authors' calculations based on *Quadros de Pessoal*.

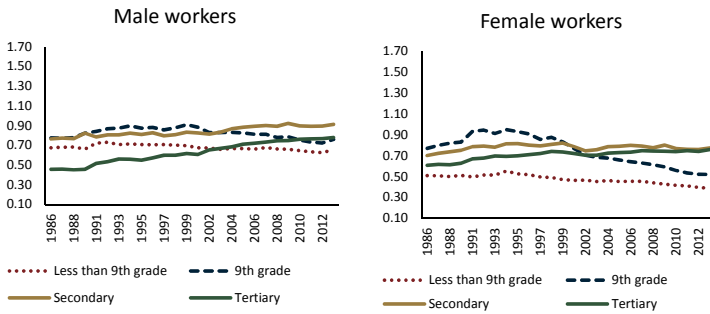
Note: Charts depict the difference in the logarithm of real hourly wages at different points of the distribution.



(A) Q90 vs Q10



(B) Q50 vs Q10



(C) Q90 vs Q50

FIGURE A.3: Real wage dispersion by schooling attainment.

Source: Authors' calculations based on *Quadros de Pessoal*.

Note: Charts depict the difference in the logarithm of real hourly wages at different points of the distribution.

Appendix B: Estimation results

	Baseline specification (2)					Specification (2) with additional controls				
	1986	1996	2005	2010	2013	1986	1996	2005	2010	2013
Educ. (years)	0.0608*** (0.0001)	0.0719*** (0.0001)	0.0730*** (0.0001)	0.0715*** (0.0001)	0.0706*** (0.0001)	0.0478*** (0.0001)	0.0580*** (0.0001)	0.0640*** (0.0001)	0.0638*** (0.0001)	0.0637*** (0.0001)
Age	0.0580*** (0.0002)	0.0458*** (0.0002)	0.0413*** (0.0002)	0.0404*** (0.0002)	0.0423*** (0.0002)	0.0489*** (0.0002)	0.0406*** (0.0002)	0.0382*** (0.0002)	0.0363*** (0.0002)	0.0374*** (0.0002)
Age sqrd.	-0.0006*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0005*** (0.0000)	-0.0004*** (0.0000)	-0.0003*** (0.0000)	-0.0003*** (0.0000)	-0.0003*** (0.0000)
Sex (male=1)	0.1858*** (0.0007)	0.2526*** (0.0006)	0.2567*** (0.0006)	0.2449*** (0.0005)	0.2477*** (0.0006)	0.1530*** (0.0007)	0.1957*** (0.0007)	0.2154*** (0.0006)	0.2036*** (0.0006)	0.1992*** (0.0006)
Firm size (log)	0.0778*** (0.0002)	0.0832*** (0.0002)	0.0715*** (0.0001)	0.0497*** (0.0001)	0.0446*** (0.0001)	0.0590*** (0.0002)	0.0621*** (0.0002)	0.0539*** (0.0001)	0.0335*** (0.0001)	0.0303*** (0.0001)
Tenure	0.0078*** (0.0001)	0.0117*** (0.0001)	0.0166*** (0.0001)	0.0165*** (0.0001)	0.0164*** (0.0001)	0.0097*** (0.0001)	0.0122*** (0.0001)	0.0174*** (0.0001)	0.0162*** (0.0001)	0.0159*** (0.0001)
Tenure sqrd.	-0.0000*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	-0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0003*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)
Intercept	-2.1323*** (0.0033)	-1.9018*** (0.0035)	-1.7187*** (0.0034)	-1.6105*** (0.0035)	-1.7051*** (0.0039)	-2.0086*** (0.0042)	-1.6877*** (0.0043)	-1.5613*** (0.0043)	-1.4275*** (0.0043)	-1.5136*** (0.0046)
Region controls	N	N	N	N	N	Y	Y	Y	Y	Y
Industry controls	N	N	N	N	N	Y	Y	Y	Y	Y
R-squared	0.54	0.53	0.49	0.45	0.45	0.63	0.59	0.53	0.50	0.49
N	990215	1464732	1981128	2042134	1904805	990215	1464732	1981128	2042134	1904805

TABLE B.1. Wage regressions - OLS.

Source: Authors' calculations based on *Quadros de Pessoal*.

Notes: Coefficients obtained from OLS regressions using specification (2) pooling data for men and women. Standard errors in parentheses.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

	Baseline specification (4)					Specification (4) with additional controls				
	1986	1996	2005	2010	2013	1986	1996	2005	2010	2013
9th grade	0.3309*** (0.0011)	0.2824*** (0.0009)	0.2127*** (0.0007)	0.1936*** (0.0007)	0.1832*** (0.0008)	0.2360*** (0.0009)	0.1966*** (0.0008)	0.1663*** (0.0007)	0.1577*** (0.0007)	0.1508*** (0.0008)
Second. educ.	0.4365*** (0.0016)	0.4604*** (0.0010)	0.4138*** (0.0008)	0.3839*** (0.0008)	0.3720*** (0.0008)	0.3156*** (0.0012)	0.3304*** (0.0009)	0.3278*** (0.0008)	0.3117*** (0.0008)	0.3058*** (0.0008)
Tertiary educ.	0.7643*** (0.0026)	0.9508*** (0.0017)	0.9356*** (0.0011)	0.8767*** (0.0010)	0.8496*** (0.0010)	0.6521*** (0.0019)	0.8005*** (0.0014)	0.8321*** (0.0010)	0.7883*** (0.0009)	0.7697*** (0.0009)
Age	0.0569*** (0.0002)	0.0437*** (0.0002)	0.0397*** (0.0002)	0.0401*** (0.0002)	0.0425*** (0.0002)	0.0471*** (0.0002)	0.0381*** (0.0002)	0.0360*** (0.0002)	0.0356*** (0.0002)	0.0370*** (0.0002)
Age sqrd.	-0.0006*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0005*** (0.0000)	-0.0004*** (0.0000)	-0.0004*** (0.0000)	-0.0003*** (0.0000)	-0.0004*** (0.0000)
Sex (male=1)	0.1947*** (0.0007)	0.2543*** (0.0006)	0.2581*** (0.0005)	0.2484*** (0.0005)	0.2525*** (0.0006)	0.1606*** (0.0007)	0.1964*** (0.0007)	0.2167*** (0.0006)	0.2068*** (0.0006)	0.2030*** (0.0006)
Firm size (log)	0.0784*** (0.0002)	0.0841*** (0.0002)	0.0713*** (0.0001)	0.0500*** (0.0001)	0.0450*** (0.0001)	0.0584*** (0.0002)	0.0620*** (0.0002)	0.0532*** (0.0001)	0.0333*** (0.0001)	0.0303*** (0.0001)
Tenure	0.0068*** (0.0001)	0.0123*** (0.0001)	0.0179*** (0.0001)	0.0173*** (0.0001)	0.0169*** (0.0001)	0.0087*** (0.0001)	0.0127*** (0.0001)	0.0187*** (0.0001)	0.0170*** (0.0001)	0.0164*** (0.0001)
Tenure sqrd.	-0.0000 (0.0000)	-0.0002*** (0.0000)	-0.0003*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	-0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0003*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)
Intercept	-1.8229*** (0.0033)	-1.4761*** (0.0035)	-1.2544*** (0.0034)	-1.1626*** (0.0035)	-1.2639*** (0.0039)	-1.7812*** (0.0042)	-1.3597*** (0.0043)	-1.1697*** (0.0042)	-1.0422*** (0.0043)	-1.1261*** (0.0046)
Region controls	N	N	N	N	N	Y	Y	Y	Y	Y
Industry controls	N	N	N	N	N	Y	Y	Y	Y	Y
R-squared	0.52	0.52	0.49	0.46	0.46	0.61	0.59	0.54	0.51	0.50
N	990215	1464732	1981128	2042134	1904805	990215	1464732	1981128	2042134	1904805

TABLE B.2. Wage regressions - OLS.

Source: Authors' calculations based on *Quadros de Pessoal*.

Notes: Coefficients obtained from OLS regressions using specification (4) pooling data for men and women. Standard errors in parentheses.

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

	1986			1996			2005			2010			2013		
	P10	P50	P90	P10	P50	P90	P10	P50	P90	P10	P50	P90	P10	P50	P90
Educ. (years)	0.0385*** (0.0001)	0.0611*** (0.0001)	0.0762*** (0.0002)	0.0394*** (0.0001)	0.0719*** (0.0001)	0.0887*** (0.0002)	0.0371*** (0.0001)	0.0732*** (0.0001)	0.0902*** (0.0002)	0.0326*** (0.0001)	0.0716*** (0.0001)	0.0885*** (0.0002)	0.0309*** (0.0001)	0.0709*** (0.0001)	0.0881*** (0.0002)
Age	0.0503*** (0.0002)	0.0530*** (0.0002)	0.0705*** (0.0004)	0.0235*** (0.0001)	0.0405*** (0.0002)	0.0655*** (0.0004)	0.0188*** (0.0001)	0.0386*** (0.0002)	0.0599*** (0.0004)	0.0147*** (0.0001)	0.0368*** (0.0002)	0.0600*** (0.0004)	0.0135*** (0.0001)	0.0387*** (0.0002)	0.0646*** (0.0004)
Age sqrd	-0.0006*** (0.0000)	-0.0006*** (0.0000)	-0.0007*** (0.0000)	-0.0002*** (0.0000)	-0.0004*** (0.0000)	-0.0006*** (0.0000)	-0.0002*** (0.0000)	-0.0003*** (0.0000)	-0.0005*** (0.0000)	-0.0001*** (0.0000)	-0.0003*** (0.0000)	-0.0005*** (0.0000)	-0.0001*** (0.0000)	-0.0003*** (0.0000)	-0.0006*** (0.0000)
Sex (male=1)	0.1522*** (0.0008)	0.1906*** (0.0008)	0.2174*** (0.0015)	0.1710*** (0.0005)	0.2575*** (0.0007)	0.3158*** (0.0014)	0.1533*** (0.0004)	0.2568*** (0.0006)	0.3393*** (0.0012)	0.1244*** (0.0003)	0.2407*** (0.0006)	0.3268*** (0.0012)	0.1206*** (0.0003)	0.2391*** (0.0006)	0.3379*** (0.0012)
Firm size (log)	0.0667*** (0.0002)	0.0766*** (0.0002)	0.0813*** (0.0004)	0.0711*** (0.0001)	0.0829*** (0.0002)	0.0801*** (0.0003)	0.0537*** (0.0001)	0.0716*** (0.0002)	0.0735*** (0.0003)	0.0287*** (0.0001)	0.0488*** (0.0001)	0.0529*** (0.0003)	0.0257*** (0.0001)	0.0429*** (0.0001)	0.0506*** (0.0003)
Tenure	0.0073*** (0.0001)	0.0074*** (0.0001)	0.0084*** (0.0003)	0.0116*** (0.0001)	0.0136*** (0.0001)	0.0038*** (0.0003)	0.0129*** (0.0001)	0.0167*** (0.0001)	0.0106*** (0.0002)	0.0112*** (0.0001)	0.0167*** (0.0001)	0.0112*** (0.0002)	0.0128*** (0.0001)	0.0170*** (0.0001)	0.0088*** (0.0002)
Tenure sqrd	-0.0001*** (0.0000)	-0.0000*** (0.0000)	-0.0000 (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0000*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	0.0001*** (0.0000)
Intercept	-2.1000*** (0.0041)	-2.0584*** (0.0035)	-2.1252*** (0.0066)	-1.4851*** (0.0019)	-1.8398*** (0.0035)	-1.9750*** (0.0073)	-1.1600*** (0.0016)	-1.6994*** (0.0035)	-1.8373*** (0.0073)	-0.8478*** (0.0015)	-1.5666*** (0.0032)	-1.7899*** (0.0077)	-0.8485*** (0.0015)	-1.6566*** (0.0033)	-1.9629*** (0.0085)
N	990215	990215	990215	1464732	1464732	1464732	1981128	1981128	1981128	2042134	2042134	2042134	1904805	1904805	1904805

TABLE B.3. Wage regressions - Quantile regressions.

Source: Authors' calculations based on *Quadros de Pessoal*.

Notes: Coefficients obtained from QR using specification (3). Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

	1986			1996			2005			2010			2013		
	P10	P50	P90	P10	P50	P90	P10	P50	P90	P10	P50	P90	P10	P50	P90
9th grade	0.1995*** (0.0013)	0.3219*** (0.0014)	0.4409*** (0.0022)	0.1359*** (0.0009)	0.2636*** (0.0011)	0.3751*** (0.0019)	0.0996*** (0.0006)	0.1904*** (0.0008)	0.2861*** (0.0016)	0.0862*** (0.0004)	0.1747*** (0.0007)	0.2580*** (0.0017)	0.0813*** (0.0004)	0.1653*** (0.0007)	0.2534*** (0.0018)
Second. educ.	0.2628*** (0.0017)	0.4312*** (0.0020)	0.5866*** (0.0027)	0.2364*** (0.0011)	0.4497*** (0.0012)	0.6192*** (0.0021)	0.1955*** (0.0007)	0.3962*** (0.0009)	0.5699*** (0.0017)	0.1635*** (0.0005)	0.3548*** (0.0008)	0.5352*** (0.0018)	0.1530*** (0.0005)	0.3418*** (0.0008)	0.5285*** (0.0018)
Tertiary educ.	0.5473*** (0.0054)	0.8431*** (0.0029)	0.8663*** (0.0035)	0.6331*** (0.0038)	1.0206*** (0.0021)	1.1082*** (0.0024)	0.6006*** (0.0020)	0.9837*** (0.0013)	1.1092*** (0.0018)	0.4940*** (0.0016)	0.9108*** (0.0011)	1.0676*** (0.0018)	0.4502*** (0.0014)	0.8800*** (0.0010)	1.0503*** (0.0019)
Age	0.0499*** (0.0002)	0.0508*** (0.0002)	0.0709*** (0.0004)	0.0221*** (0.0001)	0.0371*** (0.0002)	0.0626*** (0.0004)	0.0174*** (0.0001)	0.0347*** (0.0002)	0.0573*** (0.0004)	0.0141*** (0.0001)	0.0343*** (0.0002)	0.0582*** (0.0004)	0.0139*** (0.0001)	0.0366*** (0.0002)	0.0632*** (0.0004)
Age sqrd.	-0.0006*** (0.0000)	-0.0006*** (0.0000)	-0.0008*** (0.0000)	-0.0002*** (0.0000)	-0.0004*** (0.0000)	-0.0006*** (0.0000)	-0.0002*** (0.0000)	-0.0003*** (0.0000)	-0.0005*** (0.0000)	-0.0001*** (0.0000)	-0.0003*** (0.0000)	-0.0005*** (0.0000)	-0.0001*** (0.0000)	-0.0004*** (0.0000)	-0.0006*** (0.0000)
Sex (male=1)	0.1610*** (0.0008)	0.1981*** (0.0008)	0.2189*** (0.0015)	0.1697*** (0.0005)	0.2576*** (0.0007)	0.3154*** (0.0014)	0.1517*** (0.0005)	0.2550*** (0.0006)	0.3411*** (0.0012)	0.1264*** (0.0004)	0.2415*** (0.0006)	0.3327*** (0.0012)	0.1248*** (0.0004)	0.2418*** (0.0006)	0.3459*** (0.0012)
Firm size (log.)	0.0653*** (0.0002)	0.0768*** (0.0002)	0.0811*** (0.0004)	0.0697*** (0.0001)	0.0834*** (0.0002)	0.0818*** (0.0003)	0.0525*** (0.0001)	0.0699*** (0.0002)	0.0749*** (0.0003)	0.0279*** (0.0001)	0.0467*** (0.0001)	0.0544*** (0.0003)	0.0255*** (0.0001)	0.0403*** (0.0001)	0.0515*** (0.0003)
Tenure	0.0067*** (0.0002)	0.0062*** (0.0001)	0.0076*** (0.0003)	0.0112*** (0.0001)	0.0145*** (0.0001)	0.0047*** (0.0003)	0.0132*** (0.0001)	0.0182*** (0.0001)	0.0121*** (0.0002)	0.0116*** (0.0001)	0.0173*** (0.0001)	0.0122*** (0.0002)	0.0136*** (0.0001)	0.0166*** (0.0001)	0.0094*** (0.0002)
Tenure sqrd	-0.0001*** (0.0000)	0.0000*** (0.0000)	0.0000*** (0.0000)	-0.0003*** (0.0000)	-0.0002*** (0.0000)	0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	-0.0002*** (0.0000)	-0.0002*** (0.0000)	-0.0000*** (0.0000)	-0.0002*** (0.0000)	-0.0001*** (0.0000)	0.0001*** (0.0000)
Intercept	-1.8962*** (0.0039)	-1.7218*** (0.0034)	-1.7707*** (0.0067)	-1.2250*** (0.0018)	-1.3824*** (0.0035)	-1.4688*** (0.0071)	-0.8965*** (0.0021)	-1.1694*** (0.0034)	-1.2918*** (0.0076)	-0.6226*** (0.0016)	-1.0440*** (0.0028)	-1.2513*** (0.0079)	-0.6578*** (0.0018)	-1.1268*** (0.0030)	-1.4207*** (0.0085)
N	990215	990215	990215	1464732	1464732	1464732	1981128	1981128	1981128	2042134	2042134	2042134	1904805	1904805	1904805

TABLE B.4. Wage regressions - Quantile regressions.

Source: Authors' calculations based on *Quadros de Pessoal*.Notes: Coefficients obtained from QR using specification (5). Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$