UNIVERSITY OF WAIKATO

Hamilton
New Zealand

Do Returns to Schooling Go Up During Transition? The Not So Contrary Case of Vietnam

Tinh T. Doan
University of Waikato

John Gibson
University of Waikato, Motu and CGD

Department of Economics

Working Paper in Economics 08/09

October 2009

John Gibson
Department of Economics
University of Waikato,
Private Bag 3105,
Hamilton, New Zealand, 3240

Fax: +64 (7) 838 4331
Email: jkgibson@waikato.ac.nz

Abstract

A key stylized fact about transition economies is that the returns to schooling rise as economic reform progresses. Existing research suggests that Vietnam is an exception to this pattern, with a decrease in males' return from 1992 to 1998, and little increase in the return to females' education (Liu, 2006). This exception may be because of the gradual economic reform applied in Vietnam, whilst in Eastern European countries the "Big Bang" transformation was conducted. Therefore to see whether Vietnam is still a counter example, we re-examine the trend in the rate of return to schooling in Vietnam over the 1998-2004 period, where the reforms have had a longer time to have an effect.

Keywords

economic transition
returns to schooling
Vietnam

JEL Classification

J31, O15

Acknowledgements

We thank, without implication, Mark Holmes and participants at the 2008 ESAM conference for useful comments, and the Vietnam Ministry of Education and Training for financial support.

1. Introduction

A key stylized fact about transition economies is that the returns to schooling tend to rise as economic reform progresses (Orazem and Vodopivec, 1995). This rise marks the movement away from distorted labor markets and the effects of longer-term changes in patterns of human capital formation. Moreover, the increase in returns to schooling in transitional economies is especially marked for women, as found in the Czech Republic and Slovakia (Chase, 1998), and in Russia, Ukraine, Hungary, and Poland (Brainerd, 1998).

Existing research suggests that Vietnam disobeys this pattern. Rates of return were very low at the beginning of transition, at only 1.6 percent in 1992 (Glewwe and Patrios, 1998; Moock et al, 2003). In contrast, global average rates of returns around the same time were nine percent (Psacharopoulos, 1994). But rather than rising through time, estimates reported by Liu (2006) suggest falling rate of returns for men in Vietnam from 1992 to 1998 and little increase for women. One possible reason for the divergence from the patterns in Eastern Europe is that gradual economic reform policies have been applied in Vietnam and the Vietnamese government has intervened in the economy substantially, whilst in Eastern European countries the "Big Bang" reforms were introduced.

Therefore to see whether Vietnam is still a counter example, we re-examine the trend in the rate of return to schooling in Vietnam over the 1998-2004 period, when the reforms may have had longer time to have an effect. Moreover, recent years have seen continued development of the private sector which has stimulated competition in the labor market with consequent changes in relative wages. Concurrently, income inequality has also risen during this later reform period, with the Gini index rising from 0.35 in 1994 to 0.42 in 2002 (GSO, 2004). Since participation rates in the wage labor market are rising over this period, the analysis relies not only on the basic Mincerian earning function but also accounts for sample selection bias. Furthermore, since there has been a changing wage premium for state sector jobs and a changing share of the wage labor force in the state sector, the analysis also allows for endogenous sector choice using a treatment effects estimator to provide more robust estimates of the trend in the returns to schooling.

The next section reviews studies of the returns to schooling in transition economies. Section 3 discusses the data and econometric specifications, while Section 4 contains the discussion of the results. Discussion of possible explanations for the changing returns and conclusions are presented in Section 5.

2. Literature on Return to Schooling in Transition

Existing studies show that rates of return to schooling increase over time in transitional economies. For example, returns to schooling increased from 3.6 percent in 1988 to 12.2 percent by 1993 in China, from 1.5 percent in 1989 to 5.4 percent by 1994 in Estonia, and from 2.9 percent in 1986 to 7 percent by 1996 in Poland (Psacharopoulos and Patrinos, 2004, Table A4). Most of these studies use Ordinary Least Squares (OLS) estimation which does not allow for endogenous school choice, but when Heckman and Li (2004) used Instrumental Variables (IV) they found even higher rates of return, of around 14 percent for four-year college attendance in China. This compares to rates of return estimated by Chow (2001) for China in the 1980s which were much closer to zero. Johnson and Chow (1997) note that employment in the stationary public sector which dominated in China's urban areas in the late 1980s leads to lower rates of returns. Zhang, Zhao, Park and Song (2005) also suggest that economic reform and technical changes have enhanced competition between workers in China, with the newly-skilled rewarded at an increasing rate.

Yet studies of Vietnam for either the single year 1992 or the 1992-98 period find low rates of return to schooling, which show little rise over time (Moock et al, 2003; Glewwe and Patrinos, 1998; Liu, 2006; Gallup, 2002). It is notable that the study by Liu (2006) was after some period of economic reform yet the estimated rate of return was still low and even appeared to slightly decrease for males. However, these studies may not have captured the full effects of the transition to a market-oriented economy given the gradual nature of reforms in Vietnam. Hence it is important to see what is the trend in the returns to schooling over 1998-2004.

In addition to the timing issue, some existing studies on Vietnam ignored the important problem of sample selection bias (Heckman, 1979). Since there was a rising participation rate in the wage labor market during transition this omission may bias not only the level of the estimated rate of return but also the trend over time. Consequently, in this paper we control for sample selection bias, and also for the related problem of the endogeneity of self-selection into the state sector. State sector employment may matter since the wage differentials in state enterprises are generally less than those in the private market economy (Johnson and Chow, 1997). A final weakness of some previous studies on Vietnam is that they merge lower secondary and upper secondary education into one group, which is implausible because the income differential between the two groups is observed to be quite large (see Table 2). So a finer disaggregation of education levels is used in this current study.

3. Data and Model Specification

Data used in this paper are from the 1998 and 2004 rounds of the Vietnam Living Standards Survey (VLSS). There are 5,999 households with 28,624 household members in 1998, and 9,188 households with 40,419 household members in 2004. Only the sub-sample aged from 15 to 60 are used in the estimation, which yields 3,365 from VLSS1998 (2,009 males and 1,356 females), and 7,177 from VLSS2004 (4,499 males and 2,678 females).

To estimate the returns to schooling, the Mincerian earnings equation is used:

$$LnY_i = \alpha + \beta_1.S_i + \beta_2 Exp_i + \beta_3 Exp_i^2 + \varepsilon_i \quad (1)$$

where: LnY is the natural log of hourly wages including bonuses, allowances, and subsidies (both in cash and in-kind), S is years of schooling, Exp is potential experience (calculated as age minus schooling years minus six) and the experience squared term, Exp^2 is added in the model to allow a non-linear pattern in lifecycle earnings.

One drawback of the basic Mincerian earnings function shown is that the coefficient β_1 in equation (1) represents the average rate of returns to one additional year of schooling regardless of the level of schooling. Hence it precludes the rate of return varying with education levels. To allow for this, an extended earnings function converts continuous years of schooling S into dummies representing the different levels of education:

$$LnY_i = \alpha + \beta_1 Prim_i + \beta_2 lowerSec_i + \beta_3 upperSec_i + \beta_4 Voc_i + \beta_5 Univ_i + \beta_7 Exp_i + \beta_8 Exp^2_i + \varepsilon_i \quad (2)$$

where: Prim, lowerSec, upperSec, Voc, and Univ are dummy variables representing primary, lower secondary, upper secondary, vocational, and university as the highest level of schooling of individual *i*. The estimated coefficients indicate the difference (higher/lower) in earnings of a corresponding education level relative to the unschooled reference group. The rate of return (r) to each year of schooling of a corresponding each level is calculated as follows:

$$r = \beta_j / S_j \quad (3)$$

where S_j is the years required to complete corresponding education level j in equation (2) (j = 1-5 for primary, lower secondary, upper secondary, vocational and university level).

3.1 Sample Selection Bias-Corrected Estimation

Sample selection bias results when the subset of wage earners used for the Mincerian earnings function is not randomly sampled from the general population. To address this problem we apply the Heckman (1979) sample selection model as follows:

Wage equation: $$w_i = z_i \beta_1 + u_{1i} \quad (4)$$

where z_i is a vector of schooling and experience variable for individual i

Selection equation: $$h_i^* = x_i \beta_2 + u_{2i} \quad (5)$$

where h_i^* is a latent variable and w_i is observed if $h_i = 1$, and $h_i = 1$ if $h_i^* > 0$, while w_i is not observed if $h_i = 0$, and $h_i = 0$ if $h_i^* \leq 0$. Furthermore, the assumptions about the errors are that: $u_{1i} \sim NID(0, \sigma^2)$ and $u_{2i} \sim N(0,1)$ and $cov(u_{1i}, u_{2i}) = \rho_{12}$.

In the first estimation stage, a binary Probit model on all observations (those in wage employment and those not) is used to estimate the correction term λ_i, which is the inverse Mill's ratio or Heckman's Lambda: $\lambda_i = \phi(x_i\ \beta_2)/\Phi(x_i\beta_2)$. This term is then included in the second stage of the augmented earnings function:

$$w_i = z_i\ \beta_1 + \sigma_{12}.\ \lambda_i + \eta_i \tag{6}$$

These two equations can also be estimated in one single step, which is more efficient (StataCorp, 2001). Identification is achieved by excluding variables (X_i) such as household size and household non-wage income from the earnings function but including them in the selection equation. The justification for this is that these affect participation probabilities, through changing the opportunity cost of being in the wage labor force, but an employer is unlikely to pay a different wage rate depending on one's household size or non-labor income.

3.2 Correcting for Endogenous Choice of Employment Sector

Workers in the state sector obtain many favorable conditions such as job stability, full labor insurance, pension benefits and possibly a wage premium over private sector workers. On the other hand, they may face a more compressed wage distribution than private sector workers (Johnson and Chow, 1997). While it is possible to include an employment sector variable in the earnings equation to allow for these differences, the model must recognize that workers may self-select into the state sector rather than being randomly allocated. Consequently, an equation should also be estimated for choice of employment sector and if the error term of the sector placement equation is correlated with the error term of the wage equation this endogenous placement (which can be considered as a "treatment effect") may cause bias in OLS estimated coefficients of the wage equation.

To overcome this problem, a model for endogenous treatment effects is used as follows:

$$w_i = z_i\ \beta + k_i\psi + u_i \tag{7}$$

where z_i is a vector of schooling and experience variables of individual i, and coefficient ψ shows the effect on wages of whether or not the treatment is applied i.e. whether one works in the state sector (k=1) or not (k=0). The sector choice depends on a latent variable, $k_i{}^*$ with the selection equation as follows:

$$k_i{}^* = y_i\ \alpha + e_i \tag{8}$$

$k_i = 1$ if $k_i{}^* > 0$ ($k_i = 1$ if observations are in state sector)
$k_i = 0$ if $k_i{}^* \leq 0$ ($k_i = 0$ if observations are in non-state sector)
and $u_i \sim NID(0, \sigma^2)$, $e_i \sim N(0,1)$, and $cov(u_i, e_i) = \rho$

In the first stage of estimation, a binary Probit model for wage earners is employed to estimate the determinants of working in the state sector and to derive the correction term or Mills' ratio. The Mills' ratio is then included in the augmented earnings equation. In contrast to the sample selectivity correction, there is a Mills' ratio for both the treated and untreated observations. Identification is achieved by including a variable in the sector choice equation which is not in the earnings function. Previous studies have used the share of the household workers (excluding the *i*th individual) who are in the state sector as the identifying variable (Gibson and Fatai, 2006), since some personal contacts may be needed to get one of these preferred jobs, and the same approach is used here.

4. Results

The descriptive statistics show that the educational attainment of wage earners in Vietnam averages about 9 years, and by 2004 was slightly higher for females than males (Appendix A). This gender pattern reflects the fact that female wage earners are a more selected sample than are male wage earners (27 percent of the working age females are in wage labor versus 42 percent of the males), so amongst non-wage earners women's education is lower than men's. The average hourly wage rate was 2,513 dong (US$0.18) in 1998 and in nominal terms had risen to 4,566 dong (US$0.29) by 2004. The average wage rates of women are lower than for men, but were rising faster, from 77 percent of the male average in 1998 to 91 percent by 2004.

Table 1 contains the basic earnings function estimates. All of the coefficients are significant at the one percent level but the explanatory power of the model is substantially higher in 2004 than in 1998. The coefficient on years of schooling implies an average private rate of return to an additional year of education of 2.7 percent in 1998 rising to 8.4 percent in 2004, for the model with males and females pooled together. These results are in sharp contrast to those of Liu (2006) who found static or declining rates of return between 1992 and 1998. However the results are quite consistent with what is found in other transition economies (e.g., Zhang et al., 2005) report a seven percentage point rise in the rate of return to schooling in China from four percent in 1988 to ten percent in 2001).

The rise over time in the rate of return to education is shared by both men and women. Specifically, the rate of return to an extra year of schooling rose by 5.4 percentage points for males (from 2.2 to 7.6 percent) and by 6.6 percentage points for women (from 3.6 to 10.2 percent). However, even though the rate of return is higher and rose faster for women, their *level* of returns is still lower than for men. According to the coefficient on the male intercept dummy variable in the first column of Table 1, hourly wages were 14.5 percent higher for men than for similarly educated and experienced women in 1998.[1] This percentage gap was almost unchanged in 2004.

[1] The percentage is calculated for dummy variables in a semi-logarithmic regression as 100 x ($e^{\beta i}$ -1).

Table 1: Basic Mincerian Earning Function Estimates

Variable	1998			2004		
	All	Male	Female	All	Male	Female
Years of schooling	0.0274	0.022	0.036	0.0864	0.076	0.102
	(7.60)**	(4.35)**	(7.69)**	(35.52)**	(24.14)**	(26.74)**
Years of experience	0.015	0.0118	0.017	0.0245	0.024	0.025
	(3.95)**	(2.24)*	(3.38)**	(8.03)**	(6.22)**	(4.89)**
Experience squared	-0.00055	-0.00056	-0.0005	-0.0005	-0.0005	-0.0004
	(5.7)**	(5.65)**	(3.68)**	(6.11)**	(5.06)**	(3.16)**
Sex (male=1)	0.136			0.134		
	(4.66)**			(7.05)**		
Constant	0.291	0.527	0.148	0.223	0.464	0.049
	(6.97)**	(9.61)**	(2.99)**	(6.70)**	(11.36)**	(1.00)
R^2	0.0526	0.0448	0.0684	0.1825	0.1497	0.2267
F-statistics	38.91**	23.2**	36.0**	356.35**	213.5**	262.6**
No of obs	3,365	2,009	1,356	7,177	4,499	2,678

Notes:
Robust *t*-statistics in parentheses, statistically significant at 10% (+), at 5% (*), and at 1% (**); dependent variable: hourly wage in log, hourly wage is measured in VND 1,000 (and for all Tables hereafter).

4.1 Changes in Returns to Educational Levels

The basic Mincerian earnings function is convenient because the returns to schooling are a single parameter but it is also interesting to consider an extended earnings function where each educational level has a separate parameter. These extended results are reported in Table 2 with five educational levels specified: primary, lower secondary, upper secondary, vocational and university. The excluded base category is those with no education.

It is apparent that the wage gaps between higher education groups and the group with no education have widened over time. In 1998, those whose highest education was upper secondary were rewarded 21.3 percent [=100 x ($e^{0.193}$ -1)] more than the no-schooling group, but by 2004 this gap had expanded to 86 percent. An even greater rise is apparent for university graduates, whose hourly wage was 88 percent more than the base group in 1998 but 245 percent more in 2004.

Table 2: The Mincerian earning function estimations for educational level

Variable	1998			2004		
	All	Male	Female	All	Male	Female
Years of	0.018	0.015	0.019	0.025	0.024	0.025
experience	(4.76)**	(2.81)**	(3.74)**	(8.36)**	(6.49)**	(5.11)**
Experience	-0.0006	-0.0006	-0.0005	-0.0004	-0.00046	-0.0004
squared	(6.6)**	(4.9)**	(4.1)**	(5.6)**	(4.7)**	(2.9)**
Sex	0.148			0.150		
	(5.11)**			(8.11)**		
Primary	0.071	-0.021	0.202	0.160	0.114	0.235
	(1.37)	(0.28)	(2.89)**	(2.05)*	(1.10)	(2.02)*
Lower	-0.021	-0.123	0.151	0.159	0.121	0.215
secondary	(0.40)	(1.77)+	(2.01)*	(2.03)*	(1.16)	(1.84)+
Upper	0.193	0.037	0.448	0.620	0.473	0.844
secondary	(3.35)**	(0.48)	(5.59)**	(7.84)**	(4.51)**	(7.18)**
Vocational	0.106	-0.079	0.359	1.081	0.998	1.206
	(1.40)	(-0.70)	(3.79)**	(12.63)**	(8.65)**	(9.68)**
University	0.631	0.626	0.692	1.239	1.133	1.424
	(9.18)**	(6.36)**	(7.52)**	(15.26)**	(10.49)**	(11.85)**
Constant	0.413	0.722	0.196	0.588	0.827	0.443
	(7.58)**	(9.54)**	(2.91)**	(7.38)**	(7.82)**	(3.77)**
R^2	0.0790	0.0794	0.0873	0.2339	0.2008	0.2827
F-statistics	29.6**	18.2**	19.08**	270.3**	154.2**	165.3**
No of obs	3,365	2,009	1,365	7,177	4,499	2,678

Notes: See Table 1.

There was also a very substantial rise in returns to vocational education. In 1998 a vocational graduate only earned an hourly wage 11.2 percent higher (and statistically insignificantly so) than a similarly experienced person of the same gender who had no educational qualifications. But by 2004 the coefficient on the dummy variable for vocational qualifications is highly statistically significant and implies an average wage which is 195 percent higher than the base group. This strong rise in the returns to vocational education mirrors a finding by Zhang et al (2005) for China, who explain it by the FDI inflow and subsequent tradable goods expansion which led to a high demand for skilled labor. Similarly, the rising returns to vocational training in Vietnam may partly reflect the need for skilled employees working in factories, given the role of FDI and exports in Vietnam's transition to a market economy.

The returns to each level of education can also be used to derive rates of return per year, under assumptions about the number of years of study needed to progress from one education level to the next. In 1998, these percentage rates of return were 1.47, -0.23, 1.77, 0.80 and 5.50 for primary, lower secondary, upper secondary, vocational and university education respectively. But by 2004 there was an across-the-board rise, to percentage rates of return of 3.5, 1.9, 7.2, 13.9 and 15.3. Except for the dip between primary and lower secondary, the patterns confirm the stylized fact in the literature that rates of return increase with the education level (Gibson and Fatai, 2006; Todaro and Smith, 2006).

The other apparent feature from Table 2 is the higher rates of return to females than to males. At all five levels of education specified, and for both years, the coefficients are higher for the female sub-sample than the male sub-sample. However, while it appears that family investment in female schooling may yield much higher rates of return than a similar investment in male children such an inference must be balanced against the much greater selectivity of the sample of female wage earners. Since only a small fraction of working-age women are in the wage-earning labor force, inferences based on such a selected sample may be misleading.

4.2 Changes in Selectivity-Corrected Returns to Education

After correcting for sample selection bias the estimated rates of return to education are somewhat higher than in the OLS estimates reported in Tables 1 and 2. However the basic feature of the data, of a significant rise in the rates of return between 1998 and 2004, is not altered. The full results of using the Heckman (1979) selection-correction are reported in Table 3, alternately measuring education by years of schooling or with five dummy variables for the highest level of education. These estimates are made on samples which pool males and females, since the maximum likelihood estimator of the selection correction model requires large samples. Hence the results in Table 3 are most comparable with the results in the “All” columns of Tables 1 and 2.

The rise in the average private rate of return to a year of schooling is illustrated in Figure 1, for both the OLS and selectivity-corrected estimates. Both sets of estimates show a rise of over five percentage points between 1998 and 2004. The rise in the returns to various levels of education is illustrated in Figure 2, with just the results from the selectivity-corrected estimates shown. The returns rise at all levels of education, with the second highest rise, of over ten percentage points, at university level and the highest rise, of 15 percentage points, at vocational level.

The estimates of the selection and wage equations show that the residuals are positively correlated, for both years and regardless of whether education is measured by school years or by highest level. Specifically, the coefficient on the inverse Mills’ ratio, λ varies from 0.16 to 0.29 and is always highly statistically significant. This implies a positive correlation between the selection equation errors and the wage equation errors, since $\lambda=\rho\sigma$ (and σ must be positive). In other words, individuals with a comparative advantage in entering the wage-earning labor force also have a comparative advantage in earning more than observationally similar workers. Hence the observed wage is higher than the wage that would prevail for a sample of individuals selected at random from the working-age population.

Figure 1: OLS and selection-bias corrected returns to an additional year of schooling

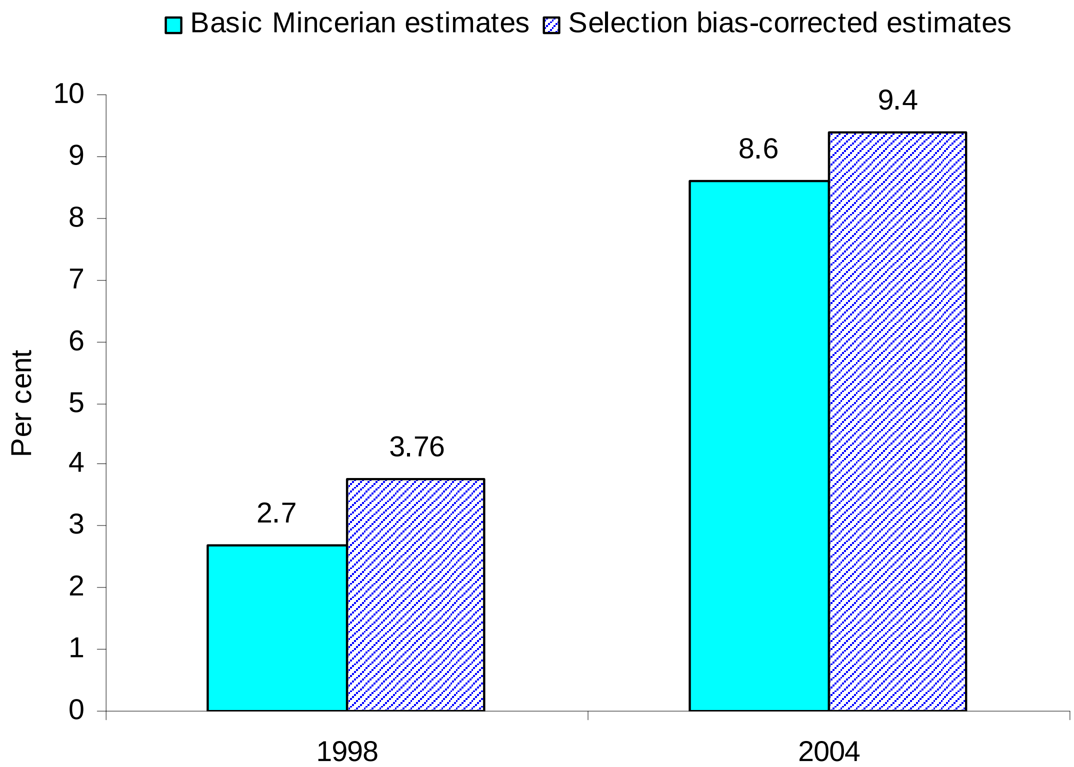

Figure 2: Selection-bias corrected average returns to educational levels

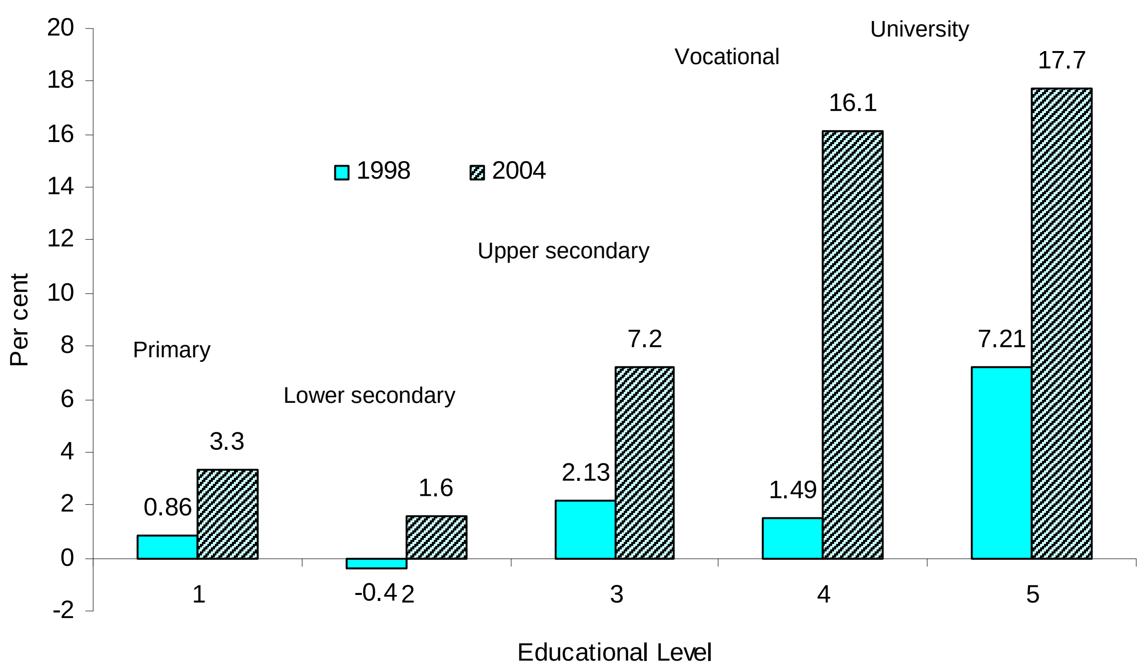

Table 3: Maximum Likelihood Estimation for Selection and Wage equation

	1998		2004		1998		2004	
Variable	Selection equation	Wage equation	Selection equation	Wage equation	Selection equation	Wage equation	Selection equation	Wage equation
School year	0.052	0.0376	0.0745	0.094				
	(14.25)**	(9.45)**	(21.56)**	(31.8)**				
Primary					-0.124	0.042	-0.049	0.151
					(2.42)*	(0.78)	(0.58)	(1.92)+
Lower					-0.037	-0.037	-0.123	0.136
secondary					(0.73)	(0.69)	(1.46)	(1.73)+
Upper					0.241	0.228	0.219	0.625
secondary					(4.29)**	(3.89)**	(2.54)*	(7.85)**
Vocational					0.500	0.189	1.345	1.178
					(7.33)**	(2.44)*	(10.47)**	(13.29)**
University					0.848	0.767	1.719	1.343
					(11.7)**	(10.6)**	(14.7)**	(15.73)**
Experience	0.00129	0.016	0.0949	0.0398	0.003	0.019	0.091	0.035
	(0.49)	(4.08)**	(31.8)**	(10.2)**	(1.16)	(4.92)**	(30.19)**	(9.61)**
Experience	-0.0003	-0.0006	-0.0025	-0.0009	-0.0003	-0.0007	-0.002	-0.0007
squared	(4.72)**	(6.33)**	(30.15)**	(8.53)**	(5.76)**	(7.10)**	(28.56)**	(7.33)**
Sex (male=1)	0.473	0.245	0.466	0.203	0.482	0.245	0.481	0.200
	(21.14)**	(7.60)**	(22.3)**	(9.13)**	(21.49)**	(7.66)**	(22.55)**	(9.52)**
Household	-0.0176		0.035		-0.016		0.037	
size	(2.81)**		(5.20)**		(2.54)*		(5.25)**	
Non-wage	-0.0735		-0.182		-0.0778		-0.196	
income[(a)]	(7.69)**		(14.89)**		(7.92)**		(15.13)**	
Constant	-0.436	-0.282	-1.61	-0.184	-1.060	-0.016	-0.995	0.349
	(25.28)**	(3.39)**	(28.99)**	(2.45)*	(16.21)**	(0.20)	(10.48)**	(3.66)**
lambda (λ)		0.289		0.219		0.254		0.161
		(8.51)**		(6.07)**		(6.95)**		(4.62)**
Wald χ^2 (all slopes=0)		203.4**		1051.8**		276.86**		1282.3**
Selectivity test: ρ=0		$\chi^2(1)$ = 67.93**		$\chi^2(1)$ = 36.49**		$\chi^{2(}1)$=48.69**		$\chi^2(1)$=21.46**
No of obs	22,591	3,365	20,549	7,177	22,591	3,365	20,549	7,177

Notes:
Dependent variable of the selection equation (observed wage earner =1, non-wage earner=0), dependent variable of wage equation is hourly wage in log. The (absolute) *z*-statistics in parentheses, statistically significant at 10% (+), at 5% (*), and at 1% (**); (a) Total household non-wage income divided by 10,000.

The positive coefficients on either school years or on the dummy variables for higher levels of education in the selection equations show a benefit of education which is omitted from standard wage equations, which is that it provides a higher probability of entering into wage work. To help interpret this effect, the probit coefficients from the selection equation are transformed into marginal effects, showing the change in probability of being in wage employment for a unit change in the explanatory variable. These equations are also estimated on separate sub-samples for males and females, with the results displayed in Table 4.

Table 4: Marginal effects of characteristics on probability of wage employment

Variable	1998			2004		
	All	Male	Female	All	Male	Female
School year	0.0104	0.0102	0.0098	0.0260	0.0210	0.0289
	(13.79)**	(7.67)**	(11.22)**	(20.98)**	(11.75)**	(17.25)**
Experience	-7.81e-06	0.0071	-0.0041	0.0350	0.0484	0.0214
	(0.01)	(6.89)**	(7.22)**	(32.49)**	(31.14)**	(15.04)**
Exp squared	-.00005	-0.00021	0.00004	-0.0009	-0.0012	-0.0006
	(4.52)**	(9.08)**	(3.04)**	(30.70)**	(28.13)**	(15.30)**
Sex[(a)]	0.1007			0.1672		
	(20.68)**			(22.41)**		
Household size	-0.0024	-0.0050	-0.0009	0.0146	0.0089	0.0161
	(1.81)+	(2.18)*	(0.59)	(5.89)**	(2.59)**	(4.75)**
Non_wage income/10,000	-0.0139	-0.0246	-0.0073	-0.0651	-0.0845	-0.0468
	(7.35)**	(7.51)**	(3.91)**	(14.52)**	(14.46)**	(8.38)**
Wald chi2	923.06**	255.52**	350.04**	1789.4**	1231.7**	541.2**
Prob > chi2	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Probability of being wage-earner at x-bar	0.1314	0.1989	0.0917	0.3405	0.4184	0.2642
Observations	22,815	9,689	13,126	20,852	10,778	10,074

Notes: Robust z-statistics in parentheses, statistically significant at 10% (+), at 5% (*) and at 1% (**), ([a]) dy/dx is for discrete change of dummy variable from 0 to 1.

There appears to be a substantial rise over time in the impact of education on the probability of being in wage employment. In 1998 each additional year of schooling raised the probability by just over one percentage point (averaged over males and females). But by 2004 the marginal effect of an extra year of schooling had risen to 2.6 percentage points. In other words, people with higher education have a higher likelihood of having wage jobs and the strength of this relationship increased over time. It is also notable that in both years the marginal effect of education was higher for females than for males, which is consistent with the pattern of female wage workers being a more educationally selected group than the male wage workers.

The final notable result from Table 4 is the overall rise in the predicted probability of wage employment for an individual with average characteristics, increasing from 13 percent in 1998 to 34 percent in 2004. This rapid increase in the probability of being in wage work is likely due to the rapid industrialization in Vietnam's urban and peri-urban areas which have helped absorb surplus labor from rural areas. The increasing significance of household size as a positive predictor of wage employment is also consistent with this surplus labor interpretation; the opportunity cost of having a household member work in wage employment is lower for a larger family since other family members can continue to work either on-farm or in some non-farm informal enterprise.

4.3 Changing returns to education with endogenous sector choice

The econometric evidence reported above on the rapid rise in the returns to schooling does not account for one possible confounding factor which is the role of the state sector. Prior to the 1990s when Vietnam was still a command economy the government set wage scales and the state budget was allocated to employees via state enterprises and administrative units. But economic reform led to public sector downsizing and private sector employment increasing rapidly. In fact by 2005 the private sector provided 88.3 percent of total employment while the state sector contributed only 11.7 percent (Duong Ngoc, 2006, p.39).

Although the state sector in Vietnam is a small contributor to employment it is a major contributor to industrial output and a major destination for investment. In 2005 the state sector received just over one-half of the total capital in the economy and contributed 39.3 percent of GDP (Ngoc Lam, 2006). Thus the state sector is capital-intensive and also enjoys privileges from monopolies in many highly profitable sectors of the economy such as mineral and oil exploitation, telecommunications, construction, airlines, transportation, and power supply. Workers in the state sector are therefore likely to be paid a wage premium since they have more capital to work with and there is the possibility of monopoly rents to be shared. Consequently, changes in the state sector wage premium and also in the changing share of the state sector in total wage employment could affect the trend in the rate of return to schooling.

To see whether the pattern of rising returns to schooling is robust, augmented Mincerian wage equations were estimated where along with schooling, experience and gender there is also a dummy variable for whether the job was in the state sector. This sector variable is treated as endogenous since workers are likely to purposively rather than randomly choose which sector to work in. The results in Table 5 support the decision to treat the employment sector as endogenous, since there is a significant (at the one percent level) correlation between the disturbances of the sector choice equation and the wage equation.

A large part of the returns to schooling appears to operate through the impact that schooling has on the sector of employment. The probit coefficients from the selection equation imply that an additional year of schooling raised the probability of wage workers being in the state sector by 0.3 percentage points in 1998 and by a much more substantial 2.8 percentage points in 2004. Since there appears to be a large state sector wage premium, of 22 percent in 1998 and 86 percent in 2004, a big part of the return to education is through the raised probability it provides of obtaining a high-paying state sector job. Once this factor is controlled for, the directly estimated rate of return to an extra year of schooling is lower than in the basic Mincerian wage equation, at 1.6 percent in 1998 and 4.7 percent in 2004. However, even these lower rates of return, which are applicable to private sector workers, represent a large rise from 1998 to 2004. In that respect, the basic pattern established from the previous results where employment sector is not accounted for still stand.

Table 5: Returns to schooling with endogenous self-selection into the state sector

Variable	1998		2004	
	Selection Equation	Wage Equation	Selection Equation	Wage Equation
School year	0.093	0.016	0.228	0.047
	(6.46)**	(3.33)**	(32.37)**	(12.08)**
Experience	0.026	0.013	0.045	0.018
	(2.43)*	(3.37)**	(7.34)**	(5.77)**
Experience squared	-0.00014	-0.0005	-0.0003	-0.0005
	(0.58)	(5.59)**	(1.76)+	(5.73)**
Sex (male=1)	0.016	0.158	-0.272	0.179
	(0.18)	(4.38)**	(7.07)**	(9.24)**
Sector (state=1)		0.199		0.622
		(7.78)**		(13.43)**
Public sector labor share	5.55		6.95	
	(17.39)**		(3.8)**	
Constant	-3.033	0.346	-4.012	0.455
	(17.59)**	(7.87)**	(17.0)**	(12.45)**
lambda (λ)		-0.092		-0.189
		(2.75)**		(7.39)**
Wald χ^2 (5) (all slopes=0)		179.82		1543.48
Rho (ρ)		-0.114		-0.269
Wald test of independent equations (ρ=0)		χ^2(1)=7.40**		χ^2(1)=52.5**
No of observations	3,365		7,177	

Note:
Dependent variable of the selection equation (wage earner in state sector =1, wage earner in non-state sector=0). Dependent variable of wage equation is log of hourly wage. The (absolute) z-statistics in parentheses, statistically significant at 10% (+), at 5% (*) and at 1% (**).

A notable feature of the estimates for the model with endogenous state sector employment is that some factors which raise the probability of working in the state sector lower the wage. For example, in 2004 a male wage worker was 24 percent less likely to be in the state sector than a female worker with the same schooling, experience and family network of state sector workers. Yet males earn a considerable wage premium over females so the selection mechanism for state sector jobs appears to go in the opposite direction to what comparative advantage would predict. This pattern is also apparent from the significant negative coefficients on the Mills' ratios in each year, which implies a negative correlation between the selection equation errors and wage equation errors and suggests that there may be some non-market aspects to state sector employment.

5. Discussion and Conclusions

The results reported in this paper on returns to schooling in Vietnam using VLSS1998 and VLSS2004 differ very substantially from previous literature. The rates of return for 2004 are much higher than estimates for either 1992 or 1998 reported by Glewwe and Patrios (1998), Gallup (2002), Moock et al (2003) and Liu (2006). Moreover, while Liu (2006) found a declining trend between 1992 and 1998 in the rates of return for men and essentially no trend for women, the current results show a very rapid rise in the rates of return to schooling between 1998 and 2004. The increased rates of return are especially marked for women and for workers with post-secondary (i.e., vocational or university) level education. The trend of increased rates of return appears to be robust to both forms of self-selection considered, that of selection into the wage labor force and selection into a state sector job.

What could account for such a rapid rise in the rates of return to schooling, especially given the previously sluggish change reported in the literature? The period studied here coincides with further opening by Vietnam and integration into the global economy, deeper reforms, and a consequent investment boom with accelerated structural change that has generated many technical-skilled jobs. Investment grew dramatically, from 32 percent of GDP in 1998 to 38.4 percent in 2004, with almost all of this investment into industry and services (Vietnam Economic Times, 2007). Consequently, the contribution of industry to GDP rose from 32.5 percent in 1998 to 40.2 percent in 2004 while low-productivity agriculture reduced its role in GDP from 26 percent down to 22 percent (Vietnam Economic Times, 2006, p.64). There was also considerable growth in foreign trade, such that overall openness (the ratio of exports plus imports to GDP) reached 127 percent by 2004.

In terms of reforms, changes in labor market laws from the early 1990s were having increasing effects in the period studied. Initial reforms in 1993 to the labor contract system introduced the "basic wage" as the minimum wage. But employers often relied on the basic wage to compute actual wages for employees without concern for appropriate differentials for educational attainment, skills and productivity. Further impetus for negotiating and signing employment contracts came in 1994 when the Labor Code was passed, allowing employers more flexibility in hiring and firing workers. This greater flexibility is also likely to have offered greater mobility for workers, allowing the more highly schooled to seek out jobs which paid appropriate wage differentials for their skills. Moreover, these more educated workers may have benefitted from the transfers of technical and managerial skills that the boom in FDI is likely to have caused.

It is unlikely that the forces behind the rapid rise in the returns to schooling in Vietnam had reached their peak by 2004. Vietnam joined the WTO in 2006 with a commitment to further opening of markets, including the labor market, so growing competition between employers is likely to affect the returns to schooling. In this regard it is notable that even though the largest rise in returns to education between 1998 and 2004 was for those with post-secondary vocational qualifications, there may still be insufficient supply of graduates in this skill category. According to Ngoc Dao (2007, p.21-23) the structure of education and training in Vietnam has been inappropriate, and Vietnam is now facing the problem of an "excess of teachers-university graduates *but* shortage of skilled-workers". Indeed, less than one-third of the tertiary educated wage earners in the VLSS2004 sample were technical and vocational graduates, yet this may be the group who benefit the most from the boom in foreign investment and the rising demand for technically skilled workers. Hence a continued rise in the rates of return to schooling is likely, making Vietnam less of a contrary case to the pattern observed in other transition economies.

Appendix A

Mean and standard deviation of variables of wage earner sub-sample

Variable	All		Male		Female	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
VLSS 1998						
Hourly wage rate	2.513	2.412	2.771	2.642	2.147	1.987
Schooling years	9.070	3.907	9.097	3.800	9.032	4.054
Experience	15.732	11.553	16.852	11.650	14.142	11.227
Age (year)	32.167	11.992	33.253	12.131	30.625	11.622
Non-wage earners' schooling years	7.642	3.604	8.121	3.533	7.330	3.616
No of wage earners	3,365		2,009		1,356	
Fraction of working age who are wage earners	14.9%		20.9%		10.4%	
VLSS 2004						
Hourly wage rate	4.566	4.351	4.726	4.480	4.302	4.115
Schooling years	9.065	3.581	8.902	3.534	9.337	3.642
Experience	17.155	10.744	17.951	10.742	15.840	10.620
Age (year)	33.567	10.803	34.242	10.850	32.451	10.634
Non-wage earners' schooling years	8.220	2.857	8.396	2.848	8.071	2.856
No of wage earners	7,177		4,499		2,678	
Fraction of working age who are wage earners	34.9%		42.4%		27.0%	

Sources: VLSS1998 and VLSS2004. Hourly wage rates are in 1,000 Vietnam Dong, and in 1998 the average exchange rate was 13,765 Dong/USD and 15,676 Dong/USD in 2004.

References

Brainerd, E. (1998). Winners and losers in Russia's economic transition. *American Economic Review.* 88(5), 1094-1116.

Chase, R.S. (1998). Markets for communist human capital: Return to education and experience in Czech Republic and Slovakia. *Industrial and Labor Relations Review.* 51(3), 401-423.

Chow, G.C (2001). *China's Economic Transformation.* Malden, MA: Blackwell.

Duong Ngoc. (2006). Labour - An outstanding issue. *Vietnam Economic Times.*

Gallup, J.L. (2002). The wage labor market and inequality in Vietnam in the 1990s. The World Bank.

Gibson, J., and Fatai, O.K. (2006). Subsidies, selectivity and the returns to education in urban Papua New Guinea. *Economics of Education Review.* 25, 133-146.

Glewwe, P., and Patrinos, H. A. (1998). The role of the private sector in education in Vietnam. *Living Standards Measurement Study.* Working Paper No 132. The World Bank.

GSO. (2004). Results of Vietnam Household Living Standard Survey 2002", *Vietnam General Statistical Office.* Hanoi.

Heckman, J. (1979). Sample selection as a specification error. *Econometrica.* 47(1), 153-162.

Heckman, J., and Li, X. (2004). Selection bias, comparative advantage and heterogeneous returns to education: Evidence from China in 2000. *Pacific Economic Review.* 9(3), 155-171.

Johnson, E.N., and Chow, G.C. (1997). Rates of return to schooling in China. *Pacific Economic Review*. 2(2), 101-113.

Liu, A.Y.C. (2006). Changing wage structure and education in Vietnam, 1992-1998: The roles of demand. *Economics of Transition.* 14(4), 681-706.

Moock, P., Patrinos, H., and Venkataraman. (2003). Education and earnings in a transition economy: the case of Vietnam. *Economics of Education Review.* 22, 503-510.

Ngoc Dao. (2007). Economic structure in 2006: Structural change results and issues. *Vietnam Economic Times*.

Ngoc Lam. (2006). Economic structure and structural changes. *Vietnam Economic Times*.

Orazem, P. and Vodopivec, M. (1995). Winners and losers in transition: Returns to education, experience and gender in Slovenia. *World Bank Economic Review.* 9(2), 201-230.

Psacharopoulos, G. (1994). Returns to investment in education: A global update further update', *World Development,* 22(9), 1325-1343.

Psacharopoulos, G., and Patrinos, H.A. (2004). Returns to investment in education: A further update. *Education Economics,* 12(2).

StataCorp. (2001). *Stata use's Guide Release 7.* College Station, Texas: Stata Press.

Todaro, M.P., and Smith, S.C. (2006). *Economic Development.* Pearson Addison-Wesley.

Vietnam Economic Times. (2006). Economy 2006-2007, Vietnam and the World.

Vietnam Economic Times. (2007). Economy 2007-2008, Vietnam and the World.

VLSS. (1998 and 2004). Vietnam Household Living Standard Survey: Datasets and Results. *General Statistic Office*, Hanoi, Vietnam.

Zhang, J., Zhao, Y., Park, A., and Song, X. (2005). Economic returns to schooling in urban China, 1988 to 2001. *Journal of Comparative Economics.* 33, 730-752.