What Education Pays? Non-Linear Returns to Schooling Among Filipino Men

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Abstract

The issue of possible non-linearities in the relationship between log wages and schooling has received a good deal of attention in the literature on the United States, as well as in Less Developed Countries (LDCs). In this paper, I use data from a recent household survey for the Philippines, the 1998 Annual Poverty Indicator Survey (APIS), to test the fit of the log-linear specification for Filipino men. I present results based on a number of estimation strategies, including discontinuous spline regressions, semi-parametric regressions with a large number of dummies for years of schooling and experience, and kernel regressions. The basic conclusions of the paper are two. First, there appear to be large differences between the rates of return to education across levels in the Philippines. In particular, the wage premia to both primary and secondary education are lower than those for tertiary education. Second, within a given level, the last year of schooling is disproportionately rewarded in terms of higher wages. That is, there are clear sheepskin effects associated with graduation from primary school, secondary school, and university.

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1. Introduction

The relationship between education and wages is one of the most studied topics in labor economics. Since the 1970's, researchers have found a strong, positive relationship between educational attainment and wages in country after country. Most of this research has taken Mincer's (1974) formulation as a starting point. This formulation states that the log of the observed wage rate is an additive function of a linear term in years of completed schooling and a quadratic term in experience. The coefficient on the years of completed schooling is frequently referred to as the "rate of return" to education (for example Card 1999; Willis 1986).[2]

The log-linear relationship between wages and years of completed education has an obvious implication: Controlling for experience, every year of education, no matter at what level, adds to log earnings by the same amount. For analytical purposes, it is convenient to break down this assumption into two components. First, the log-linear specification assumes that the rate of return to education is the same across schooling levels. This assumption will not hold if the mean rate of return is higher for some levels than others—say, secondary school compared to primary school—or for some kinds of schooling—say, university education compared to post-secondary vocational training. Such differences could be driven by a multitude of factors, including differences in the quality of education across levels, changes in the supply of and demand for graduates which have an effect on the price which these graduates can command in the labor market, serious barriers to entry at some levels, or a pattern of development which rewards some skills more highly than others. Second, the log-linear specification assumes that the rate of return to an additional year of schooling is no higher for the last year of schooling in a particular level—primary, secondary, university—than in any other intervening year. This may not hold if credentials themselves are rewarded, a hypothesis known as the "sheepskin effect" in the literature.

The assumption of linearity in the relationship between log wages and schooling has received much attention in the United States, where the issue is still not resolved (see, for example Card 1999; Heckman, Layne-Farrar, and Todd 1996; Hungerford and Solon 1987; Park 1994). There are also literally hundreds of empirical studies on the relationship between wages and education in Less Developed Countries (LDCs) (for reviews see Psacharopoulos 1973, 1981,

[2] Strictly speaking, this is only the case if a number of conditions are met (for example, if the direct private costs of education are zero, and if there is no grade repetition). These conditions are unlikely to hold perfectly in practice. In this paper I refer to the coefficient on education in the Mincerian regression as the "rate of return" for convenience, even though it would be more accurate to describe it as the "wage premium" or "wage increment".

1985, 1989 and 1994). Many of these studies have attempted to identify differences in the rate of return across levels, but the data and methods used are often poor at best (see, for example, Bennell 1996 for a critique). Two studies which are methodologically similar to the analysis conducted in this paper focus on Brazil (Lam and Schoeni 1993; Strauss and Thomas 1996). Both find convexity in the relationship between market wages and education, and significant, large sheepskin effects associated with the completion of particular years of schooling.

In this paper, I use data from a recent household survey conducted in the Philippines to test the fit of the log-linear relationship between wages and schooling. Like many middle-income countries, the Philippines has a large number of people in the labor force who have received very different amounts of schooling—ranging from no formal education whatsoever, to completed university education. This allows me to inspect the relationship between log wages and education over a broader range than would be the case in the United States. The Philippines has also seen important economic changes in the last two decades, including a supply-push to make secondary education universal, volatility in various economic indicators, including aggregate GDP growth, and substantial differences in the rate of growth of output and productivity across sectors. These changes, many of which are not yet complete, may result in differences in the rate of return to education across levels and across years which are large, at least in the short run.

The basic conclusions of the paper are two. First, there appear to be large differences between the rates of return to education across levels in the Philippines. In particular, the wage premia to both primary and secondary education are lower than those for university education. Some of the convexity in the relationship between the wage rate and schooling may be caused by the omission of variables which control for ability and motivation, which induce self-selection into higher levels of schooling. I argue, however, that much of the difference in the rates of return is probably driven by characteristics of the Philippine education system and labor market. Second, within a given level, the last year of schooling is disproportionately rewarded in terms of higher wages, so that there are clear sheepskin effects associated with graduation from primary school, secondary school, and university. Here, too, self-selection probably plays a role, since more motivated students are less likely to drop out before completing a given level of schooling. But much of the sheepskin effect could also be caused by the lack of accurate information which employers have about the underlying productivity of potential employees, or by requirements of particular diplomas for access to some jobs.

The rest of the paper proceeds as follows. Section 2 gives a brief review of education and labor markets in the Philippines, including a discussion of recent literature. Section 3 describes the data set. In Section 4, I discuss the econometric specification. Section 5 presents results on

the differences in rates of return across levels, and across years within levels. Section 6 presents some very simple simulations about the effects that large differences in the rates of return of primary school, secondary school, and university graduates might have on the lifetime stream of expected earnings. Section 7 concludes with a general discussion and suggestions for further research.

2. Education in the Philippines

Basic education in the Philippines is acquired in two levels: primary education (six years), and secondary education (four years). Graduates from secondary school can continue on to university (five years) or post-secondary vocational training (up to four years).[3] As in most countries, there are both public and private providers of education, and the latter are particularly important at higher levels of education: Private providers of education account for a very small fraction of total enrollment in primary school (7%), but much larger fractions for secondary school (22%), and a majority of tertiary education (77%). Total expenditures on education in the Philippines stood at 6.9% of GDP in 1997—roughly comparable to the amount spent by other countries in a similar income bracket. The fraction of total expenditures on education which is made by private households in the Philippines (43%) is high relative to other countries.

The Philippines has achieved impressively high enrollment rates in primary and secondary education for its income level. Administrative data indicate that the gross enrollment rate in secondary school—the fraction of children aged 13-16 enrolled in school—was 75% in the Philippines in 1997 (Asian Development Bank and World Bank 1999, p. 158). But this high aggregate figure for the Philippines masks important differences by region: The gross enrollment ratio in secondary school in the Autonomous Region of Muslim Mindanao, for example, was only 29.7%. There are also differences by gender: Analysis of the Annual Poverty Indicator Survey (APIS) shows that in rural areas 84.1% of girls aged 13-16, but only 71.1% of boys were attending school in 1998; in urban areas, comparable figures were 90.4% for girls and 84.7% for boys.[4]

The low quality of basic education is an oft-voiced concern in the Philippines (Asian Development Bank and World Bank 2000). Poorly prepared and unmotivated teachers, inadequate teaching materials, and decaying infrastructure are all cited as problems. So is a

[3] Much of this section is based on Asian Development Bank and World Bank (1999).
[4] Differences by gender and region are significant at the 0.1% level or better. Similar figures, based on the 1998 Philippines Demographic and Health Survey, can be found in Filmer (1999).

misallocation of resources in the public system, with a very large percentage of total expenditures going to teacher salaries (87.7% in 1999, up from 74.3% in 1990), to the detriment of expenditures on maintenance and other operating expenditures (8.8% in 1999, down from 16.5% in 1990) (Asian Development Bank and World Bank 1999, 112). Filipino students perform poorly on international tests, and poorly on tests designed by the Department of Education, Culture and Sports which are intended to measure minimum acceptable levels of achievement. The quality of education provided by the private sector appears to be noticeably higher (Jimenez et. al. 1995).

The low quality of education may be one of the factors contributing to the poor economic performance of the Philippines. Since 1984, real GDP per capita has essentially been stagnant: The total change in real GDP per capita between 1984 and 1999 was 2.6%, a very small amount indeed in comparison with many of the Philippines' East Asian neighbors, which posted *annual* growth rates above 5% for much of the period. Productivity in the Philippines declined for the economy as a whole (by about 0.9% over the fifteen-year period), and fell dramatically in industry (declining by 17.0% during the period).

There are numerous studies about the relationship between schooling and labor market performance in the Philippines (for example, Tan and Paqueo 1989; Hossain and Psacharopolous 1994; Lanzona 1997; Maluccio 1998). Most of the studies which use national-level data are by now outdated. The Lanzona (1997) and Maluccio (1998) papers are based on a panel of households interviewed in 1978, 1983, and 1994. Because the households selected are all in a single province of the 78 which exist in the Philippines (the province of Camarines Sur, in the Bicol region), it is not clear whether the results in these papers are applicable to the country at large.

3. The data set

This paper is largely based on the 1998 Annual Poverty Indicator Survey (APIS). The APIS covered 38,710 households, including 195,791 individuals. It is the first nationwide multipurpose survey conducted in the Philippines, and includes a basic consumption and income module, as well as modules addressing the educational attainment, health status, and degree of contraceptive use by various household members. Exceptionally, the 1998 APIS also included a module intended to capture the household-level effects of the dual financial and El Nino crises in the Philippines in 1998 (see Datt and Hoogeveen 2000).

About 60% of the households in the 1998 APIS are part of a panel with a survey carried out in 1997, the Family Income and Expenditure Survey (FIES). The FIES, which has been conducted in three-year intervals since 1985, has important advantages over the APIS—most notably, the fact that the consumption and income modules are much more extensive. But the FIES is not truly a multi-purpose survey. It includes information on only very basic household demographics and, for example, no information on the educational attainment of household members other than the household head. The large difference in the length of the consumption and income modules in the APIS and the FIES means that it is not possible to compare changes in income or consumption in any straightforward way.

The wage module in the APIS is quite detailed. It asks all household members aged 5 and older about salaries and wages earned in two three-month periods prior to the survey (April 1 to June 30, and July 1 to September 30). Data is gathered separately for the "primary" job and "others". The value of "basic salaries and wages, in cash or in kind" is recorded separately from the value of "allowances, honoraria, tips, housing, clothing, food, etc., in cash or in kind" (separately for both jobs, where applicable). In addition, the survey asks about the number of full days worked, the average number of working hours in these full days, the number of less-than-full days worked, and the average number of hours worked in these less-than-full days (again, separately for both jobs, where applicable). Information on the number of hours and days is collected separately for the months of July, August, and September (but no information on days or hours is collected for April through June). Like most surveys of its kind, the APIS also includes questions about the sector of employment, business classification, and class of worker for those employed. Finally, all adults who have not held a job at all in the six months prior to the survey are asked if they have been looking for work and, if they have not, why.

I construct mean hourly wages for the three-month period prior to the survey by adding the "wage" and "allowance" components for the main job, and dividing this by the total number of hours worked. I consider only men working in the wage-earning sector, rather than all employed men.[5] This is because very few self-employed men report "wages", and it is not clear what to make of the wage data for those who report being self-employed and earning wages on the same job. For similar reasons, the analysis in this paper is limited to the main job, because some men whose main job is in the wage-earning sector have a second job where they are self-employed.

[5] The APIS gives seven categories for class of worker: (i) worked for private household; (ii) worked for private establishment; (iii) worked for government / government corporation; (iv) self employed without any employee; (v) employer in own family operated farm or business; (vi) worked with pay on own family-operated farm or business; (vii) worked without pay on own family-operated farm or business. The analysis in this paper is restricted to categories (i), (ii), (iii), and (vi) above.

Note, however, that only a tiny fraction (0.34%) of men in the sample who hold a wage-earning main job report holding a second wage-earning job, so restricting the analysis to the main job is unlikely to make much of a difference.

There are 33,110 individuals (21,882 males and 11,228 females) who report all of the necessary information for the construction of hourly wages, including data on wages, allowances, days and hours worked on the main job. I deflate all wages with provincial-level price deflators—one for each of the 73 provinces in the country, as well as, separately, the 11 biggest cities (Balisacan 2000). The education module in the APIS is also reasonably detailed. For the purposes of this paper, the relevant question asks about the highest grade completed by all household members aged 6 and older. With the exception of those currently enrolled in school, the APIS provides no information about the kind of school attended by current workers, including whether it was public or private.

4. Estimation strategy

To minimize two econometric problems, the analysis in this paper is limited to men only. First, it is likely that non-random selection into the labor market is more of a concern for women than for men.[6] Note, however, that selection into the wage-earning sector may still be an issue for men in the Philippines, and this could bias the results (for a general discussion of these issues see Strauss and Thomas 1995). Second, like most household surveys, the APIS did not collect information on labor market experience, so I approximate actual experience with "potential experience", defined as X A – S – 6, where A stands for age in years, and S is the years of completed schooling. This formulation, which is standard in the literature, assumes that all children start schooling at age 6 and complete S years of schooling in exactly S years, and begin accumulating labor market experience as soon as they leave school. The use of "potential experience" rather than actual experience may introduce biases in the estimation of the effects of experience on earnings—for example, if children start school late, or if repetition rates are anything other than zero. "Potential experience" will then tend to overstate actual experience,

[6] I do not use the correction for selection proposed by Heckman (1976) because of the absence of a variable which affects the probability of participation in the labor market but does not have an independent effect on wages. One oft-used candidate, namely the amount of non-wage income available to members of a household, including public transfers, pensions, interest and dividends, is not a significant predictor of participation in the labor market conditional on age and education in the APIS sample. Moreover, as is well known, even where such an identifying variable can be found, the corrected Heckman estimates are based on restrictive assumptions about the distribution of the error terms—normally, homoskedasticity and joint normality in the residuals from the switching and wage regressions.

and the coefficient on experience may be a (downward) biased estimate of the returns to actual experience. More worrying is the fact that "potential experience" could seriously over-estimate the number of years of experience of women if women who are currently in the labor market have taken time out at an earlier point in their career to raise children (for example Altonji and Blank 1999). Finally, because of the very small number of observations, I exclude from the sample men whose highest degree is post-secondary vocational training or post-college education (Master's or Ph.D.).

To estimate the relationship between log wages, schooling and experience, I start with the standard Mincerian log-linear specification:

$$(1) \qquad \log W_i = \alpha + \beta S_i + \chi X_i + \delta X^2_i + \varepsilon_i,$$

where S is the years of completed schooling, X is the number of years of potential experience, ε is an error term, and the subscript i refers to individuals.

One concern with this bare-bones specification is the effect of unobserved ability. If ability affects wages and is positively correlated with schooling, as is generally thought to be the case, the estimated β will be biased up (for a discussion of these issues in the United States see, for example, Card 1999; Griliches 1977; for Less Developed Countries, see Schultz 1988). I attempt to correct for this problem by presenting additional results which include the education of the parents as controls (see Heckman and Hotz 1986, and Lam and Schoeni 1993 for a similar approach; a general discussion can also be found in Card 1999, and the references therein). Controls for parental education also seek to capture the effect of investments made by parents in their children that are not captured by years of completed schooling, but which may be positively correlated with schooling. These could include the quality of the schools attended, expenditures on other education inputs, and the time spent by parents teaching their children. Note, however, that measurement error in years of schooling could make estimates which control for parental education more (not less) unreliable than the bare-bones Mincerian estimates. This is because estimates that introduce controls which are correlated with own schooling, including measures of parental education, tend to reduce the signal-to-noise ratio, and increase the downward bias in the coefficient on own schooling which results from measurement error (Griliches 1977; Lam and Schoeni 1993).

The standard Mincerian framework constrains the coefficient on education to be the same across education levels, and across years within a given level. To test the accuracy of this assumption, I compare the results from the Mincerian specification with estimates based on spline

functions and semi-parametric regressions. A discontinuous spline function, proposed by Hungerford and Solon (1987), and which has since formed the basis for much of the literature on "sheepskin effects" in the United States, can be specified as follows:

(2) $\log W_i = \alpha + \chi X_i + \delta X^2_i + \beta S_i + \varsigma D6_i + \varepsilon[(S_i-6)*D6_i] + \tau D10_i + \lambda[(S_i-10)*D10_i] + \varphi D15_i + \iota_{i,}$

where S and X are the years of completed schooling and potential experience, as before, D6, D10, and D15 are dummy variables for those who have completed *at least* six, ten, and fifteen years of schooling, respectively, $[(S_i-6)*D6_i]$ is an interaction term between the D6 dummy and S-6, $[(S_i-10)*D10_i]$ is an interaction term between the D10 dummy and S-10, and ι_i is the error term in the regression. In this specification, the mean rate of return to the first five years of education is given by the parameter β; the rate of return to the sixth year of education is given by the sum of the parameters β and ς; the mean rate of return to the first three years of secondary schooling is given by the sum of the parameters β and ε; the rate of return to the last year of secondary school is given by the sum of the parameters β, ε, and τ; the mean rate of return to the first four years of university is given by the sum of the parameters β, ε, and λ; while the rate of return to the last year of university is given by the sum of β, ε, and φ. An alternative, which suppresses the three dummy variables (but not the interaction terms), gives the mean rate of return to education by level.

I also consider two semi-parametric regression models. In the first, experience and squared experience are still assumed to have an additive effect on log wages, but education is captured by fifteen schooling dummies:

(3) $\log W_i = \alpha + \varsigma_x S_i + \varepsilon X_i + \lambda X^2_i + \varepsilon_i,$

This imposes fewer restrictions than a spline function (see also Hungerford and Solon 1987; Lam and Schoeni 1993; Strauss and Thomas 1996). Finally, I estimate a very flexible regression model in which the log of the individual wage rate is regressed on a set of 15 schooling dummies, 10 experience dummies, and 150 interactions between experience and schooling:

(4) $\log W_{isx} = \varsigma_x S_{is} + \varepsilon_s X_{ix} + \lambda_{sx} (S_{ie} * X_{ix}) + \iota_{isx},$

where the subscripts i, s, and x stand for individuals, years of schooling, and experience cohort. Years of experience have been lumped together into five-year groups, such that the first group (experience = 1) corresponds to workers with 0 to 5 years of experience, the second group (experience = 2) corresponds to workers with 6 to 10 years of experience, and so on. I do this to avoid the very small sample sizes which would have resulted had I created dummy variables for every year of experience. This model therefore estimates 160 coefficients (fifteen dummy variables have to be dropped to avoid perfect colinearity). An important appeal of this approach is that it allows the effect of education on wages to vary by experience (see Card 1999, pp. 1805-06). An obvious limitation is that this estimation strategy will produce noisy estimates in the absence of very large sample sizes.

5. Results

I present summary statistics on the sample of wage-earning men in the APIS in Table 1, non-parametric density estimates of the logarithm of the wage distribution in Figure 1, and histograms for the years of completed schooling in Figure 2. All of these results are weighted by the appropriate expansion factors, and are presented for three groups: All men who report earning wages, men aged 25-65, and a sub-sample of men for whom we have information on the number of years of completed schooling of both parents. Table 1 shows that these men tend to be much younger: The mean age is 33 for all wage-earning men, 38 for men aged 25 to 65, and 22 for men for whom information on parental education is available. The reason for this is that we have information on parental education only for those children of the household head who are still living at home, in households where both parents are present. Figures 1 shows that the distribution of log wages is approximately normal for all three groups, with means at 2.71 (all men), 2.86 (men aged 25 to 65), and 2.39 (the younger men). The lower mean wages of the younger men are, presumably, a result of less years of labor market experience, as well as of the fact that these younger men have slightly less education (8.5 years, rather than 8.8 years for all wage-earning men in the sample, or 9.1 years for men aged 25 to 65).[7] Figure 2 shows that these means hide important spikes in the distribution. In the full sample, for example, 19.9% of the men have completed exactly six years of education, 25.5% have completed exactly ten years, and

[7] These summary statistics make apparent some of the limitations of the sub-sample of men for whom we have information on parental education: The sample size is small (2,355 observations), and some of these young men may not have completed their schooling. The fact that all of these men are still living at home raises a possible selection problem, since formation of an independent household is arguably endogenous.

12.5% have completed exactly fifteen years, corresponding to graduation from primary school, secondary school, and university, respectively.

Table 2 presents results based on the Mincerian framework for all wage-earners, with and without provincial level dummies (columns 1 and 2), for men aged 25-65 only (column 3), and for the sub-sample of men for whom we have information on the education of both parents (columns 4 and 5). Table 2 shows that the mean rate of return to education is large and significant—between .110 and .148. Inclusion of provincial level dummies reduces the estimated coefficient by about 12.7%, suggesting that there may be important omitted province-level variables.[8] The education of both parents has a significant, positive impact on log wages, with the effect of father's education being noticeably larger;[9] a comparison of the results in columns 4 and 5 shows that inclusion of controls for parental education reduces the coefficient on own education by about 17%. There is no way of telling from these data whether this difference can be attributed primarily to a reduction in ability bias or to an increase in the bias due to measurement error.[10]

Tables 3 presents the results from estimating the spline function models for the sample of Filipino men in the APIS. I present ten sets of results, corresponding to specifications for all males reporting wage earnings, with and without provincial dummies (columns 1 and 2), for the smaller sample of men aged 25 to 65 (column 3), and for the sample of men for whom we have information on parental education (columns 4 and 5). Results for the full model, including the years-of-education dummy variables, are reported in the upper panel of the table, while results without these dummies are reported in the lower panel.

A number of things are worth noting about the results in Table 3. First, although the change in the R-squared is modest, a joint F-test on the additional coefficients shows that the more flexible spline function is a significant improvement in fit over the prototypical Mincerian equation in all of the specifications, including those with the smaller sample sizes. Second, the mean rate of return appears to vary considerably by level. The continuous spline specification for the full sample of men, for example, indicates that the mean rate of return to education in the

[8] If provinces with higher average education levels also have higher wage premia associated with an additional year of schooling, inclusion of provincial dummies will lead to a decrease on the coefficient on years of schooling. Persistently higher wage premia in some provinces could be explained, in turn, by a positive correlation between the quantity and quality of schooling, or by differences across provinces in the relative demand for and supply of workers with different amounts of education in conjunction with imperfect labor mobility.

[9] Contrast this with the results in Heckman and Hotz (1986), who find that mother's schooling has a larger impact than father's schooling on wages for a sample of Panamanian men.

Philippines is smallest for primary education (.094), slightly larger for secondary education (.100), although the increase is not significant, and much larger for tertiary education (.167), a change in slope which is significant at any conventional level. This would suggest considerable convexity in the relationship between the wage rate and years of completed schooling. Adding the provincial dummies further increases the difference between the mean rate of return for a year of primary school (.062) and a year of university (.171). When parental education is included as a control, the slope of the relationship between log wages and education drops both for primary education (from .094 to .078) and university (from .196 to .162). Finally, there are large, significant sheepskin effects, especially at the university level. In the full sample, these range from .121 for primary school, to .129 for secondary school, to a whopping .276 for university. Including provincial-level dummies greatly reduces the sheepskin effects at the primary and secondary levels, but not at university. Note that the sheepskin effects are less sensitive to the inclusion of the controls for parental education, although the estimates are not always significant because of the small sample sizes.

Before discussing some possible explanations for these results, I present estimates based on the semi-parametric specification in equation (3) in Table 4 and Figure 3. Like Table 3, these estimates show differences in the mean slope of the earnings function between primary education and university education. For example, in the full sample of men, without the provincial dummy variables, the (unweighted) mean step size for the first six years of education is .089, compared to a mean step size for the last five years of education of .167—almost twice as large. Note also that the slope of the earnings function at the primary and secondary levels is much flatter when provincial dummies are included, consistent with the results in Table 3. Controls for parental education reduce the step-sizes throughout. Finally, the results in Table 4 also show very large sheepskin effects for graduation from university. The sheepskin effects associated with primary and secondary school are smaller, and disappear when provincial-level dummies are included in the regression.

What explains the convexity of the earnings function for Filipino men? No doubt, omitted individual-level variables, such as ability, motivation, and family connections, all of which are probably correlated with schooling, play a role. But the results in Tables 3 and 4 show that controls for parental education reduce all the estimated coefficients on own schooling, and have only a modest effect on the *differences* in the estimated rates of return across levels.

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[10] Using data on the Bicol region, Maluccio (1998) estimates that the reliability ratio of own schooling is 85% to 90%, and that measurement error accounts for a downward bias of about 20% in the rate of return to education estimated by simple OLS.

Of course, measures of parental education are not perfect controls for ability. Further informal evidence that much of the observed difference in the rate of return across levels is not explained by self-selection into more schooling comes from a comparison of estimates from the United States and the Philippines. The correlation between schooling and ability is probably no bigger in the Philippines than in the United States. If anything, one would expect the converse to be the case: In the Philippines, where there is no functioning credit market for student loans, and where the average direct private cost borne by households of sending a single child to university is estimated to be almost one half of mean household income, there are probably many more high-ability students from disadvantaged backgrounds who cannot afford to go to university than in the United States. And yet, in the results presented by Hungerford and Solon (1987) for the United States, for instance, the mean step-size for college is *smaller* than the mean step size for primary education. This does not seem unreasonable: It is what one would expect if there were diminishing marginal returns to education.

Ability does not seem to be the only culprit. Two other possibilities come to mind. First, the high rates of return to tertiary education could be the result of differences in the relative demand for and supply of workers with different amounts of education. Some support for this idea is given by gross enrollment figures for the Philippines: Between 1980 and 1997, gross enrollment ratios increased in primary school (by 4.4%, from 113 to 118) and secondary school (by 15.4%, from 65 to 75), and decreased at the tertiary level (by 19.2%, from 26 to 21) (Asian Development Bank and World Bank 1999, p. 149).[11] Moreover, these changes in supply took place at a time when, if anything, there appears to have been an increase in the demand for more skilled workers in the Philippines—for example, in the booming electronics sector. Disaggregated data on sector of employment, which was collected in the APIS but was not

[11] However, gross enrollment rates at the tertiary level in the Philippines were well above those of its neighbors throughout the 1990's, including Thailand, China, Indonesia, and Malaysia (but not Korea) (World Development Indicators database, The World Bank). There is also some discrepancy between these figures and those in the World Development Indicators database, which actually show an *increasing* enrollment ratio for tertiary education in the Philippines over the 1990's. The difference between these estimates appears to be that the World Development Indicators include enrollment in non-university tertiary institutions, whereas the Asian Development Bank and World Bank (1999) report do not. Alternatively, I could have used data on educational attainment (rather than enrollment), such as those found in Barro and Lee (1993) or Nehru, Swanson and Dubey (1994), for the comparisons in this paper. These data sets have been criticized on numerous grounds (Behrman and Rosenzweig 1994). For the purposes of my analysis, the main disadvantage of these data sets is that they do not provide post-1987 data. See also Pritchett (1996).

available for this paper, might help clarify this point by showing whether there were large differences in wage premia across sectors.[12]

A second explanation might be that there are large differences in quality across levels in the Philippines, with higher quality education being provided in university (for a general discussion of this issue see Strauss and Thomas 1995, pp. 1971-72, as well as Behrman and Birdsall 1983). If there are indeed quality differentials, these could be related to a quality-diluting expansion in primary and secondary enrollments. Moreover, a large fraction of students currently attending university in the Philippines are enrolled in technical subjects, such as engineering, which are likely to be associated with high rates of return (World Bank 2000, pp. 15-16). Whether the fact that a higher fraction of university education than primary or secondary education is provided by the private sector also has implications for quality is a matter of conjecture, although Jimenez et. al. (1995) argue that this is plausible.

What accounts for the very large sheepskin effects found in the Philippines? Here too omitted level variables, such as motivation, probably play a part. But the sheepskin effects, especially for graduation from university, are quite insensitive to inclusion of controls for parental ability. And once again, the size of the sheepskin effect for graduation from college is much larger in the Philippines than in the United States: .276 compared to .090 for the discontinuous spline specification, and .382 compared to .176 for the semi-parametric specification. The very large sheepskin effects estimated for the Philippines are probably explained, at least in part, by information imperfections in the labor market which lead employers to rely on credentials as a shortcut to assessing skills and future productivity. Surprisingly, the sheepskin effects associated with graduation from primary school, secondary school, and university in the Philippines appear to be *larger* in the private sector than in the public sector: Separate estimation of discontinuous spline functions for government and private sector employees for the full sample of men, for example, shows a university graduation effect that is more than twice as large in the private sector (.343, significant at the 0.1% level) as in the public sector (.153, significant at the 5% level).[13]

Figure 4, finally, presents results based on an estimation of equation (4) above for the full sample of men. Each point on a graph corresponds to the expected earnings for a given

[12] Pritchett (1996), for example, argues that the rate of return to education is generally highest in sectors in which there is exogenous technological progress. Amongst other evidence, he cites work by Foster and Rosenzweig (1996) on India, which suggests that rates of return to education went up in villages where the Green Revolution brought agricultural innovation, but not in technologically-stable villages.

[13] Note also that Table 4 and Figure 3 show that the estimated sheepskin effects are not the result of a particularly small increase in earnings for those who have all-but-the-last year of schooling within a given level (that is, people with 5, 9, or 14 years of education), something which has been argued for the United

combination of education and experience, while the vertical lines correspond to six years of education (completed primary) and ten years of education (completed secondary). Figure 4 provides some further visual (and noisy!) support for the results in Tables 3 and 4. There frequently appears to be a marked step in log wages upon graduation from a given level, especially university; and the slope of the earnings function often looks flatter for the first ten years of education than for the last five—especially for some of the older cohorts. There appears to be no pattern in the relationship between sheepskin effects and experience. For example, the wage premia associated with completion of the last year of university are .504, .374, .387, .213, .355, .488, .180, .240, -.010, and 1.083 for the first through tenth experience cohorts, suggesting that credentials are no less important later in a worker's life.[14] Note, finally, that it is impossible to disentangle cohort effects from period effects with a single cross-section of data, so that it is not clear what could be driving the difference in the relative slopes of the lines across the graphs. For example, the steeper slope for the first years of education in the first graphs could be due to a recent improvement in primary or secondary education (a period effect) or the fact that the low returns to primary and secondary education only bite in later years of a worker's labor market experience (a cohort effect).[15]

6. Estimating the expected earnings of primary school, secondary school and college graduates

I turn finally to some very simple simulations of the effect which differences in the rate of return might have on the expected earnings streams of graduates from primary school, secondary school, and university. To answer this question, note that the points in Figure 4 can also be grouped by years of education (rather than by experience category). The net present value of total lifetime earnings could then be approximated by the area below a line connecting the wage-experience interactions for a given number of years of education, once the scale of the y-axes had

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States (Park 1994). It is the last year of schooling within a given level, rather than the year immediately before it, which appears to be an unexpected departure from linearity in the Philippines.

[14] Another way of estimating the effect of experience on credentials, suggested to me by Harold Alderman, would interact the dummy variables D6, D10, and D15 in the discontinuous spline specification with experience. The results from such an estimation strategy show that the coefficients on these interactions tend to be significantly negative for graduation from primary school and secondary school, and significantly positive for graduation from university.

[15] With repeated cross-sections over many years, it might be possible to relate high wages across cohorts in a given year to period effects (for example, a strong labor market), and persistently high wages for a given cohort across years to cohort effects (for example, a better-than-average education). One shortcoming of such an approach, however, is that if there is persistence in wages—so that workers who start at low wages tend to continue with low wages—then period effects can (mistakenly) begin to look like cohort effects.

been adjusted in order to turn hourly wages into yearly wages, and future earnings had been appropriately discounted.

These simulations should not be taken too seriously, since they make two important assumptions. First, they do not take into account underlying individual differences in ability, motivation, or family connections, all of which can have an effect on wages. Second, they assume that information on the rates of return of earlier cohorts can be applied to today's labor market entrants, which may not hold (although it is unclear what additional information potential labor market entrants might have). The purpose of the simulations is merely to give a rough picture of the effects of large differences in the rates of return to education.

To reduce the biases which could be introduced by the choice of functional form, and to avoid having to continue working with five-year experience groups, I run nonparametric (kernel) regressions of log hourly wages on years of experience for the full sample of wage-earning men who have completed exactly six, ten, and fifteen years of schooling (corresponding to primary school, secondary school, and university graduates).[16] I take retirement from the labor force to take place no later than at age 65, although higher values turn out not to make much difference to the results. The graphs corresponding to these regressions are presented in Figure 5.[17]

The direct private costs of education also play a role in determining the relative benefits of different amounts of schooling. Disaggregated data on the private costs of education per household are available in the 1997 Family Income and Expenditure Survey (FIES), but not in the 1998 APIS, while data on household composition, including the number of children attending different education levels are available in the APIS, but not in the FIES. I use the panel of households included in both the FIES and APIS to estimate the private unit cost of primary, secondary, and tertiary education by running an auxiliary regression of the total amount spent by a household on education in 1997 on the number of children attending each education level in 1998:

$$(5) \qquad Ex_h = \hat{a}P_h + \tilde{a}S_h + \ddot{a}T_h + \mathring{a}_h$$

[16] All nonparametric regressions are Fan regressions with a quartic kernel (see Fan, 1992). The bandwidth was fixed by ocular inspection and trial and error.

[17] Note that these regressions directly estimate the log hourly wage rate for all individuals in the sample, including the youngest ones (children who have completed primary school and have no experience, aged 12 years old). This approach is quite different from that frequently applied elsewhere, where the wages of minors are taken to be some arbitrary fraction of those earned by adults. For example, Tan and Paqueo (1989) assume that children aged 11-12, 13-16, and 17-18 earn 20%, 50% and 75% of the earnings of adults aged 19, respectively.

where Ex is the total amount spent by household h on education in 1997, P, S, and T are the total number of children in household h who are attending primary school, secondary school, and university in 1998, and the constant in the regression has been suppressed. The parameters â, ã, and ä will provide a reasonable approximation to the private unit costs, by level, if attendance rates in 1998 are close to those in 1997.[18]

Estimates of the private unit costs of education are presented in Table 5 for all households, and for households broken down by 1997 consumption quintiles. I report both the absolute figures for expenditures, as well as the fraction that these represent of total household consumption. Table 5 shows that the private unit costs of education increase steeply by level: The mean unit cost of primary education is 898 pesos, while the mean unit cost of tertiary education is 13,334 pesos—almost 50% of mean household consumption. Table 5 also shows that there are even larger differences in the unit costs for a given education level by income quintile: The richest households spend 19 times as much as the poorest households for every child in primary school, and almost 10 times as much for every child in university education.

I combine information on the average unit costs estimated in the first column of Table 5 and the profile of expected lifetime earnings in Figure 5 to compare the total expected lifetime earnings of individuals who have completed primary school, secondary school, and college. To understand how this works, consider a student who has just completed primary school. This student has two options: To enter the labor market and begin to earn wages, or continue in school, paying the private cost of education and deferring the wages he will earn. Given that he has completed primary school, the net present value of total expected lifetime earnings if he enters the market immediately can be represented by:

$$(6) \qquad NPV(Y_s) = W_{s,x} + W_{s,x+1}/(1+r) + W_{s,x+2}/(1+r)^2 + \ldots + W_{s,x+n}/(1+r)^n,$$

[18] I would particularly like to thank Hans Hoogeven for help with this calculation. This procedure assumes that household composition, the fraction of children who are attending school at every level, and actual (rather than measured) unit costs remained constant between 1997 and 1998. There is no reason to suppose that household composition or unit costs changed in the Philippines between 1997 and 1998. Attendance rates may have dropped, however, because of the effect of the East Asian crisis on incomes in 1998, even though the Philippines was much less affected by the crisis than other countries in the region. Administrative data show no significant impact of the crisis on *enrollment* rates in primary or secondary school; but attendance rates may still have fallen if households were more likely to keep their children at home or doing odd jobs. A weaker assumption for the calculation of unit costs is that attendance rates in primary school, secondary school, and university in 1998 are a constant *fraction* of attendance rates in 1997, such that $E_p^{98} = øE_p^{97}$, $E_s^{98} = øE_s^{97}$, and $E_c^{98} = øE_c^{97}$, where the subscripts stand for education levels, superscripts for years, and the parameter ø is less than one. Under this assumption, the proposed

where W is the total wages earned in a year, r is the discount rate, n+1 is the total number of years in the labor market, and the subscripts s and x stand for the years of schooling and experience, respectively. Note that the quantities W need not be the same at each level of experience x. Indeed, they will change as additional years of experience are rewarded (or penalized) in the labor market. Contrast this with the net present value of total expected lifetime earnings if this young man decides to postpone entering the labor market in order to acquire z more years of schooling. If the total number of years a man can devote to experience or schooling is fixed at n+1, this can be described by:

$$(7) \qquad NPV(Y_{s+z}) = -C - C/(1+r) - C/(1+r)^2 - \ldots - C/(1+r)^{z-1} + W_{s+z,x}/(1+r)^z + W_{s+z,x+1}/(1+r)^{z+1} + \ldots + W_{s+z,x+n-z}/(1+r)^n,$$

where C is the direct private cost of each additional year of education, which is allowed to vary by education level. A comparison of equations (6) and (7) makes apparent the nature of the trade-off between labor market participation and schooling. If more schooling is valued in the labor market, the expected wage rate $W_{s+z,x}$ should exceed $W_{s,x}$, and $W_{s+z,x+1}$ should also exceed $W_{s,x+1}$, although the size of the gap between the wage rate for those with s and s + z years of schooling need not be constant for all years of experience x. There are, however, three factors which reduce the value of additional schooling. First, there are direct private costs of schooling. Second, for a fixed retirement age, more schooling means fewer years during which these (higher) wages can be earned, so that there are n +1-z years of experience rather than n + 1. Finally, because earnings are deferred with more schooling, they will be discounted. In sum, additional schooling will seem most attractive when there are large wage differentials, long periods of expected participation in the labor market, small private costs, and low discount rates.

In Figure 6, I present results for Filipino men for three different discount rates—0.03, 0.06, and 0.09.[19] Figure 6 shows that primary schooling only is the strategy which maximizes expected lifetime earnings for individuals who intend to stay in the labor market for only a short period of time—13 years or less at the lowest discount rate, and 19 years or less at the highest discount rate. Otherwise, completing college is the dominant strategy.[20] Of course, this is only a reasonable

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estimation procedure will over-estimate unit costs by a factor 1/ø, but will produce accurate estimates of the *ratio* of unit costs across levels.

[19] The average interest rate on deposits in the Philippines between 1990 and 1999 was 12.3%, and average inflation was 9.6% (International Monetary Fund, various years).

[20] Note that these thresholds may be high for women, who often spend a large number of years outside the labor market during their childbearing years. Insofar as the basic pattern of earnings streams holds for women as well, therefore, women may rationally choose not to invest in education above the primary level.

approximation for the marginal student: If many more workers were to acquire university education the returns to university graduates would presumably drop as well. Still, the results in Figure 6 do suggest that the high marginal returns to completed tertiary education in the Philippines make a college degree a very good investment indeed. The private costs of tertiary education are also high, but these costs are offset by the higher wages earned by college graduates if they remain in the labor market for a long period of time. Even so, the direct private costs of tertiary education may be too high for students from poor households—especially since they are unlikely to be able to borrow against future income.

7. Conclusion

The Philippines has made important strides extending primary and secondary education to many of its citizens. Even so, or perhaps as a result, the private rates of return to primary and secondary schooling are low relative to those associated with university education. This paper also shows that there are very large sheepskin effects—much larger than those found in the United States—for graduation from primary school, secondary school, and especially university. One puzzling question which this paper does not address is how the high education levels and the reasonably high overall returns to education are consistent with the declining productivity and very low levels of growth found in the Philippines in the last two decades (for a general discussion, see Pritchett 1996).

Returns to education are not static, in the Philippines or elsewhere, and the shape of the earnings function may well change over time.[21] A more ambitious research agenda would focus on a developing country or countries with a number of large cross-sectional data sets. One further advantage of this approach is that it would allow for distinctions to be made between the period and cohort effects which underlie changes in the relative shape of the earnings function.

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[21] Indeed, because 1998 was a year in which the Philippine economy was affected by the East Asian crisis, it might not be a representative year to study the relationship between wages and education—even though the effects of the crisis were nowhere near in magnitude to those felt by many of the Philippines' neighbors: The year-on-year change in real GDP per capita for the third quarter of 1998, corresponding to the period for the wage data in our sample, was only –2.2%.

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Table 1: Summary Statistics

Variable	Median	Mean	Standard Deviation
ALL MEN			
Log Hourly Wage Rate (1998 Filipino pesos)	2.75	2.71	.859
Weekly Hours Worked	43.69	41.12	15.08
Age	33	34.29	12.38
Years of Completed Schooling	9	8.78	3.56
MEN AGED 25 TO 65			
Log Hourly Wage Rate (1998 Filipino pesos)	2.89	2.86	.821
Weekly Hours Worked	44.31	42.34	14.25
Age	38	38.96	9.77
Years of Completed Schooling	10	9.06	3.62
MEN WITH AVAILABLE DATA ON PARENTAL EDUCATION			
Log Hourly Wage Rate (1998 Filipino pesos)	2.36	2.39	.909
Weekly Hours Worked	40.62	37.09	16.53
Age	22	22.20	5.65
Years of Completed Schooling	9	8.53	3.58

Table 2: Returns to Education, Log-Linear Specification

	Specification				
	1	**2**	**3**	**4**	**5**
Years of Completed Schooling	.126***	.110***	.123***	.148***	.123***
	(.002)	(.002)	(.002)	(.005)	(.006)
Experience	.043***	.042***	.031***	.058***	.061***
	(.001)	(.001)	(.002)	(.008)	(.008)
Squared Experience (x 100)	-.056***	-.056***	-.039***	-.116**	-.125***
	(.003)	(.002)	(.004)	(.033)	(.033)
Father's education	--	--	--	--	.029***
					(.007)
Mother's education	--	--	--	--	.017*
					(.007)
Provincial dummies	No	Yes	No	No	No
R-squared	.300	.469	.253	.338	.357
Number of observations	21,765	21,765	16,738	2,355	2,355

Note: The dependent variable is the log of hourly wages. Coefficients and standard errors corrected for heteroskedasticity and clustering are reported in parentheses. Specifications 1 and 2 include all men who report earning wage income; Specification 3 is limited to men aged 25 to 65; Specifications 4 and 5 are limited to men whose relationship to the household head is "son" in households in which both the household head and his or her spouse are members and report their education levels. A constant was estimated but is not reported.

* Significant at the 5% level; ** Significant at the 1% level; *** Significant at the 0.1% level.

Table 3: Returns to Education, Spline Specification

	Specification				
	1	**2**	**3**	**4**	**5**
Experience	.043***	.044***	.036***	.058***	.060***
	(.001)	(.001)	(.002)	(.008)	(.008)
Squared Experience (x 100)	-.060***	-.061***	-.050***	-.121***	-.128***
	(.003)	(.002)	(.004)	(.033)	(.032)
Years of Completed Schooling	.062***	.054***	.058***	.055	.043
	(.011)	(.010)	(.013)	(.031)	(.032)
Dummy for S> = 6 (D6)	.121***	.031	.098**	.156	.142
	(.031)	(.027)	(.036)	(.091)	(.090)
D6 * (S-6)	.008	.008	.008	.039	.038
	(.013)	(.011)	(.015)	(.039)	(.040)
Dummy for S> = 10 (D10)	.129***	.047*	.124***	.135	.137
	(.025)	(.021)	(.030)	(.072)	(.071)
D10 * (S-10)	.043***	.058***	.042***	.033	.014
	(.010)	(.009)	(.012)	(.040)	(.038)
Dummy for S = 15	.276***	.269***	.288***	.340*	.336*
	(.037)	(.032)	(.040)	(.164)	(.158)
Father's Education	--	--	--	--	.027***
					(.007)
Mother's Education	--	--	--	--	.014*
					(.007)
Provincial Dummies	No	Yes	No	No	No
R-squared	.311	.489	.267	.352	.367
F-test	48.37***	112.27***	46.81***	8.50***	5.62***
Number of observations	21,765	21,765	16,738	2,355	2,355
Experience	.043***	.044***	.036***	.059***	.061***
	(.001)	(.001)	(.002)	(.008)	(.008)
Squared Experience (x 100)	-.060***	-.060***	-.049***	-.124***	-.131***
	(.003)	(.002)	(.004)	(.032)	(.032)
Years of Completed Schooling	.094***	.062***	.084***	.094***	.078***
	(.006)	(.005)	(.007)	(.019)	(.020)
D6 * (S-6)	.006	.007	.009	.035	.037
	(.008)	(.007)	(.009)	(.026)	(.026)
D10 * (S-10)	.067***	.102***	.072***	.067***	.047**
	(.006)	(.005)	(.007)	(.018)	(.018)
Father's Education	--	--	--	--	.027***
					(.007)
Mother's Education	--	--	--	--	.014
					(.007)
Provincial Dummies	No	Yes	No	No	No
R-squared	.308	.487	.264	.348	.364
F-test	81.80***	263.40***	84.99***	14.81***	8.43***
Number of observations	21,765	21,765	16,738	2,355	2,355

Note: The dependent variable is the log of hourly wages. Coefficients and standard errors corrected for heteroskedasticity and clustering are reported in parentheses. Specifications 1 and 2 include all men who report earning wage income; Specification 3 is limited to men aged 25 to 65; Specifications 4 and 5 are limited to men whose relationship to the household head is "son" in households in which both the household head and his or her spouse are members and report their education levels. A constant was estimated but is not reported.
F-test: F-test in columns 1 and 3 is joint F-test on the two interaction terms; F-test on columns 2 and 4 is joint F-test on the two interaction terms and the three dummy variables.
* Significant at the 5% level; ** Significant at the 1% level; *** Significant at the 0.1% level.

Table 4: Returns to Education, Semi-Parametric Specification

	Specification				
	1	**2**	**3**	**4**	**5**
Experience	.043***	.044***	.036***	.058***	.060***
	(.001)	(.001)	(.002)	(.008)	(.008)
Squared Experience (x 100)	-.060***	-.061***	-.050***	-.123***	-.129***
	(.003)	(.002)	(.000)	(.033)	(.032)
S = 1	.100	.089	-.049	-.127	-.197
	(.153)	(.128)	(.164)	(.272)	(.282)
2	.024	.052	-.001	.169	.187
	(.072)	(.060)	(.079)	(.224)	(.238)
3	.138**	.110	.180**	-.093	-.120
	(.049)	(.042)	(.058)	(.129)	(.133)
4	.017	.033	-.014	.124	.117
	(.037)	(.030)	(.044)	(.104)	(.104)
5	.061	.025	.072	.078	.065
	(.033)	(.029)	(.040)	(.087)	(.086)
6	.193***	.102***	.163***	.208*	.182*
	(.028)	(.023)	(.031)	(.082)	(.081)
7	.049	.087***	.021	.005	-.011
	(.029)	(.025)	(.036)	(.075)	(.074)
8	.072*	.026	.088*	.153	.136
	(.033)	(.028)	(.040)	(.084)	(.083)
9	.097**	.077**	.103**	.137	.132
	(.030)	(.024)	(.034)	(.081)	(.080)
10	.187***	.105***	.175***	.206**	.191**
	(.023)	(.019)	(.026)	(.066)	(.065)
11	.086**	.125***	.084	.065	.016
	(.033)	(.028)	(.038)	(.091)	(.090)
12	.141***	.119***	.113**	.163	.140
	(.037)	(.031)	(.042)	(.109)	(.108)
13	.110**	.132***	.142***	.231	.217
	(.037)	(.031)	(.039)	(.203)	(.196)
14	.118	.055	.128	.047	-.004
	(.073)	(.062)	(.078)	(.450)	(.401)
15	.382***	.441***	.361***	.466	.427
	(.070)	(.059)	(.074)	(.409)	(.358)
Father's education	--	--	--	--	.027***
					(.007)
Mother's education	--	--	--	--	.015*
					(.007)
Provincial Dummies	No	Yes	No	No	No
R-squared	.312	.489	.268	.353	.369
Number of observations	21,765	21,765	16,738	2,355	2,355

Note: The dependent variable is the log of hourly wages. Coefficients and standard errors corrected for heteroskedasticity and clustering are reported in parentheses. Specifications 1 and 2 include all men who report earning wage income; Specification 3 is limited to men aged 25 to 65; Specifications 4 and 5 are limited to men whose relationship to the household head is "son" in households in which both the household head and his or her spouse are members and report their education levels. A constant was estimated but is not reported.

* Significant at the 5% level; ** Significant at the 1% level; *** Significant at the 0.1% level.

Table 5: Direct private costs of education, by education level and income quintile

		All HH	Quintile				
			1 (poorest)	2	3	4	5 (richest)
Primary	**Amount**	898	299	557	894	1646	5681
	% of HH income	3.1%	4.0%	4.4%	4.8%	5.7%	7.2%
Secondary	**Amount**	2745	719	1367	1913	3259	7694
	% of HH income	9.4%	9.7%	10.9%	10.3%	11.2%	9.8%
Tertiary	**Amount**	13334	2127	3955	6004	8863	20740
	% of HH income	45.6%	28.7%	31.5%	32.4%	30.5%	26.4%

Note: Quintiles are 1997 consumption quintiles.

Figure 1: Density estimates of Mean Log Hourly Wages

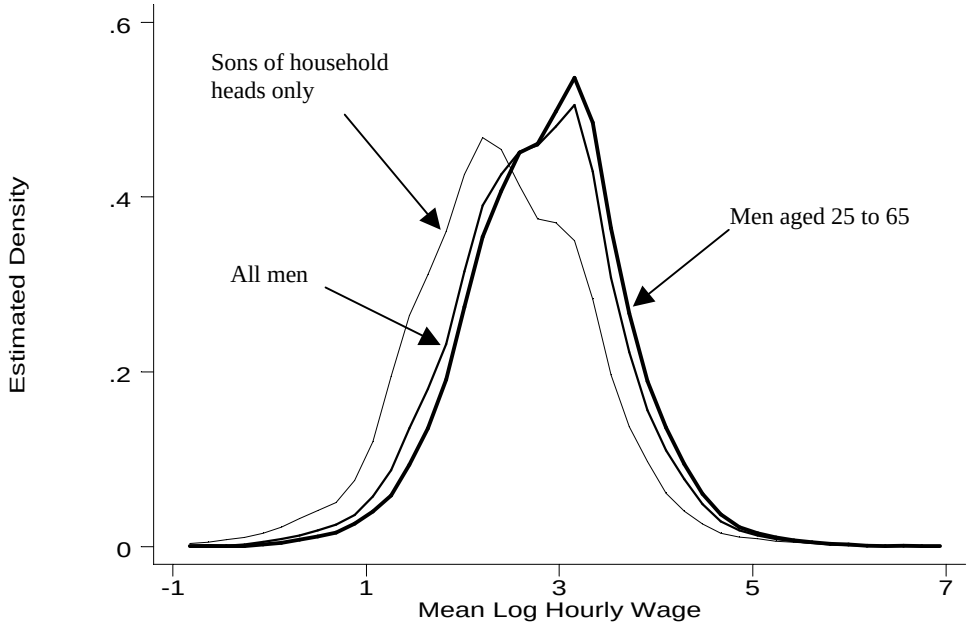
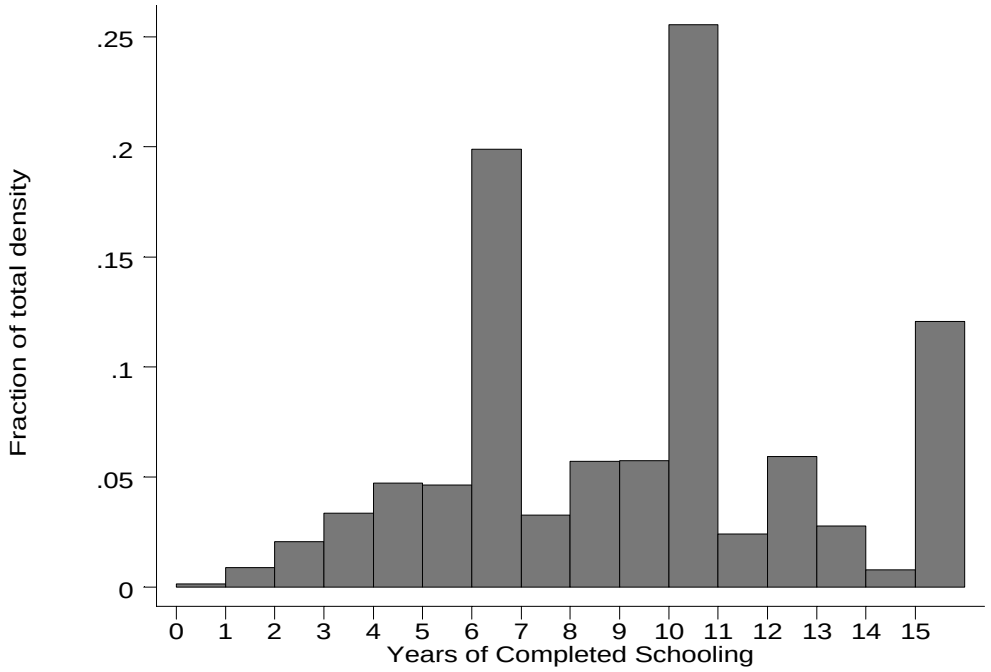
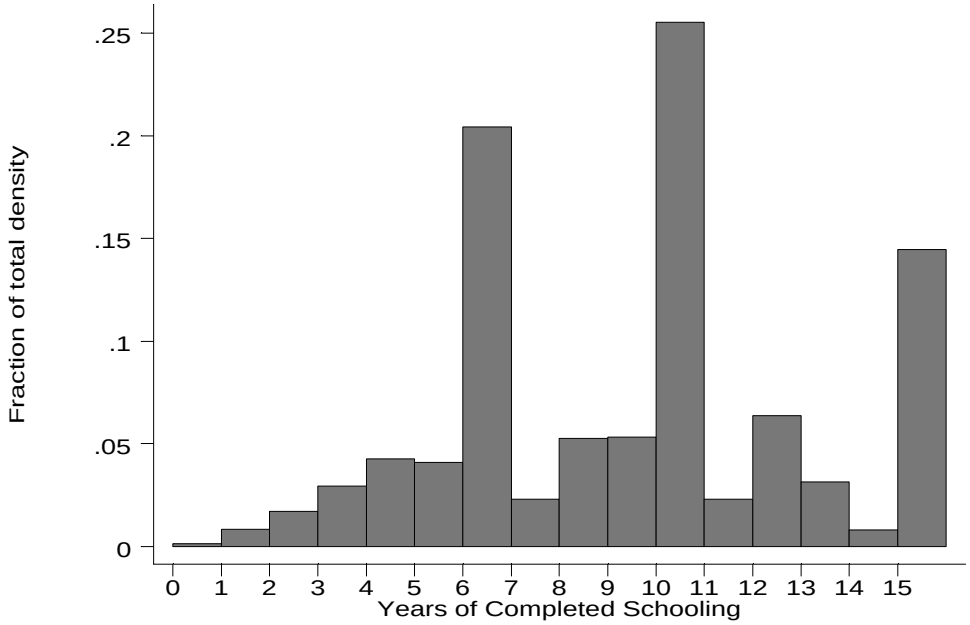


Figure 2: Histograms of Years of Completed Schooling

(i) All men



(ii) Men aged 25 to 65



(iii) Sons of household heads only

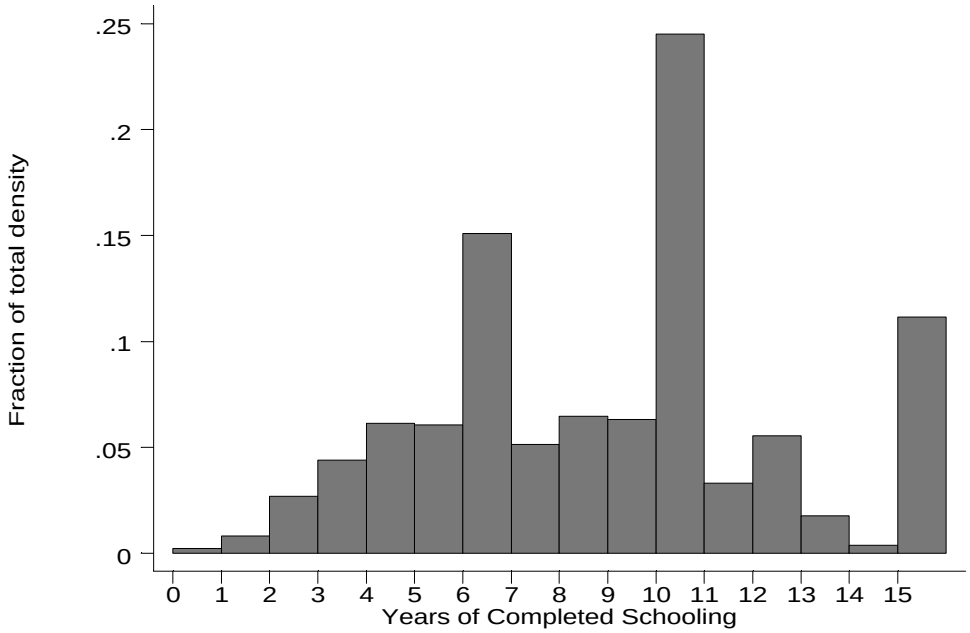
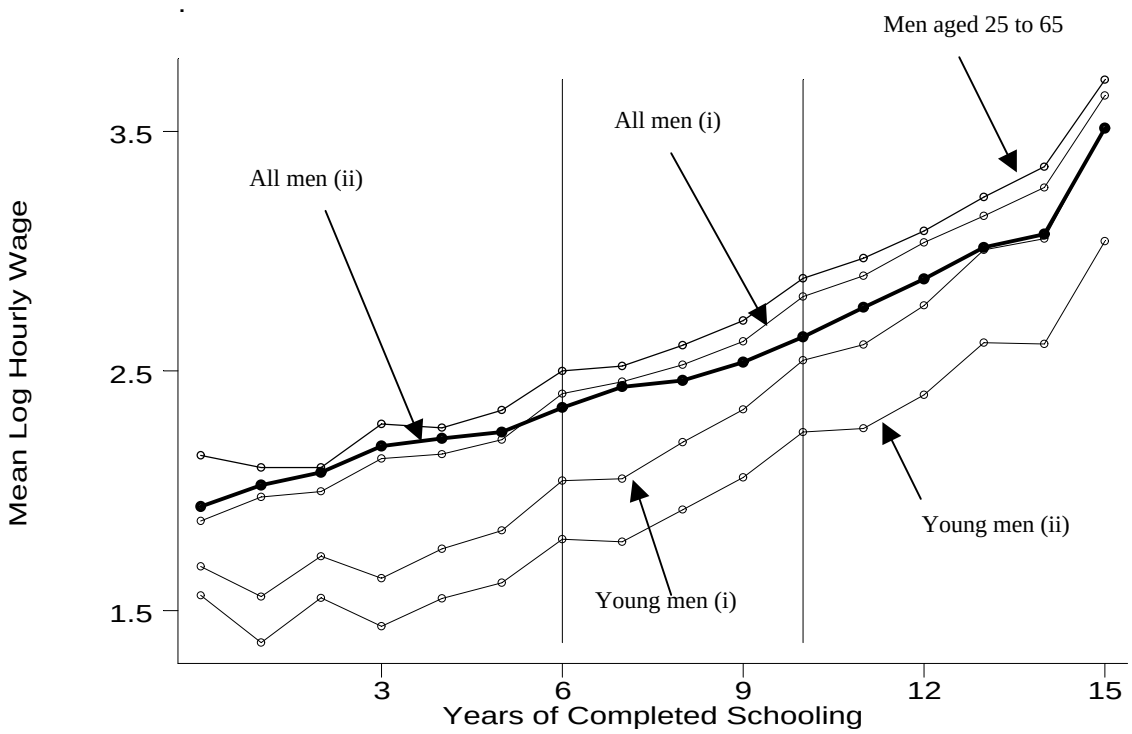


Figure 3: Mean Log Hourly Wage, by Years of Completed Schooling



Note: Coefficients for line labeled "all men (i)" come from a regression without provincial dummies, while coefficients for line labeled "all men (ii)" are from a regression with provincial dummies. Coefficients for line labeled "young men (i)" come from a regression limited to the sample of men for whom data on parents' schooling are available, but do not include controls for parental schooling; coefficients for line labeled "young men (ii)" come from a regression with this same sample which includes controls for parental education. The intercept for each of the lines includes the effect of experience at the mean of the sample.

Figure 4: Mean Log Hourly Wage, by Experience Category and Years of Completed Schooling

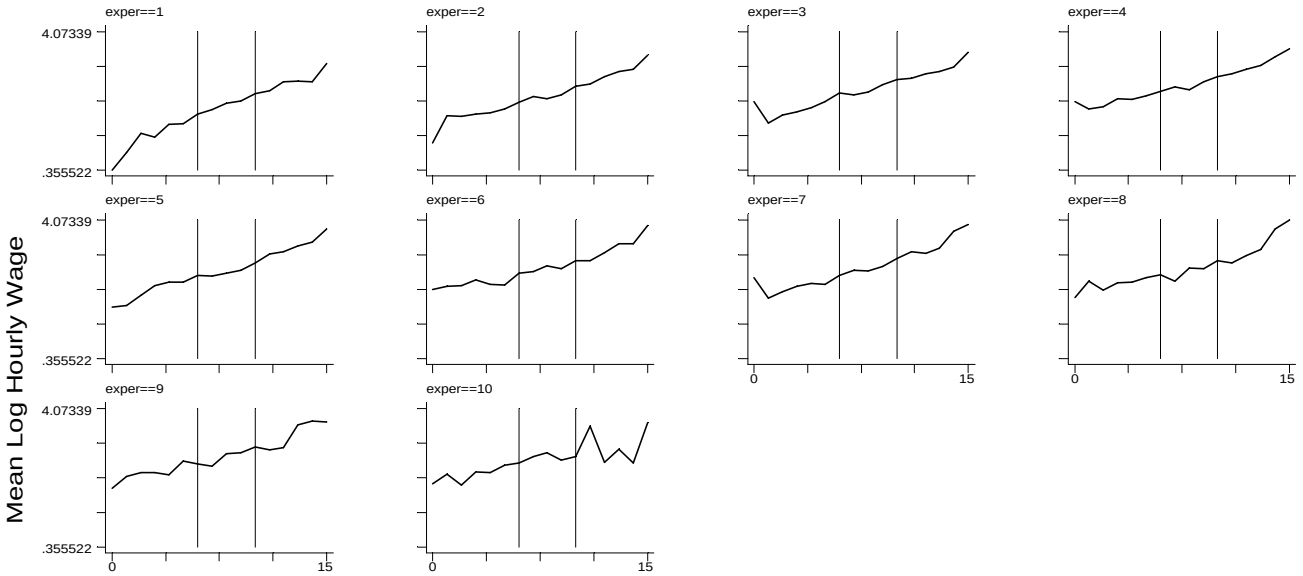


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Figure 5: Earnings Profiles for Primary School, Secondary School, and University Graduates

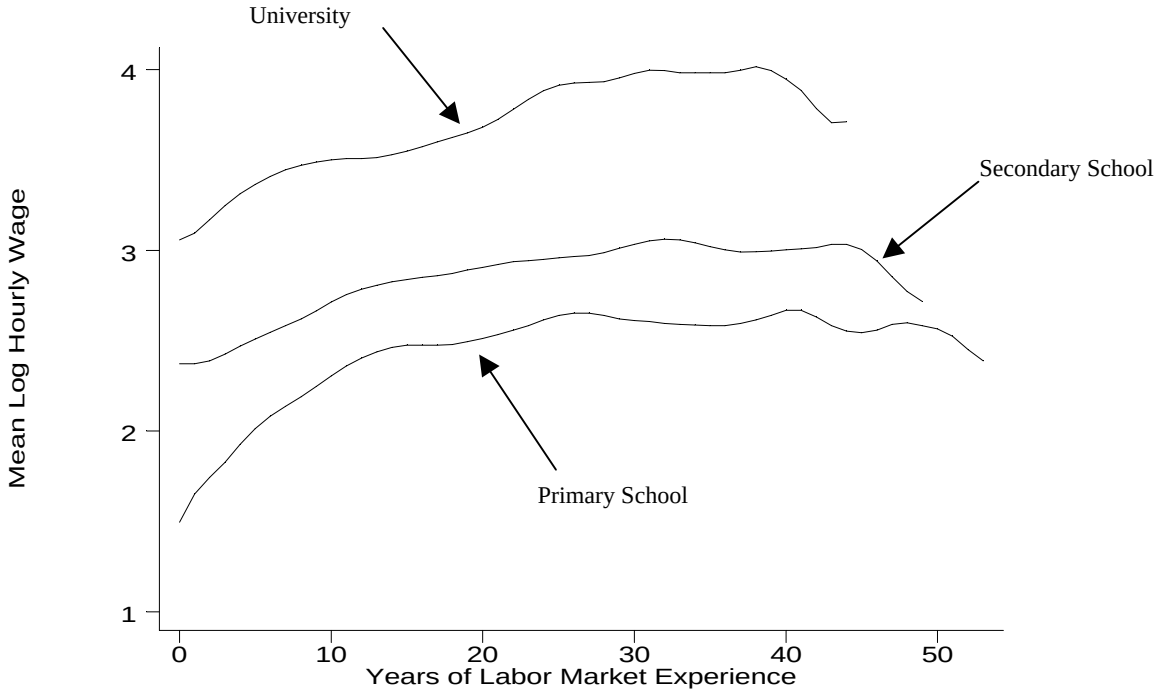
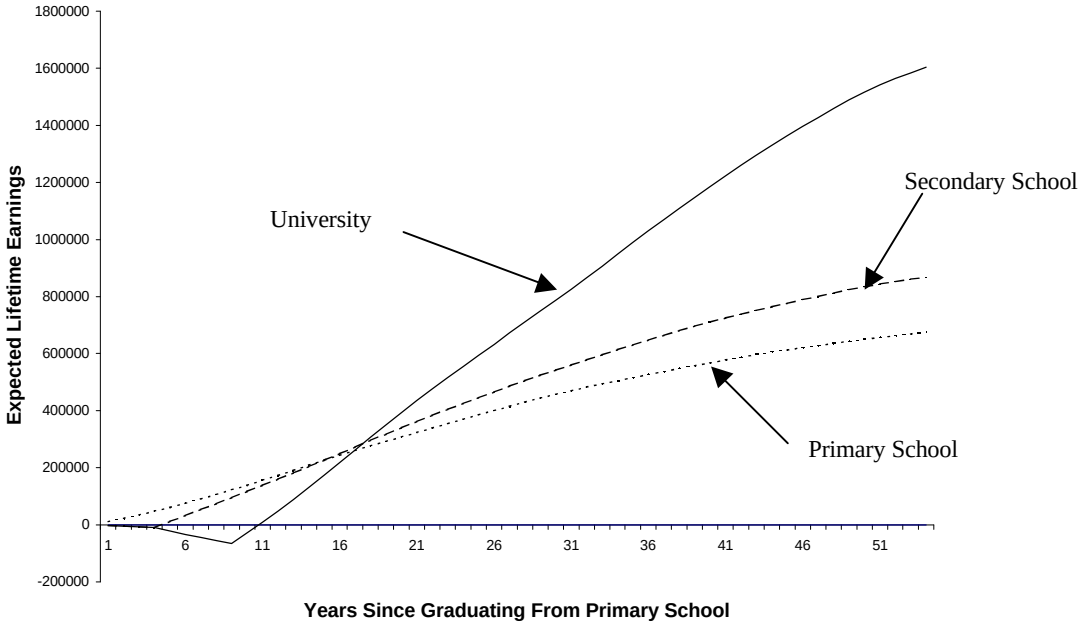
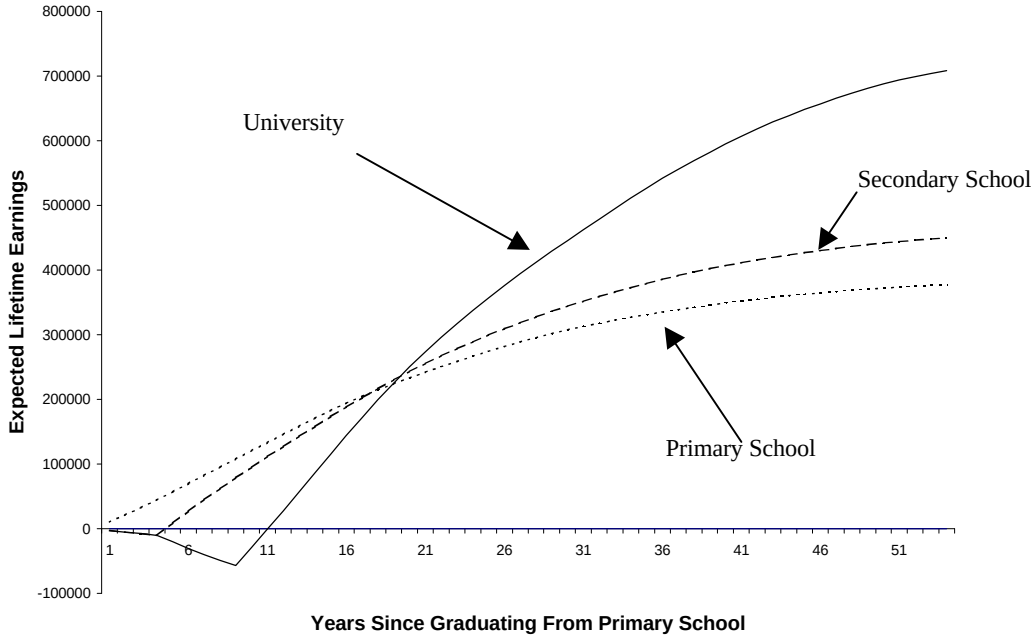


Figure 6: Expected Lifetime Earnings of Filipino Men, by Education level and Discount Rate

Discount Rate: 0.03



Discount Rate: 0.06



Discount Rate: 0.09