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Kazuya Masuda and Yoko Sakai

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Institute of Economic Research
Hitotsubashi University
2-1 Naka, Kunitachi, Tokyo, 186-8603 JAPAN
http://cei.ier.hit-u.ac.jp/English/index.html
Tel:+81-42-580-8405/Fax:+81-42-580-8333

Secondary education and international labor mobility: Evidence from the free secondary education reform in the Philippines

Kazuya Masuda[a] and Yoko Sakai[b]

[a]Institute of Economic Research, Hitotsubashi University, 2-1 Naka, Kunitachi, Tokyo 186-8603, Japan.

Tel.: +81-4 42-580-8312; Email: masuda@ier.hit-u.ac.jp

[b]Independent researcher. Email: sakaiiyoko@gmail.com

Abstract

International labor mobility is a key factor for a well-functioning labor market. Although educational attainment is known to affect regional labor mobility within a country, evidence of a relationship between schooling and international labor mobility is limited, particularly in developing countries. This study uses the across-cohort variation in the exposure to the 1988 free secondary education reform in the Philippines to examine the impact of years of education on the propensity of working abroad. The results suggest that attaining another year of schooling increases the likelihood of working abroad by 3 and 8 percentage points for men and women, respectively. These results suggest that education improves the ability to deal with negative economic shocks by allowing individuals to find employment in the international labor market.

Keywords: Labor mobility, Migration, Education, Philippines, Free secondary education

JEL Classification Codes: J61, R23, I20

1. Introduction

International labor mobility is a key factor for a well-functioning labor market.[1] While a growing literature has shown that extra educational attainment causally affects regional labor mobility in the United States and in European countries, evidence on international labor migration is surprisingly limited. Understanding whether and how education affects the international labor mobility of individuals in developing countries is particularly of interest because international migration and remittance has been shown to reduce poverty (e.g., Adams and Page, 2005) and increase household investment in children's education (e.g., Edwards and Ureta, 2003) in the migrant-sending country. Since these regions have been experiencing dramatic educational expansion, the question of how education changes international labor mobility will influence labor market development in sending and hosting countries. The present study, in the context of a developing country, thus examines whether individuals become more or less likely to work abroad as they are educated.

The direction of the effects of education on labor mobility is theoretically ambiguous and empirically not straightforward. The human capital framework argues that an individual moves if the benefit from migration outweighs the cost, and generally this suggests that education increases international labor mobility. Behind this prediction are three main mechanisms: attaining an extra year of education 1) improves the skills and knowledge that help individuals find alternative employment in a distinct labor market, 2) mitigates credit constraints by increasing borrowing credibility, and 3) reduces the psychological costs associated with international migration through a tolerance for new experiences and the acquisition of the

[1] On the one hand, this allows firms in the host country to employ labor at the lowest wage in a geographically broad international market. On the other hand, unemployed workers in the sending country may find employment opportunities in another country.

language spoken in the host country. On the contrary, education may reduce the propensity of an individual to work abroad. McHenry (2013) argues that additional schooling at a low level enhances local labor market contracts and increases the opportunity cost of moving.

Early economics studies often used cross-sectional data, reporting a positive association between individual educational attainment and the frequency of regional migration (e.g., Greenwood, 1997). Identifying the causal effects of education on labor mobility is, however, challenging because of the non-random assignment of individual educational attainment. If, ideally, educational attainment is randomly assigned to each individual, OLS would provide a consistent estimate of the causal effects of education on the outcomes of interest. In reality, however, individuals endogenously decide how many years to stay in school. Thus, if unobserved factors are correlated with both educational attainment and the decision to migrate overseas, a spurious relationship between these two factors may arise.[2] Against such a background, one possible solution to consistently identify the causal effects of education on the propensity of working abroad is to use the exogenous variation in schooling created by education reform.

For this purpose, the present study examines the 1988 free secondary education reform in the Philippines to investigate the effects of education on international labor mobility. This reform eliminated the tuition fee of public secondary schools to expand access to secondary education. Since the reform came into effect in the academic year of 1988, there is cross-cohort discontinuity in exposure to the reform. Moreover, since the reform was announced after pivotal cohorts were born, exposure to the reform around the cutoff point is plausibly random. Indeed,

[2] For example, as unobserved parental experience in international migration positively affects both educational attainment and the preference for international migration of their offspring, the ordinary least squares (OLS) estimate will be upward biased. Other possible sources of omitted variable bias include individual innate ability, career aspirations, and parents' preference for investment.

Sakellariou (2006) studies the same reform in the Philippines to estimate the rates of returns to education, and this has been shown to disproportionately increase the educational attainment of the cohorts born in 1975 or later. By using this plausibly exogenous variation, we study the effects of education on international migration by employing the two-stage least squares (2SLS) approach. The Labor Force Survey (LFS) of 2006–2014 provides information on whether an individual works abroad as well as his/her educational attainment and birth year.

The results show that after controlling for endogeneity, attaining another year of education increases the probability of working abroad when individuals are in their 30s by 2.0 percentage points for men and 5.7 percentage points for women compared with the pre-reform mean of the outcome, equivalent to 44% and 96% increases, respectively, indicating economically large effects. The instrumental variable (IV) estimates are two-to-six times larger than the OLS estimates, suggesting that the OLS estimates reported in the literature may be downward biased. Our results are robust to using different polynomial selections and window sizes. In summary, we, for the first time, find that education improves labor mobility in the context of international migration from developing countries.

The present study contributes to the large literature on labor market returns to education. Quasi-experimental studies often use compulsory education reforms as a natural experiment and mostly report that an extra year of education is associated with higher labor market earnings.[3] Recent studies show that the labor market return to education is not only limited to higher wages,

[3] In developed countries, recent studies use compulsory education reforms as a natural experiment and report that increased education raises wages in the United Kingdom (Oreopoulos, 2006; Devereux and Hart, 2010; Dolton and Sandi, 2017), France (Grenet, 2013), and Australia (Leigh and Ryan, 2008). As for evidence from developing countries, Aydemir and Kirdar (2017) find that wage-related returns to education in Turkey are 7–8% for women and 2–2.5% for men.

but also associated with greater regional labor mobility within a country.[4] In Western countries in Europe, Machin et al. (2012) use the cross-state variation in the timing of the compulsory education law (CEL) reform to instrument individual educational attainment, finding that individuals become more likely to migrate within a country as they are educated. Similarly, Weiss (2015) uses the cross-country variation in the timing of the CEL reform in eight European countries and concludes that education makes people more mobile. Bauernschuster et al. (2014), in Germany, argue that the educated are more mobile over longer distances because they are more open to cross-cultural boundaries and migrate to places whose cultures differ from those at home. In summary, studies in Europe find a positive association between education and labor mobility.

By contrast, evidence from the United States is mixed. On the one hand, Malamud and Wozniak (2012) use the variation in college education due to draft-avoiding behavior among the Vietnam War generation and find that attaining a college degree is positively associated with the likelihood that an individual lives in a state other than his/her birth state. On the other hand, McHenry (2013) exploits the variation in education at a lower level by using the cross-state variation in the CEL reform in the United States and finds that individuals are less likely to migrate across states as they are educated. Taken together, existing evidence usually examines the effects of education on labor mobility within a country and empirical evidence is inconclusive.

[4] Another related but distinct literature on non-monetary returns to education includes its effects on health (Albouy and Lequien, 2009; Arendt, 2005; Crespo et al., 2014; Gathmann et al., 2015; Kemptner et al., 2011; Braakmann, 2011; Silles, 2008; Clark and Royer, 2013; Lleras-Muney, 2005; Behrman, 2015), fertility (Fort et al., 2016; Monstad et al., 2008; Cygan-Rehm and Maeder, 2013; Masuda and Yudhistra, 2018), crime (Brugard and Falch, 2013; Hjalmarsson et al., 2015; Machin et al., 2011; Bell et al., 2016), teenage marriage and risky sexual behavior (Masuda and Yamauchi, 2017; Keats, 2014; Adu Boahen and Chikakko, 2017), civic participation (Milligan et al., 2004), and religiosity (Hungerman, 2014; Cesur and Mocan, 2018; Mocan and Pogorelova, 2017; Masuda and Yudhistra, 2018).

The only exception is Fenoll and Kuehn (2017), who examine the association between education and international migration in Europe by using the cross-country variation in the timing of the CEL reform from 1950 to 1990. Consistent with McHenry (2013) but inconsistent with other studies, they find that an additional year of education reduces the number of individuals who migrate across country by 9%. However, to the best of the authors' knowledge, evidence on the causal effects of education on international labor mobility from developing countries is lacking. This is unfortunate because the developing world sends the largest number of international migrants. The present study, therefore, provides the first evidence on the causal effects of education on the propensity of working abroad in the context of South–North migration from a developing country.

This study also contributes to the growing literature on the effects of free secondary education programs in developing countries. By examining free secondary education reforms in Africa in the 21st century,[5] recent empirical studies find that such reforms have short-run effects on students' access to secondary education (Gajigo, 2016) with few negative effects on their academic achievement (Duflo et al., 2017; Masuda and Yamauchi, 2017; Brudevold-Newman, 2016; Blimpo et al., 2015; Garlick, 2013).[6] No study, however, examines the long-run effects of these reforms because programs in sub-Saharan Africa only began in the late 2000s. Therefore, the findings of this study for the first time show that such reform affects labor market performance after a few decades, at least in some developing countries.

[5] The reforms started in South Africa in 2007, the Gambia in 2001–2004, Uganda in 2007, Kenya in 2008, Tanzania in 2016, and Rwanda in 2007.

[6] There is one empirical study from developed countries. In West Germany, Riphahn (2003) finds that fee abolition at the secondary level increases educational attainment.

Lastly, our study contributes to the literature on the determinants of international migration. The determinants of migration have been of interest since Borjas (1987) investigated the selection of immigrants from developing countries to developed countries. A series of papers have studied the factors related to international migration such as the wage distribution and educational attainment (Borjas, 1987; Chiquiar and Hanson, 2005; Docquier and Marfouk, 2006), migration networks in the host country (McKenzie and Rapoport, 2011), and remittances (Gibson and McKenzie, 2011; Rapoport and Docquier, 2006; Yang, 2008, 2011). Nonetheless, evidence on the causal impact of education on the migration decision in developing countries is scarce. Our study thus uncovers the clear causality of educational attainment and international mobility in a developing country, suggesting that higher education drives more international migration from low-income countries.

The remainder of this paper is organized as follows. Section 2 provides background information on international migration and the 1988 education reform in the Philippines. Section 3 discusses the data. Section 4 reports the empirical strategy and Section 5 describes the results. Lastly, Section 6 concludes.

2. Background

2.1 Education and international migration in the Philippines

The Philippines is a low–medium-income country; however, compared with other low–medium-income developing countries, educational attainment in the Philippines is higher.[7]

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[7] At least 58.5% of individuals had completed their upper secondary education in 2013 (World Bank Open Data, 2018). By comparison, the rate in Thailand was 32% in 2016 and in Indonesia was 31% in 2014.

Despite having a better educated population, the unemployment and underemployment rates in the Philippines in the 21st century are high. The National Statistics Office of the Philippines reported the unemployment rate in January 2018 as 5.3% and the underemployment rate as 18%. In particular, women have fewer job opportunities than men. The labor force participation rates among 15-year-olds and over for men and women were 76.4% and 46.8% in 2017, respectively. Hence, more than half of women do not participate in the labor market. Yamauchi and Tiongco (2013) find that returns to education for women are lower than those for men and argue that women are discriminated against in the labor market in the Philippines.

On the contrary, Filipino women are known to work abroad more often than men do. Figure 1 presents the level and trend of international migration by sex from LFS data (1988–2014) in the Philippines. Since 1988, the estimated share of overseas workers has steadily increased for both men and women aged 25 years or over. In 1988, the estimated share of male and female overseas workers were two percent, and these shares increased two times more for men and three times more for women by 2014. As international migration is an important alternative for women to obtain employment, it has become common in the Philippines in the 21st century. Figure 1 depicts the level and change in average educational attainment by sex during the same period. In 1988, men and women attained about eight years of education on average; however, this steadily increased until 2014. Taken together, these data suggest that the level of education is positively associated with the propensity of working abroad.

To understand the types of employment opportunities in the host country for Filipino migrants, Table 1 shows the number of land-based overseas Filipino workers (new hires only) by occupation in 2014. The majority of migrants engage in low skilled jobs such as household service and waiting/bartending. Therefore, increased educational attainment may lead to a lower probability of working overseas. Yet, high skilled jobs such as nursing are also popular occupations among Filipinos (Carino, 1994). In the Philippines, students need to complete three years at nursing school or four years at nursing college, and then pass the national vocational qualification test to officially work as a nurse. Additionally, some host countries have their own certification systems. Even if the selection process is complicated and difficult, high skilled occupations are also popular among Filipinos. Given these mixed employment opportunities in host countries, it is empirically unclear whether individuals are more or less likely to work abroad if they are educated.

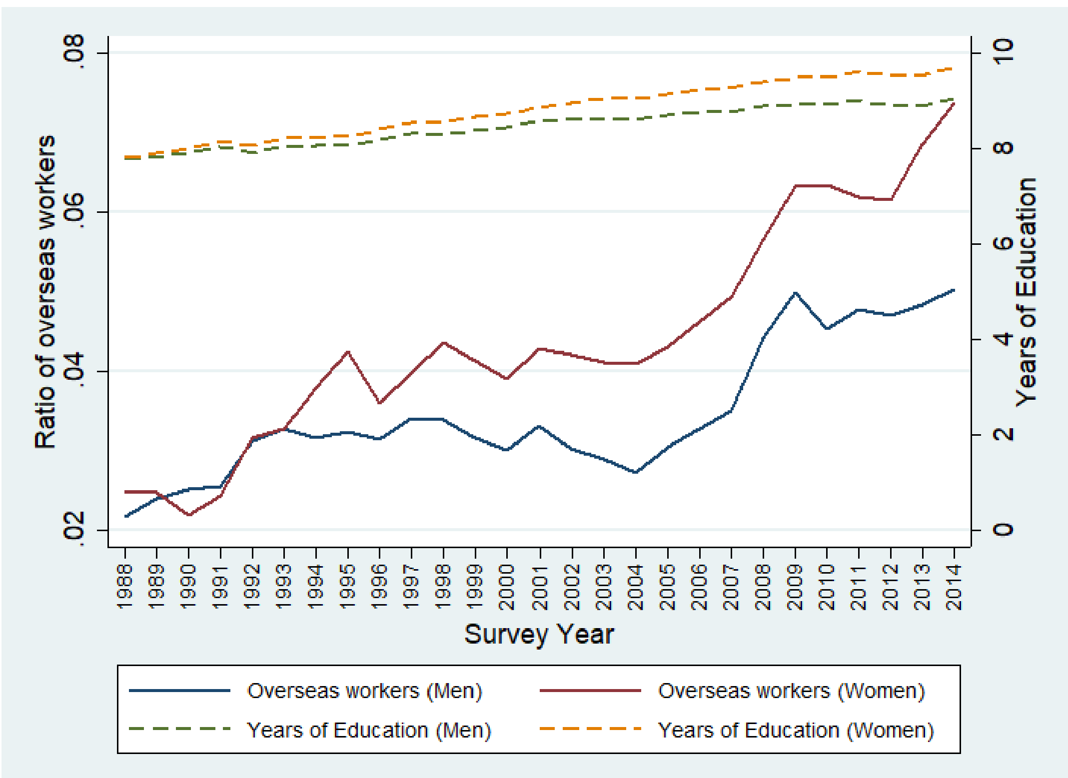

Figure 1: Change in the share of overseas workers and educational attainment in the Philippines

Table 1: Number of land-based overseas Filipino workers by occupation, new hires in 2014

Occupation	
Household Service Workers	183,101
Nurses Professional	19,815
Waiters, Bartenders and Related Workers	13,843
Caregivers and Caretakers	12,075
Charworkers, Cleaners and Related Workers	11,894
Laborers/Helpers General	11,515
Wireman and Electrical Workers	8,226
Plumbers and Pipe Fitters	7,657
Welders and Flame-Cutters	7,282
Cooks and Related Workers	5,707

Notes: The data come from the Philippines Overseas Employment Administration (2014).

The number of workers includes both men and women.

To determine whether an individual goes abroad to work, the level of direct and indirect costs plays a critical role, and these costs should be associated with educational attainment. The costs of working overseas for higher educated people should be less than the costs for less educated people because educational attainment 1) improves the skills and knowledge that help individuals find alternative employment in a distinct labor market, 2) mitigates credit constraints by increasing borrowing credibility, and 3) reduces the psychological cost associated with international migration through a tolerance for new experiences and the acquisition of the language spoken in the host country. For example, those who acquire the host country's language can more easily assimilate into the culture of that country. Fenoll and Kuehn (2017) find that students who learn foreign languages tend to migrate to countries in which those languages are spoken, presumably because their financial and psychological costs of moving abroad are much lower than those who are not familiar with other languages. Therefore, reform that promotes access to primary or secondary education may increase the probability of an individual working abroad.

2.2 Education system in the Philippines and 1988 education reform

Over our study period, the formal education system in the Philippines consisted of three levels: primary, secondary, and tertiary.[8] Primary education contains six compulsory grades and secondary school involves four years. Students are required to start the first grade of primary school at age six, and therefore they graduate secondary school at age 15 or 16. After completing secondary education, students may transit to tertiary school, which consists of college, masters, and doctoral programs.

World Bank (2018) shows that the gross enrollment ratio of secondary school has been increasing in the Philippines over the past three decades. In 1981, the enrollment ratio of women was 70% compared with 58% for men. In 2013, these had increased to 93% and 84%, respectively.[9]

[8] In the Philippines, the educational system started to shift from K-10 to K-12 in 2012. The new education system extends compulsory education and consists of six years of primary school, four years of middle school, and two years of high school. After high school, there is four years of university. This policy change does not affect our study, since our sample only includes those above 21 years in 2014 who had already completed education.

[9] In 2006, the public secondary school enrollment rate was 79.4% and the private secondary school enrollment rate was 19.9%.

In 1988, the Philippines Congress passed legislation on free secondary education including technical and vocational schools. The Free Public Secondary Education Act of 1988 declares, "It is the policy of the State to provide for a free public, secondary education to all qualified citizens." Hence, students who enroll in secondary school pay no tuition fee; however, school fees related to membership in the school community such as identification cards, student organizations, and publications needed to be paid (Free Public Secondary Education Act of 1988 (No. 6655)). This Act commenced in school year 1988/1989. In the Philippines, children usually transit to secondary school at age 12. Thus, people born in 1975 were the first cohort fully exposed to this reform.[10] Figure 2 shows the change in the average years of schooling for different birth cohorts. Consistent with Sakellariou (2006), we observe a jump in educational attainment around the cutoff point for both men and women.

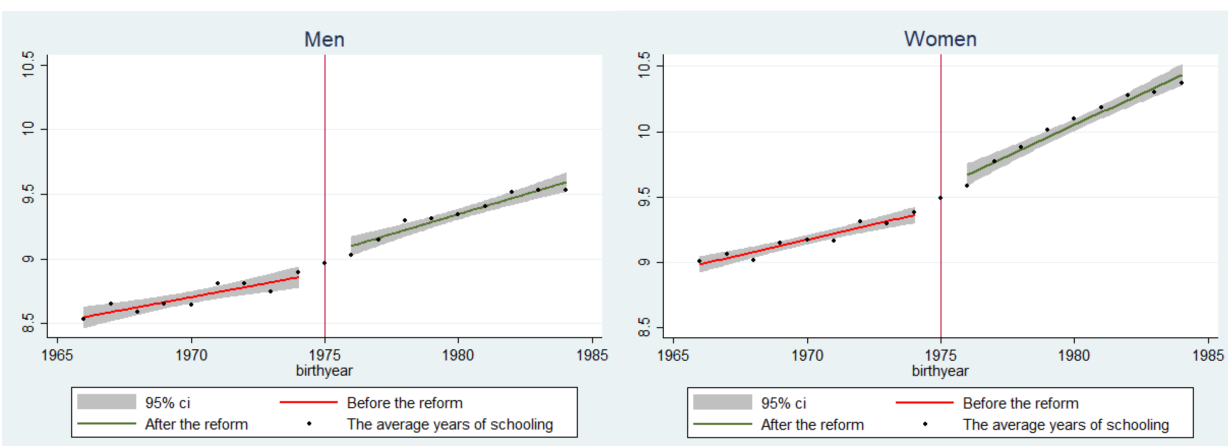

Figure 2: Average years of schooling for different birth cohorts

[10] The program abolished the school fees of secondary school entrants in 1988, and students who were already attending secondary school in 1988 had to pay tuition fees until their graduation.

The Constitution of the Republic of the Philippines (1987) under the Aquino regime mandated the right of children to receive equal educational opportunities. Article XIV, Section 2(2) explains the free provision of primary and secondary education. It, however, does not mention the exact timing of the reform implementation and thus students and their families could not have been able to expect exact timing and content. Therefore, parents could not have been able to act strategically to benefit from the change and the exposure to the reform is plausibly exogenous. Further, this reform affected only secondary school students; hence, to manipulate the running variable, parents would have needed to anticipate the reform 12 years before it was announced and delayed the conception of their children, which is unlikely in practice. Therefore, it is reasonable to assume that the exposure to the reform around the cutoff is exogenous and that the individual characteristics of the cohorts around the cutoff point are plausibly similar.

In this setting, where the treatment was assigned as if randomly around the cutoff point, we applied fuzzy regression discontinuity designs in our analysis and used the dichotomous variable for being born in 1975 or later as an instrument for the two-stage IV estimation. In addition, the estimate from our study is interpreted as the local average treatment effect (LATE), since we conducted this analysis with a two-stage IV estimation. In other words, we examined the impacts of educational attainment among the compliers who attend school longer if they are exposed to the reform, but not in the absence of the reform. Sakellariou (2006) uses the same reform to investigate the returns to education for men in the Philippines. Compared with his study, however, our main outcome is the decision to work abroad as a proxy for international labor mobility. Moreover, both men and women were affected by this reform; therefore, we included both in the analysis and run the regression separately to study the heterogeneous effects of education on international migration. In summary, the 1988 free secondary education reform in the Philippines provides an ideal setting within which to study the effects of educational attainment on labor mobility across countries for both men and women in a developing country.

3. Data

The data we used in this study came from the 2006–2014 waves of the LFS, a quarterly nationally representative repeated cross-sectional survey conducted by the National Statistics Office of the Philippines. We used the across-cohort variation in the exposure to the 1988 reform as a source of the exogenous variation in individual educational attainment to study whether education improves labor mobility. Given the year of this reform, the treatment group in the present study comprised individuals born in 1975 or after; those born in 1974 or before were considered to be the control group.[11] We therefore used a binary indicator for the cohort born in 1975 or after as the IV to predict individual educational attainment in the first-stage estimation of the 2SLS approach. Children born in 1975 became 13 years old in 1988, and this cohort is omitted from the main analysis because a part of this cohort, depending on the birth month, have had transited to secondary school, whereas the rest of this cohort were still in 6th grade in 1988 and exposed to the reform[12]. We have, however, also included them in the analysis as a robustness test, and our results are robust to their inclusion (see Appendix).

Ideally, we would measure the birth year of each individual directly from the LFS data set. However, such information is lacking from our primary data set, and thus we imputed the

[11] In the Philippines, children enter secondary school, unless they delay the primary school entry or repeat some grades, in the year in which they turn 12. We assumed that people born after 1975 were exposed to the policy change.
[12] Ideally, we use individual birth month to define the treatment cohort. However, this information are unavailable from our data set.

year of birth information from the age of the respondent at the time of the interview. Although the LFS is conducted in January, April, July, and October every year, to avoid measurement errors in our instrument, we only used data collected in January. By doing so, we are most likely to identify the age of respondents in 1988 and whether they were exposed to the free secondary education reform.

Our main outcome was a dummy variable indicating whether the individual worked abroad or in the domestic labor market at the time of the survey. As the present study is interested in the effects of education on workers' locations rather than on whether they work or not, the sample was restricted to individuals in the labor force. The main advantage of using the LFS is that it provides information on employment status in both local and international labor markets. This information allows us to identify the year of birth and educational attainment of both overseas and domestic workers in the same data set.

The data were collected by using a questionnaire-based interview; the interviewees, mostly household heads, answered the questions for all household members. An individual was defined as an overseas worker if he/she worked abroad at the time of the interview. As a result, this study defines only temporary overseas workers who have left their families behind in the original country as international migrant workers. In this sense, our outcomes of interest are different from those of other studies (e.g., Weiss, 2015; Machin et al., 2012), which define migration (or regional labor mobility) as whether a worker's current location is different from where they spent their youth. Therefore, their studies may include people who have moved to other places because of marriage or occupational changes for another household member. In addition, they mainly study the impacts of education on permanent migration. By contrast, our data consider only temporary labor migrants, which is more interesting when examining the domestic economy.

There may be some concern that examining the effects of education on permanent migration is empirically more relevant than its effects on temporal migration because the status of permanent migration may help identify the impacts of educational attainment on lifetime income. However, we argue that being a temporary migrant is relevant in the context of North–South migration from developing countries. Dustmann and Gorlach (2016) clarify that the behaviors and economic phenomena of temporary migrants differ from those of permanent migrants. For instance, migrants expecting a shorter stay in the host country send more remittances to their home country. Remittance is an important income source for supporting the family left behind, particularly in developing countries; therefore, examining the determinants of being a temporary overseas worker is at least as important as the case of permanent migrants in developing countries.

Migration studies often lack information on illegal migrants since the official data in host countries do not recognize them. However, we are likely to incorporate illegal migrants as well as legal migrants because the family member of the overseas worker living in the sending countries provides the information in the LFS. This is another advantage of the present study compared with existing studies.

To simplify the interpretation of the results, our primary explanatory variable is continuous years of education. For this purpose, ideally, LFS has a direct measure of completed years of education. One limitation of the LFS, however, is that among the 2006–2014 waves, a continuous indicator (i.e., years of schooling) is available only in the 2012 to 2014 waves. By contrast, those in 2006 to 2011 coded the highest grade completed as seven categorical levels: no grade completed, elementary undergraduate, elementary graduate, high school undergraduate, high school graduate, college undergraduate, and college graduate. To have a sufficient number of observations and high statistical power, the main analysis in this study uses data from 2006 to 2014. Therefore, we recategorize educational attainment at each education level[13] (no education, elementary, high school, and college) into a continuous measure. We test the robustness of our findings by limiting the sample to the 2012–2014 waves as well as using a dichotomous variable that takes a value of 1 if the respondent completed high school and otherwise 0. The results suggest that our estimates are robust to these alternative specifications.

Another advantage of using the LFS is that the data cover more than 130,000 respondents per wave over a decade. This setting allows us to identify the impact of a change in educational attainment on international migration after controlling for the flexible time trend (i.e., survey year fixed effects). Including the survey year fixed effects allows us to control for the unobserved year-specific factors that affect both demand for and the supply of international labor. For instance, employers in the host country may be reluctant to hire additional workers from other countries when they face financial shocks. Such across-year confounding factors that affect all the country at the same time are controlled for by using the survey fixed effects. In summary, we examine whether individuals are more likely to work abroad as they are educated within the same survey year by including the survey fixed effects.

Table 2 shows the change in the mean educational attainment and share of international migrant workers in the labor force between the treatment and control cohorts, using the 2006–2014 LFS in the Philippines. The sample is restricted to men and women aged above 21 so that

[13] These are 0 for no grade complete, 3 for under elementary, 6 for elementary complete, 8 for under high school, 10 for high school complete, 12 for under college, and 14 for college complete.

respondents have already completed their education.[14] The average education level of both men and women increased from the pre-reform to the post-reform periods. Before the reform, the share of the population who attained at least some secondary education for men was 65.9% and this increased to 75.8% thereafter. For women, the size of the change was similar (from 72.3% to 85.5%). These statistics suggest that the reform increased the educational attainment of men and women but that the effects of the reform on educational attainment, if any, may be stronger among women than men.

By contrast, the change in the proportion of overseas workers before and after the reform was larger among women than men. As shown in columns (1) and (3), the ratio of overseas workers among the labor force changed little for men. However, women's ratio of overseas workers almost doubled, from 5.5% to 9.6%, after the reform. These summary statistics imply that the effects of educational attainment on international labor mobility may be larger among women.

Figure 3 illustrates the reduced-form relationship between our instrument and propensity of working abroad. The left panel for the men's sample shows the small discontinuity in the share of respondents who work abroad around the cutoff point. By contrast, we observe a dramatic increase in the share of the female population who work abroad. The discontinuity in educational attainment around the cutoff point in Figure 2 suggests that education may have a causal impact on international labor mobility, particularly among women in the Philippines.

[14] In the Philippines, most students graduate college aged 21; however, 13.3% of students were aged 22 in 2014. We were interested in workers who have already completed their education. The LFS 2006–2014 includes a variable to identify whether the individual still attends school even after 21 years old, and this allowed us to exclude individuals still at school.

Table 2: Descriptive statistics

	Cohort born in 1974 or before		Cohort born in 1976 or after	
	Men	Women	Men	Women
Age in years	40.34	40.52	29.28	29.42
Years of education	8.52	9.29	9.23	10.58
Highest level of education (%)				
No Educ	1.91	1.74	1.25	0.96
Under elementary	16.16	11.03	12.46	6.18
Elementary	16.03	14.92	11.23	7.36
Under secondary	12.57	12.00	13.72	9.59
Secondary	28.51	27.32	29.89	28.27
Under college	12.91	12.29	15.20	15.77
College	11.91	20.71	16.27	31.88
Employment rate	91.50	91.85	86.20	81.78
Share of individuals working overseas	4.51	5.52	4.87	9.6
N	108,460	72,487	113,772	68,016

Notes: Data periods are from 2006 to 2014. The data cover individuals aged 21 to 54.

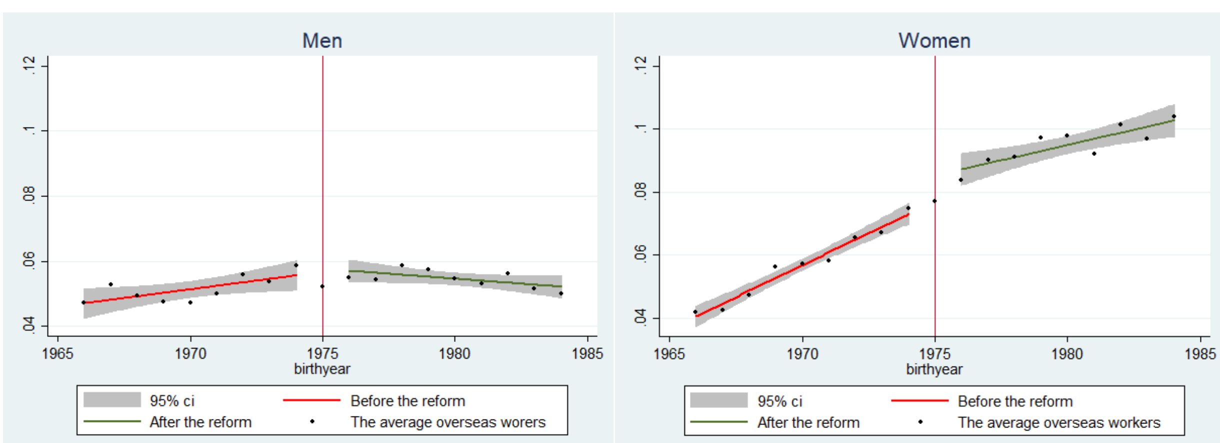

Figure 3: Change in the proportion of overseas workers

4. Empirical Strategy

Our study is interested in the extent to which educational attainment changes the propensity of working abroad in a developing country. Therefore, we consider the following causal model:

$$\mathrm{Y}_{\mathrm{ijc}} = \alpha_0 + \alpha_1 S_{ijc} + f\left(B_j\right) + \lambda_c + \epsilon_{ijc} \quad (1)$$

where $\mathrm{Y}_{\mathrm{ijc}}$ indicates the outcome variable, which takes the value of 1 if individual i, born in year j, observed in survey c works in a foreign country and 0 if he/she participates in the local labor market (including employed and unemployed people). S_{ijc} is continuous years of schooling and α_1 represents the impacts of schooling on the decision to work in another country. ϵ_{ijc} is the idiosyncratic error term. The survey year fixed effects λ_c control for the unobserved determinants of international labor mobility, which affects all the sample within the same survey year, such as the effects of the global financial crisis of the late 2000s.

The OLS estimates from equation (1), however, are likely to suffer from bias because of the endogeneity of the observed educational attainment, since this tends to be correlated with unobservable factors that affect an individual's decision to work overseas. For example, people who have high career aspirations tend to go to school and migrate to other places (Bauernschuster et al., 2014), and an unobserved factor such as this may produce the spurious positive association between educational attainment and the propensity of working abroad. To address the endogeneity of educational attainment, we thus use the across-cohort variation in the exposure to the free secondary education reform in 1988 as an instrument to identify the effect of education on the probability of working overseas.

To consistently estimate the coefficient of interest by using the 2SLS approach, an instrument should satisfy two conditions. First, the instrument, the dichotomous indicator for born in 1975 or later, should explain a sufficient degree of the variation in the endogenous variable (i.e., educational attainment). As will be shown in Section 5, the first-stage F-statistics are sufficiently larger than the rule-of-thumb crucial value of 10 proposed by Stock and Yogo (2005), thereby rejecting the null hypothesis of the weak instrument test. Thus, the first condition is most likely to be satisfied in the present study.

Second, our instrument needs to affect the decision to work overseas only through its effects on individual years of schooling. We argue that this exclusion restriction is likely to hold in the present study because the reform in 1988 was an unanticipated exogenous shock for the parents of the sample children. There may be some concern, for example, that treated mothers may manipulate the timing of conception and birth so that their children are young enough to benefit from the free secondary education reform. If such non-random assignment of the year of birth is empirically observed, the characteristics of the children, parents, and households may be systematically different between children born in 1975 and those born in 1974. This is, however, unlikely in our setting because we exploit the as-if random across-cohort variation regardless of whether children were aged 13 or older in 1988; thus, parents must have accurately anticipated in which year the reform would be implemented 13 years before it occurred. Therefore, our instrument, year of birth around the cutoff point, is unlikely to be correlated with ability, motivation, or other unobserved characteristics, and thus the estimated coefficient of educational attainment is likely to represent the effects of education on labor mobility alone. In summary, the identification assumptions on the instruments are likely to be satisfied, and we therefore use the IV method to estimate the impacts of schooling on the decision to work overseas.

Our first-stage equation of the 2SLS approach using the instrument T_j is

$$S_{ijc} = \beta_0 + \beta_1 T_j + f(B_j) + \lambda_c + \eta_{ijc} \quad (2)$$

where S_{ijc} denotes the years of schooling of individual *i* born in year *j* observed in survey *c*. The treatment group consists of those born in 1975 or after, and therefore T_j is a dummy indicating whether an individual was exposed to the program at age 13 or younger; this dummy takes 1 if the individual was born in 1975 or after and 0 otherwise. $f(B_j)$ and λ_c are polynomial functions of the birth year and survey year fixed effects. The main specification control for the quadratic function is the birth year. η_{ijc} is the error term and standard errors are robust and clustered by year of birth to allow individuals in the same cohort to correlate with others in the cohort. To investigate the heterogeneous impacts of the reform on educational attainment for men and women, we regress the equation for each sex separately. Therefore, we obtain different β_1 values for men and women, and these are our interest in equation (2).

The second-stage equation of the 2SLS is as follows:

$$\mathrm{Y_{ijc}} = \alpha_0 + \gamma_1 \widehat{S_{ijc}} + f(B_c) + \lambda_j + \epsilon_{ijc} \qquad (3)$$

where $\widehat{S_{ijc}}$ is the predicted value of S_{ijc} from equation (2). The coefficient of interest is γ_1, which we expect to be positive if an individual is more likely to work abroad as he/she is educated. To investigate the heterogeneous impacts of educational attainment on international labor mobility for men and women, we also regress equation (3) for these groups separately.

Our identification strategy using 2SLS is interpreted as a fuzzy regression discontinuity design. When adopting such a design, a bandwidth and polynomial function need to be chosen. In the presented analysis, we first apply a 15-year window (i.e., the cohort born in 1960 to those born in 1990) as a baseline. To isolate the effects of the reform on education and labor mobility from the time trend, we include a linear and a quadratic function of year of birth. We finally show several variations of bandwidth as a robustness check.

To consistently estimate the effects of the reform on educational attainment and satisfy the exclusion restriction in the second stage, another education reform around 1988 should not disproportionately affect the treated or control cohorts. Although Sakellariou (2006), who uses the same reform as a natural experiment, does not explicitly mention this point in his study, government secondary schools nationally increased dramatically from 1988 to 1990 in response to the 1987 Constitution (UNESCO, 2009).[15] This policy was implemented around the same time as the Free Public Secondary Education Act of 1988. Some may concern, therefore, that the coefficient of interest in the first stage may be overestimated because of the effects of both the 1988 free secondary education reform and the school construction program. Unlike the free secondary education reform, however, the latter policy change is likely to have affected both the cohort born in 1975 or after and that born in 1974 or earlier. Such effects, which differ across cohorts, should therefore be controlled for by using the linear and quadratic year of birth controls. Therefore, we argue that the variation in educational attainment used to estimate its effects on international labor mobility is derived from the 1988 free secondary education reform. In summary, by using the 1988 reform as a source of exogenous variation in educational attainment, we employ the 2SLS approach to identify the effects of educational attainment on

[15] Similarly, hundreds of public schools had been constructed since the start of the Aquino administration in 1986.

international labor mobility and present the results of the first- and second-stage estimations in the next section.

5. Results

Before investigating how educational attainment affects the propensity of working abroad by using the 2SLS approach, we describe the characteristics of overseas workers using the same LFS data set by estimating the following equation based on the OLS method:

$$Y_{ijc} = \gamma_0 + \gamma_1 X_{ijc} + \lambda_j + \epsilon_{ijc} \quad (3)$$

where Y_{ijc} is a binary variable taking 1 if individual *i* works in another country in survey year j. X_{ijc} includes individual characteristics such as age group, marital status, and sex and λ_j is the survey year fixed effects.[16] As shown in columns (1) to (3) in Table 3, individuals tend to work abroad in their 30s; however, compared with their 20s, they are reluctant to go abroad to work after their 40s. The coefficient of the female dummy suggests that women from the Philippines are more likely to work abroad than men. Moreover, once the sample is disaggregated by sex in columns (2) and (3), the life cycle pattern of working abroad is clearly different between men and women. For women, the 40s dummy has a significant and negative impact on their migration decisions compared with the 20s dummy. On the contrary, for men, the 40s dummy is not significantly different from zero. This finding suggests that women are less likely to work abroad once they have a family with children.

[16] The age groups are categorized as aged 21–29, 30–39, 40–49, and 50–59, and the first group serves as the reference group.

Consistent with this interpretation, the signs of the married dummy are different for men and women. For men, the married dummy has positive and significant impacts on the probability of working abroad, whereas the coefficient of the married dummy is negative and significant for women. These findings imply that during the early stage of the life cycle, women are more likely to become overseas workers than men. However, they tend to stay in their own country after they are married to take care of their families. Given these differences, we examine the effects of education on international labor mobility separately for men and women in the forthcoming sections.

Table 3: Determinants of the propensity of working abroad in the Philippines

	All (1)	Men (2)	Women (3)
1 if aged 30-39	0.013***	0.012***	0.012**
	(0.003)	(0.002)	(0.005)
1 if aged 40-49	-0.007**	0.003	-0.025***
	(0.003)	(0.003)	(0.005)
1 if aged 50-59	-0.024***	-0.009***	-0.051***
	(0.003)	(0.003)	(0.005)
1 if married	-0.009***	0.007***	-0.034***
	(0.002)	(0.002)	(0.005)
1 if female	0.018***		
	(0.004)		
Survey fixed effects	Yes	Yes	Yes
Constant	0.038***	0.023***	0.082***
	(0.003)	(0.002)	(0.005)
Observations	640,490	388,944	251,546
R-squared	0.007	0.002	0.017

Notes: The LFS 2006–2014 provides the information. All estimates are from the linear probability model. The robust standard errors clustered by year of birth are in parentheses. The sample excludes the cohort born in 1975 and is limited to the individuals aged 21 to 59 at the time of the survey.

*** p<0.01, ** p<0.05, * p<0.1

To understand whether and how much the free secondary education program increased years of education, Table 4 presents the results of the first-stage estimation for the whole sample as well as for the male and female subsamples. Our preferred specifications are in columns (3), (6), and (9), whereas in columns (1), (4), and (7), we control for the linear function of year of birth as a comparison. The coefficients of interest in columns (1), (4), and (7) are all positive and statistically different from zero, suggesting that the reform significantly increased years of schooling. The estimated regressions in columns (2), (5), and (8) include the quadratic term of year of birth, and the results are qualitatively similar, suggesting that the first-stage results are robust to controlling for the first-order and second-order polynomial functions of year of birth.

Our main results are in columns (3), (6), and (9), in which the survey year fixed effects are additionally included to control for any year-specific unobserved factors that affect the migration decision. The estimates indicate that the 1988 free secondary education policy raised years of schooling and had a similar impact on men and women (i.e., an increase of 0.35 additional years of schooling on average). There may be some concern that our findings contradict the literature, with some studies finding that educational attainment is more elastic to the tuition fee elimination among women than men because average education level in the absence of the fee elimination was typically lower for women than that of men (Gunderson and Oreopoulos, 2010). However, as shown in the previous section, women's average years of schooling in the Philippines were already higher than those of men before the reform. Therefore, it is not surprising that we find similar size effects for men and women in this context.

Table 4: First stage: The effect of free secondary education on schooling

	All			Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1 if born in 1976 or later	0.369***	0.375***	0.358***	0.363***	0.371***	0.361***	0.387***	0.378***	0.351***
	(0.087)	(0.096)	(0.092)	(0.083)	(0.088)	(0.086)	(0.120)	(0.117)	(0.11)
1 if female	1.068***	1.068***	1.071***						
	(0.081)	(0.081)	(0.081)						
Linear control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Quadratic control	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Survey year FE	No	No	Yes	No	No	Yes	No	No	Yes
Observations	505,153	505,153	505,153	308,769	308,769	308,769	196,384	196,384	196,384

Notes: Data from the 2006–2014 LFS. The robust standard errors clustered by year of birth are in parentheses. The sample excludes the cohort born in 1975 and is limited to individuals aged 21 to 59 at the time of the survey.

Table 5: IV estimates of the effects of schooling on working overseas

	All		Men		Women	
	(1) OLS	(2) IV	(3) OLS	(4) IV	(5) OLS	(6) IV
Years of schooling	0.01***	0.048***	0.011***	0.029***	0.007***	0.081***
	(0.000)	-0.007	(0.000)	(0.006)	(0.000)	(0.015)
First-stage F-statistics	-	15.3	-	17.6	-	10.2
Observations	505,153	505,153	308,769	308,769	196,384	196,384

Notes: The LFS 2006–2014 provides the information. All specifications control for the linear and quadratic functions of the year of birth and survey year fixed effects. The robust standard errors clustered by year of birth are in parentheses. The sample excludes the cohort born in 1975 and is limited to cohorts born in 1960–1990 and individuals aged 21 to 59 at the time of the survey. *** p<0.01, ** p<0.05, * p<0.1.

Table 5 presents the second-stage regression results to study how an individual's level of education affects his/her propensity of working abroad. For ease of comparison, columns (1), (3), and (5) provide the results from the OLS estimations, showing that years of schooling and propensity of international migration are positively correlated for both men and women. Column (1) shows that an additional year of schooling is associated with a 1 percentage point increase in the probability of working overseas on average. Consistent with Table 3, these results also suggest that women tend to more often go abroad to find jobs than men. Columns (3) and (5) show that another year of educational attainment is associated with a 1.1 percentage point increase for men and a 0.7 percentage point increase for women. In summary, these results suggest that consistent with the existing literature, educational attainment is positively associated with the propensity of working abroad in the Philippines. These estimates are, however, likely to suffer from bias unless the observed educational attainment is unrelated to the unobserved error term.

To consistently estimate the effects of educational attainment on the probability of working abroad, columns (2), (4), and (6) show the IV estimates, suggesting that individuals are more likely to work abroad if they are educated. Specifically, an extra year of education increases the probability of working abroad by 4.8 percentage points on average. Compared with the mean of the control cohort, this effect size is a 100% increase, suggesting the economically significant effects of education on international labor mobility. To see the heterogeneity in the effects of education, the samples are disaggregated by sex in columns (4) and (6). The results suggest that an additional year of schooling is associated with a 2.9 percentage point increase in the probability of working overseas for men and an 8.1 percentage point increase for women. These results confirm that education has heterogeneous impacts on male and female overseas workers, as shown in Figure 3. Taken together, the IV estimates suggest that even after controlling for the endogeneity of schooling, another year of education increases international labor mobility.

Theoretically, education may have positive and negative consequences on migration decisions. On the one hand, education drives overseas workers since it reduces the cost of migration such as job search costs in other countries and psychological costs. On the other hand, education lets people stay in the domestic labor market because educated people easily find jobs locally and the opportunity costs of moving increase. Our results suggest that education has a positive impact on the decision to work overseas in a developing country.

Why do individuals, especially women, decide to work overseas instead of working in their home country once they are educated? One of the possible reasons is that higher educated individuals have a higher chance of finding a well-paid job in the international labor market than the local market.[17] In particular, the wage difference for women may be larger than that for men because women suffer discrimination in the local labor market, where the returns to education for women are significantly lower than those for men (Yamauchi and Tiongco, 2013), meaning that men receive higher wages than women in the home country even though their education levels are the same. In addition, our data showed that the unemployment rates of men and women differ (Table 2). Even among individuals who obtain higher education (i.e., college graduates), women face more difficulty finding a job in the local labor market than men. Thus, educated women are motivated to work abroad instead of staying in their home country unlike educated men who choose to stay home.

Another possible reason may be the difference in the cost of working overseas by individual educational attainment. Fenoll and Kuehn (2017) show that in EU countries, being exposed to a foreign language in school increases the probability of students migrating to countries where those languages are used. In the Philippines, students start learning English in the first grade of secondary school, suggesting that attaining a secondary education promotes

[17] The Philippines Statistic Authority shows the average cash remittance per professional worker for six months in 2015 was 101,000 PHP and the average monthly wage rate for skilled workers in 2016 was 14,663 PHP. For unskilled workers, the average cash remittance per worker for six months was 41,000 PHP and the average monthly wage rate was 10,158 PHP. As remittances are part of income, it is obvious that the earnings of overseas workers are higher than those of workers in the domestic labor market, especially for high skilled workers.

international labor mobility through the acquisition of English proficiency. Other than language, the acquisition of academic and general knowledge may reduce both the financial costs and the psychological costs of moving because the speed of assimilating different cultures can be higher for people who have more knowledge of travel costs (Bauernschuster et al., 2014).

The results of the OLS and IV estimations both show that the impacts of schooling on overseas workers are positive, but the magnitudes are different. For men, the OLS (IV) estimates suggest that an additional year of schooling increases the probability of working abroad by 1 percentage point (2 percentage points). The coefficients may differ from each other for two main reasons. First, the values from the OLS estimations can suffer from omitted variable bias. In other words, unobserved factors that influence both schooling and working abroad may produce a spurious association between two variables. Our results imply that the OLS estimates are underestimated, and thus possible unobserved characteristics may negatively (positively) affect years of schooling and positively (negatively) affect the decision to work overseas. For example, parental income, unless explicitly controlled for, can be associated with both children's years of schooling and the decision to work abroad at the same time. Parental income tends to be positively associated with children's years of schooling. On the contrary, it can negatively affect migration status since children from wealthy households do not need to work abroad to send remittances home. Another possible source of bias may be the risk attitude of the individual. A risk-averse youth may stay in school longer than a risk-loving youth who tends to commit risky behavior such as teenage pregnancy. At the same time, risk-averse individuals stay in the domestic labor market for longer because working abroad is riskier owing to the lack of information on the labor markets in foreign countries. These unobserved factors may explain the underestimation of the true size effects in the OLS estimates.

Second, the IV estimates may be interpreted as LATEs, which provide the average treatment effects among compliers. In other words, the IV estimates are the average treatment effects among individuals who would not change educational attainment in the absence of the reform, but would attain an extra year of education if the public secondary school fee was eliminated. Sakellariou (2006) estimates the rates of return to schooling in the Philippines by using the same natural experiment and suggests that the subgroup affected by the 1988 reform might be high ability and have liquidity constraints. Before the reform, these individuals could complete primary school, but could not transit to secondary school even though they had high motivation and ability. Against such a background, the reform reduced the direct cost of secondary schooling, allowing those children who have high ability and liquidity constraints to continue their education.

Our IV estimates may thus provide the effects of education on international labor mobility among these compliers. Given that overseas workers from developing countries are required to complete secondary school to find employment in the host country, it is not surprising to find that our IV estimates are almost double the OLS estimates; after completing their schooling, such compliers are more likely to find high income jobs in other countries because they are qualified and highly motivated to do so.

Before concluding the present study, by way of a robustness check, we estimate equation (3) by using different windows. In Table 6,[18] we use an eight-year window (column (1)) to an 18-year window (column (11)) to assess whether our results are robust to different subsamples. The estimated coefficients and statistical significance levels are almost the same when we

[18] Tables A1–A4 in the Appendix provide the robustness checks by disaggregating the sample by sex and the results are robust.

change the window size. The first-stage F-statistics fall below 10 in the case of an 11-year or smaller window, and thus may suffer from the weak instrument problem, most likely because of the small sample size. As an additional robustness check, we also estimate the same regression equations including the cohort born in 1975. Again, consistent with our main results, the extra year of education is positively associated with the propensity of working abroad. Moreover, the coefficients of our main results and the results in Table 7 are similar, implying that our main results are robust to the inclusion of this pivotal cohort. Taken together, our results consistently show that the reform significantly and disproportionately increased the individual educational attainment of the treated cohort and that this as-if exogenously increased education is positively associated with the propensity of working abroad.

Table 6: IV estimates of the effects of schooling on working overseas with different bandwidths: All samples

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.027*	0.03**	0.034***	0.037***	0.041***	0.044***	0.046***	0.048***	0.052***	0.056***	0.059***
	(0.014)	(0.012)	(0.008)	(0.008)	(0.007)	(0.007)	(0.008)	(0.007)	(0.009)	(0.01)	(0.010)
Observations	290,261	326,553	362,105	395,130	426,391	453,528	481,235	505,153	526,510	544,156	558,989
First-stage F	3.943	5.61	7.784	10.437	12.084	14.285	14.816	15.279	15.646	15.591	16.048

Notes: Data from the 2006–2014 LFS. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The sample is limited to individuals aged 21 to 59 at the time of the survey. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).

Table 7: IV estimates of the effects of schooling on working overseas with cohorts born in 1975: All samples

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.030**	0.030***	0.034***	0.036***	0.041***	0.043***	0.045***	0.046***	0.050***	0.053***	0.055***
	(0.012)	(0.010)	(0.007)	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)	(0.007)	(0.008)	(0.009)
Observations	307,334	343,626	379,178	412,203	443,464	470,601	498,308	522,226	543,583	561,229	576,062
First-stage F	4.355	6.151	7.877	10.405	12.139	14.088	14.82	15.449	15.998	16.134	16.613

Notes: Data from the 2006–2014 LFS. Robust standard errors clustered by year of birth are in parentheses. The sample is limited to individuals aged 21 to 59 at the time of the survey. *** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).

6. Conclusion

This study examined whether individuals in developing countries become more likely to work abroad as they are educated. We exploited the across-cohort variation in exposure to the 1988 free secondary education program in the Philippines as a source of exogenous variation in educational attainment. By using LFS data from 2001 to 2014, we found that additional schooling significantly increases the propensity of working abroad among both men and women. The IV estimates suggest that an extra year of education increases the likelihood of working abroad by 2.0 percentage points for men and 5.7 percentage points for women. Compared with the preprogram mean of the outcome, these magnitudes are as large as 44% and 96% increases, respectively, suggesting the economically significant effects of education on international labor mobility. The IV estimates of the effects of education on labor mobility are larger than the OLS estimates. This finding implies that the positive association between education and migration rates reported in the empirical literature may have been largely underestimated.

The findings of this study have important implications for returns to education. Considering international job mobility as the ability for the unemployed to search for employment in the international labor market, education may improve the way of dealing with negative economic shocks in the domestic labor market. We also contribute to the literature on returns to free secondary education in developing countries by showing the long-run effects of the program on labor mobility. Given the increasing number of developing countries in sub-Saharan Africa eliminating secondary school tuition fees, this region may see greater international labor mobility in the future. In summary, the evidence of this study suggests that a program improving access to education in developing countries has large positive returns through increased labor mobility.

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Appendix

Table A1: IV estimates of the effects of schooling on working overseas with different bandwidths: Male subsample

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.011	0.017	0.020**	0.025***	0.028***	0.028***	0.029***	0.029***	0.032***	0.034***	0.036***
	(0.018)	(0.014)	(0.009)	(0.008)	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)	(0.007)
Observations	177,794	199,927	221,839	242,050	260,992	277,687	294,472	308,769	321,567	332,075	340,834
First-stage F	3.873	5.152	7.022	10.069	12.666	15.503	16.256	17.573	18.083	18.214	18.463

Notes: Data from the 2006–2014 LFS. Sample is limited to men aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses.

*** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).

Table A2: IV estimates of the effects of schooling on working overseas with different bandwidths: Female subsample

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.059**	0.051***	0.057***	0.058***	0.067***	0.071***	0.077***	0.081***	0.088***	0.093***	0.095***
	(0.024)	(0.019)	(0.013)	(0.010)	(0.013)	(0.012)	(0.014)	(0.015)	(0.017)	(0.019)	(0.019)
Observations	112,467	126,626	140,266	153,080	165,399	175,841	186,763	196,384	204,943	212,081	218,155
First-stage F	2.57	4.242	6.827	8.458	8.619	10.04	10.414	10.205	10.639	10.67	11.312

Notes: Data from the 2006–2014 LFS. Sample is limited to women aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses.

*** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).

Table A3: IV estimates of the effects of years of schooling on working overseas using 2012-14 data: Male subsample

	(1) w=8	(2) w=9	(3) w=10	(4) w=11	(5) w=12	(6) w=13	(7) w=14	(8) w=15	(9) w=16	(10) w=17	(11) w=18
Years of education	0.001	0.007	0.02	0.026	0.033	0.042*	0.044**	0.045***	0.047***	0.049***	0.05***
	(0.025)	(0.023)	(0.024)	(0.022)	(0.022)	(0.022)	(0.019)	(0.015)	(0.013)	(0.012)	(0.012)
Observations	53,477	60,252	66,930	73,719	80,344	86,898	93,725	100,431	107,181	112,081	115,663
First-stage F	15.16	14.76	9.94	11.84	12.73	11.02	11.84	13.74	12.96	14.62	15.59

Notes: Data from the 2012–2014 LFS. Years of S is the continuous indicator (i.e. years of schooling). Sample excludes the cohorts born in 1975 and is limited only to men aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column 1) to 18 (column 11).

Table A4: IV estimates of the effects of years of schooling on working overseas using 2012-14 data: Female subsample

	(1) w=8	(2) w=9	(3) w=10	(4) w=11	(5) w=12	(6) w=13	(7) w=14	(8) w=15	(9) w=16	(10) w=17	(11) w=18
Years of education	0.066	0.039	0.075	0.089	0.089*	0.115**	0.134**	0.140***	0.141***	0.143***	0.140***
	(0.078)	(0.042)	(0.056)	(0.073)	(0.047)	(0.051)	(0.056)	(0.044)	(0.039)	(0.037)	(0.032)
Observations	35,583	40,077	44,281	48,617	52,927	56,992	61,325	65,633	69,890	73,012	75,338
First-stage F	0.34	0.96	1.13	0.88	1.77	2.56	2.86	4.01	5.08	6.09	7.22

Notes: Data from the 2012–2014 LFS. Years of S is the continuous indicator (i.e. years of schooling). Sample excludes the cohorts born in 1975 and is limited only to women aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column 1) to 18 (column 11).

Table A5: IV estimates of the effects of high school completion rate on working overseas: Male subsample

	(1) w=8	(2) w=9	(3) w=10	(4) w=11	(5) w=12	(6) w=13	(7) w=14	(8) w=15	(9) w=16	(10) w=17	(11) w=18
1 if completed high school	0.131	0.199	0.231**	0.273***	0.291***	0.286***	0.291***	0.296***	0.329***	0.359***	0.394***
	(0.210)	(0.159)	(0.105)	(0.0907)	(0.0772)	(0.0677)	(0.0637)	(0.0609)	(0.0689)	(0.0806)	(0.0927)
Observations	177,794	199,927	221,839	242,050	260,992	277,687	294,472	308,769	321,567	332,075	340,834
First-stage F	1.7	2.69	4.5	7.15	9.65	12.24	13.96	15.01	14.82	14.56	14.26

Notes: Data from the 2006–2014 LFS. Sample is limited only to male. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Sample is limited only to men aged 21 to 59 at the time of the survey. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column 1) to 18 (column 11).

Table A6: IV estimates of the effects of high school completion rate on working overseas: Female subsample

	(1) w=8	(2) w=9	(3) w=10	(4) w=11	(5) w=12	(6) w=13	(7) w=14	(8) w=15	(9) w=16	(10) w=17	(11) w=18
1 if completed high school	0.520***	0.454***	0.597***	0.649***	0.765***	0.896***	0.991***	1.094***	1.213***	1.273***	1.305***
	(0.187)	(0.122)	(0.139)	(0.123)	(0.141)	(0.157)	(0.183)	(0.230)	(0.272)	(0.290)	(0.286)
Observations	112,467	126,626	140,266	153,080	165,399	175,841	186,763	196,384	204,943	212,081	218,155
First-stage F	4.12	7.16	10	11.75	11.69	13.53	13.34	10.59	10.66	10.75	11.49

Notes: Data from the 2006–2014 LFS. Sample is limited only to female. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Sample is limited only to women aged 21 to 59 at the time of the survey. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column 1) to 18 (column 11).

Table A7: IV estimates of the effects of schooling on working overseas with cohorts born in 1975: Male subsample

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.017	0.022*	0.023***	0.027***	0.029***	0.029***	0.03***	0.029***	0.031***	0.033***	0.035***
	(0.016)	(0.013)	(0.008)	(0.007)	(0.006)	(0.006)	(0.006)	(0.005)	(0.005)	(0.006)	(0.006)
Observations	188,300	210,433	232,345	252,556	271,498	288,193	304,978	319,275	332,073	342,581	351,340
First-stage F	4.403	5.922	7.312	10.243	12.763	15.244	16.074	17.428	18.127	18.467	18.792

Notes: Data from the 2006–2014 LFS. Sample is limited to men aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).

Table A8: IV estimates of the effects of schooling on working overseas with cohorts born in 1975: Female subsample

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	w=8	w=9	w=10	w=11	w=12	w=13	w=14	w=15	w=16	w=17	w=18
Years of education	0.054***	0.045***	0.053***	0.05***	0.062***	0.066***	0.072***	0.076***	0.081***	0.086***	0.088***
	(0.019)	(0.015)	(0.012)	(0.010)	(0.011)	(0.011)	(0.013)	(0.013)	(0.015)	(0.016)	(0.016)
Observations	119,034	133,193	146,833	159,647	171,966	182,408	193,330	202,951	211,510	218,648	224,722
First-stage F	2.943	4.713	7.065	8.669	9.074	10.377	10.979	10.915	11.461	11.587	12.213

Notes: Data from the 2006–2014 LFS. Sample is limited to women aged 21 to 59 at the time of the survey. Robust standard errors clustered by year of birth are in parentheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables are the quadratic function of year of birth and survey year fixed effects. The bandwidths are from 8 (column (1)) to 18 (column (11)).