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Abstract
The process of economic integration over the past two decades has been accompanied by an expanding income wedge between skilled and unskilled workers in many developing countries. This was also the case for Ugandan wage employees during the 1990s, which was a period of abrupt trade opening and market reforms. This is a surprising result for an unskilled labour abundant country like Uganda in light of a standard Heckscher-Ohlin (H-O) framework. But was the trade opening responsible for the increase in wage premia? By using a novel district-level analysis, I find that in fact increased trade reduced the returns to schooling in line with the H-O predictions. On the other hand, the intensification of domestic trade across districts during the period was associated with higher returns in those districts relatively endowed with skilled employees. This effect appears to be responsible for at least some of the rising returns to schooling among wage employees in Uganda.

Keywords: Returns to education, wage inequality, Uganda, trade, market reforms
JEL Classifications: F10, F14, F16, O12, O15

1. Introduction

The process of economic integration at the global level has been accompanied by increasing wage premia of skilled workers within many developing countries (Goldberg and Pavnick, 2007). This is a puzzling result in light of standard models of international trade à la Heckscher-Ohlin (H-O), which predict that trade liberalisation increases the rewards to the country's production factor in which it is relatively abundant (usually unskilled labour in developing countries) vis-à-vis the other factors. The literature has tried to reconcile the theory with this empirical evidence in various ways, such as by introducing the effect of skill biased technical change (SBTC), by recognising that certain developing countries are skilled abundant relatively to others (Wood, 1999), or by introducing the concept of heterogeneity across firms with higher demand for skills by exporters following trade liberalization (Verhoogen, 2008).

However these arguments are usually not applicable to many poor countries, such as most countries in sub-Saharan African (SSA), which are far from the technological frontier even after trade liberalisation and which are unskilled abundant relatively to most other developing countries. In fact the little empirical evidence available is inconclusive about the effects of trade integration on economic inequality in such countries (Meschi and Vivarelli, 2009). This paper provides new evidence on the impact of an abrupt trade liberalisation on the wage skill premium (measured by returns to schooling) in a Least Developed Country (LDC), i.e. Uganda. The timing of the trade liberalisation in Uganda (early 1990s) coincided also with a marked improvement in the relations with neighbouring Kenya, which prompted a boom in the trade between the two countries adding to the trade expansion induced by the liberalisation. The paper employs a novel identification strategy, which uses district as the unit of analysis and exploits the location of the main border-posts during the liberalisation period to predict the districts' relative exposure to trade. This strategy allows for a relatively clean identification of the effects of trade on the wage premium controlling for other potentially confounding factors (including also a proxy for technical change) and overcoming the lack of firm-level data which is common in many LDCs. The analysis confirms the prediction of the standard H-O model for an unskilled labour abundant country as Uganda: districts more exposed to international trade following liberalisation experienced smaller increases in the wage skill premium relative to the other districts.

The result is even more important considering that during the period of analysis (the 1990s) Uganda experienced an increase in the wage skill premium, which could be mistakenly interpreted as a by-product of the trade liberalisation of that period. This result is also in line with that of Amiti and Cameron (2012) for another unskilled labour abundant country (Indonesia), confirming that the standard H-O seems to be an adequate framework to think through the trade effects on inequality in such countries. As a further confirmation of the predictions of the H-O framework in this context, I also find that the increased domestic inter-district trade following the reforms in the 1990s is associated with a rise in returns to education in those districts relatively endowed with skilled labour.

The district-based empirical strategy is convenient also in that it allows focusing on the bulk of the recent rise in economic inequality in developing countries, which is within rather than between regions (Demombynes et al., 2003; Yemtsov, 2003, Elbers et al., 2005). In this way the analysis complements an expanding literature on inequality between regions within developing countries (Kanbur and Venables, 2005; the special issue of Journal of Economic Geography, 2005).

This paper is explicitly related to the literature investigating the spatial dimension of inequality. Duranton and Monastiriotis (2002) find that the different evolution of returns to education between regions explain an important fraction of the large and increasing North-South divide in the UK between 1982 and 1997. For example average earnings increased in London relatively more than in the North as returns to education in London grew faster (catching up with those in the other regions) and the share of educated labour increased relatively more quickly in London than elsewhere. However few studies go beyond the description investigating the determinants of these labour market inequalities within regions. Taylor (2006) finds that trade and technology intensity are the most relevant factors to explain within region inequality in the UK throughout the 1980s and 1990s. These results are consistent with both the theoretical predictions of the H-O model for a skilled labour-abundant Northern country and the expected effects of SBTC. He also finds that different education premia across regions are persistent over time. Such variation in inequality across space is likely to be associated with imperfect geographical mobility of labour. In fact in a context of perfect labour mobility, skilled labour would tend to move towards regions where the earnings' skill premium is higher, and unskilled labour would do the opposite, thus equalising labour market inequalities across regions. Limited geographical mobility of labour is usually more pronounced in developing countries' labour markets, and Uganda is no exception.[1]

To the best of my knowledge, only one study, by Yang (2005) on China, has analysed within-region labour market inequalities in a developing country context. He examines the determinants of returns to education across urban areas in China before and during the transition from planned to a market economy. His results reveal a large variability across cities, which is partly explained by skilled labour demand and supply related factors, such as the size of the technological and of the public sector, the presence of foreign firms, the level of infrastructure provision. Similarly to Yang (2005), I use a micro approach to analyse how returns to education in Uganda have evolved at the district level during the 1990s, but in addition my analysis explicitly tests for some of the channels through which market reforms may have influenced labour market inequalities.

The present analysis also adds to the literature investigating the evolution of returns to education during periods of market reforms in developing countries. There is no clear evidence that returns to education have evolved homogeneously in the SSA region in a time of pro-market reforms. Evidence from Ghana between 1987 and 1992 (Canagarajah and Thomas, 1997), from Tanzania (Soderbom et al., 2006) and Uganda (Appleton, 2001) in the nineties suggest the presence of rising returns to education. On the other hand, Krishnan et al. (1998) find that returns in urban Ethiopia remained stable despite labour market reforms in the early 1990s; and Soderbom et al. (2006) find high but falling returns in the Kenyan manufacturing sector over the nineties. Soderbom et al. (2006) argue that economic policies are key to account for the differences in levels and dynamics of returns to education across countries. The high returns in Kenya relative to Tanzania at the beginning of the 1990s reflected a policy environment in Kenya more favourable to the liberal operations of market forces than in Tanzania. A decade of reforms in Tanzania brought the country's polices more in line with those of Kenya. As a likely result the earnings profiles (at least in the manufacturing sector) were quite similar in the two countries at the end of the 1990s. This paper contributes to this literature by investigating explicitly the role of market reforms in the evolution of returns to education in the SSA context.

[1] According to the data used in this paper on average only around 6% of the district's labour force has changed districts between 1992 and 2000. In a typical year less than 2% of the work force change region in the UK (McCormick, 1997).

The remainder of the paper is organised as follows: the next section provides the context of the analysis describing the pro-market reforms and the evolution of the wage skill premium in Uganda in the 1990s; section 3 discusses the links between trade liberalisation and inequality; section 4 describes the data; section 5 presents the empirical approach; section 6 presents the results and section 7 concludes.

2. Uganda in the 1990s: Abrupt liberalisation and increased inequality

Uganda underwent substantial pro-market reforms during the 1990s, which have transformed the economy in a number of ways. With the help of World Bank sponsored Structural Adjustment Programme, the economy was considerably liberalised starting in 1987 with the liberalisation process intensifying during the nineties. This process included the downsizing of the public sector, the privatisation of state owned enterprises, and measures aimed at lifting constraints to trade both domestically and internationally. These measures along with improvements in transport infrastructure have led to an increase in Uganda's trade both across district and across borders. Such increases in trade are likely to have had a relevant impact on labour market inequalities, and most of the subsequent empirical analysis is devoted to testing for such impacts. Let us consider international and domestic trade in turn.

International trade liberalisation was an important component of the Ugandan reform process in the 1990s. Table 1 shows that between 1994 and 2000 Uganda halved the average tariff rate vis-à-vis the rest of the world and more than halved the one vis-à-vis the other members of the Common Market for Eastern and Southern Africa (COMESA). Tariffs were rationalised being reduced to only three non-zero rates (with a maximum of 30% rate) and almost all import bans removed by 1998. Maximum tariff rates were reduced dramatically as well. Using actual trade data Rudaheranwa (2005) calculates that the average effective rates of protection due to applied tariffs fell from 35% in 1994 to 18% in 2001. Exports were also facilitated, for example through the reduction and then the abolition (in 1996) of the tax on coffee exports (the main Ugandan export) and a duty drawback scheme for exporters was introduced.

This liberalisation was coupled by improvements in transport infrastructures and by a reduction in border-post transit times. This was achieved through an ambitious cross border initiative to promote regional trade, especially with Kenya and Tanzania, introduced in 1993. These factors contributed to a large increase in trade during the nineties. According to

COMTRADE data, imports in particular grew by 50% between 1994 and 1999 from USD 680 million to over USD 1 billion. And this figure does not likely document the magnitude of the total increase in trade between 1992 and 1999, which is the period under consideration here, as a substantial jump in international trade occurred also between 1992 and 1994. During that period trade with Kenya, Uganda's main trading partner, boomed as political relations stabilised after the two countries nearly went to war in 1992. A liberal and buoyant foreign exchange market and liberal immigration procedures by both countries enacted in those years facilitated the free flow of goods across the border.[2]

While there is data to measure the increase in Uganda's international trade, the rise in inter-district trade is more difficult to document in the absence of data on internal trade. However some pieces of evidence are consistent with a substantial increase in inter-district trade in the nineties. First, the monopoly commodity marketing boards, such as the Produce Marketing Board were dismantled in 1992. These boards *de facto* controlled trade and production by guaranteeing local agricultural producers a minimum price for their crops. Their elimination substantially freed inter-district trade in agriculture (van der Geest, 1999 p. 132).

Second, explicit restrictions to cross-district product movements were removed in 1993 (World Bank, 2008). Moreover, the process of domestic trade liberalization included also an attempt to open rural areas to markets through improvement of infrastructure. (DENIVA, 2005). This improvement contributed to a considerable reduction in overland transport costs during the nineties. Rudaheranwa (2005) calculates that the implicit taxation on Ugandan exports relating to transport costs declined from over 31% in 1994 to about 24% in 2003 for 40-foot containerized exports. This was coupled by transport policy reforms, such as the commercialisation of Ugandan Railways in the early 1990s, which improved efficiency in rail transportation and allowed the railway to compete with road transportation (Rudaheranwa, 2005).

As data on inter-district trade is not available in Uganda, one indirect way to check for the increase in this trade over the nineties is to look at the variation in prices across districts over time. I use the same data as in the rest of the analysis - i.e. the Household Surveys carried out in 1992 and 1999/2000 through the World Bank Living Standards Measurement Survey (LSMS) –

[2] Newspapers in 1994 reported that "commercial traffic is so heavy that existing customs personnel are unable to cope with it and hundreds of vehicles loaded with goods spend up to a week awaiting customs clearance" (Ojulu, 1994). And residents of the Uganda-Kenya border town of Busia defined the volume of trade across the border in that year as unprecedented in the past 20 years.

in order to compute such variation. The latter (as measured by the range/mean and the standard deviation/mean ratios) is lower in 2000 than in 1992 for commonly traded goods within Uganda as shown in Table 2. This finding is consistent with the price-converging effect of increased trade between districts during the 1990s.

These reforms have been associated with a healthy growth rate of the economy, which averaged 7.1% per year in real terms between 1992/93 and 1999/2000, comparing favourably to the 4.7% of the previous period.[3] However, in line with the experience of many developing countries which have increased their level of integration with the global economy, also Uganda has seen a rise in income inequality throughout the 1990s. During this period of intense reforms, overall inequality, as measured by the Estimated Household Income Inequality (EHII) index, grew substantially (Figure 1).[4]

This result applies also to income inequality for wage employees and self-employed alike, as found by Appleton (2001) and Calì (2009) using returns to schooling as the measure of inequality.[5] The results in the latter study suggest that the average effect of an extra year of education on employees' wages increased from 8.7% in 1992 to 14.7% in 1999-2000 and the rise in the effect of primary education is larger than that of secondary education. The breakdown by educational level shows a convex shape in the returns to education. In both periods the marginal return to post-secondary education was higher than the returns before the secondary level. This convexity is in line with findings on other Sub-Saharan African countries, e.g. Soderbom et al. (2006) for Kenya and Tanzania and Baptist and Teal (2008) for Ghana. However, unlike in the case of Tanzania, the intensity of the convexity seems to have decreased somewhat during the nineties in Uganda. The ratio of post- to pre-secondary returns shrank from a factor of around 2.4 in 1992 to a factor of around 1.4 in 2000, when the level of returns was significantly lower than in Tanzania (where it exceeded 27%). That was essentially due to a large increase in returns to primary education, which is encouraging for strategies aiming at expanding primary education

[3] Based on National Account elaborated by the IMF and the Ugandan Bureau of Statistics.

[4] The index ranges from 0 to 1 as a conventional Gini index and is built combining the Deninger and Squire (1996) inequality measure with the UTIP-UNIDO index, which is a wage inequality Theil measure for the industrial sector. The indices are developed by the University of Texas Inequality Project (UTIP) and are available at http://utip.gov.utexas.edu.

[5] Given that access to education and educational achievements are unequally distributed across the population in developing countries, and that education is one of the major determinants of wage earnings, returns to schooling capture a substantial part of a country's economic inequality. In a context of limited geographical labour mobility, such as Uganda in the 1990s, this measure of inequality tends to vary significantly across space.

provision as the one currently pursued by Uganda via UPE. To what extent the increase in the supply of (primary) educated workforce may offset this upward trend of returns to primary schooling is a question which will be tackled in the next sections.

3. Trade opening and inequality in a developing country

The rise in labour market inequality documented above appears to clash with the expected – according to the standard trade theory - labour market impact of increased trade for an unskilled labour abundant country like Uganda. The basic (2 goods, 2 factors, 2 countries) H-O framework predicts that countries specialise in the production of goods intensive in the factor which is relatively abundant at home. The Stolper-Samuelson theorem complements the H-O model by showing that trade liberalisation induces in each country an increase in the production of the good intensive in the abundant factor. Therefore freer trade would increase the demand for the relatively abundant factor of production. In a simple two-factor model with skilled and unskilled labour, the relatively abundant factor in a low income country as Uganda would be the latter. According to the Stolper-Samuelson theorem this increased demand should be reflected in higher return to unskilled labour in developing countries. Conversely, the expected decrease in the demand of the skilled-labour intensive imported products should lead to a decline in the wage of skilled labour.

The increased demand for (relatively) skilled labour associated with the rising wage premia in Uganda seems prima facie a puzzling result. To be sure this is the case for virtually all developing countries for which empirical evidence exists on the evolution of income inequality, and labour market inequality in particular, during the 1980s and 1990s (Goldberg and Pavnick, 2007). As in Uganda, these periods also coincided with a marked increase in the exposure of developing countries to international markets in terms of lowering degree of trade protection, increasing share of trade in GDP and capital flows. These findings could be reconciled with the H-O framework once its basic assumptions are relaxed in two main ways (Anderson, 2005). First, adding natural resources as a further factor of production to the model may change the balance of relative abundance in those developing countries relatively rich in natural resources. Second, if one relaxes the H-O assumption that all countries have equal access to the best available production technology, then greater openness to that technology for countries that did not access it before liberalisation may increase the relative demand for skilled labour even in low-income developing countries as better production technology is usually operated by relatively skilled labour.

However, these considerations seem to have limited application in an LDC like Uganda. In fact its abundant factor would be unskilled labour even in a three-factor model. Labour-land ratio is relatively high in Uganda, especially by African standards. In terms of population density it ranks number 49 out of 191 countries of at least 1 million inhabitants; and it ranks number 7 (out of 52) in Africa. Even the second argument has limited relevance in Uganda, as poor sub-Saharan African countries tend to have limited access to the technological frontier even after trade opening (Soderbom et al., 2006; Baptist and Teal, 2008).[6]

There are other possible alternative explanations for the increase in the wage skill premium in developing countries. For example empirical evidence suggests that usually unskilled labour intensive sectors are most protected before the liberalisation (Attanasio et al., 2004), which disproportionately affects such sectors. However the tariff data does not provide clear evidence for this pattern in Uganda, where the most protected sectors continue to be the most unskilled labour intensive sectors and tariffs have decreased substantially across all sectors. Another explanation concerns the concomitant increase in capital flows (along with liberalisation) which are usually complementary with skilled labour. Again, this does not seem to be the case in Uganda which has not received any significant amount of capital inflows following the liberalisation process.

Perhaps the most interesting alternative explanation comes from the idea that firms are in fact heterogeneous and exporters tend to be more productive and produce better quality products than non exporters (Iacovone and Javorcick, 2010). More openness to trade induces an increase in firms' productivity and/or in the quality of their product, as trade can shift resources from non-exporters to exporters (Melitz, 2003) and firms in import competing sectors overcome competition by differentiating themselves. This firms' upgrading process should induce in turn a higher demand for skill, and thus an increase in the skill premium. For example, as production for export markets tends to be relatively more skill-intensive than production for domestic markets, higher demand for exports will increase the relative demand for skilled workers. However as noted by Goldberg and Pavnick (2007) the empirical evidence on how this channel affects

[6] In any instances, in the subsequent analysis I use a technological proxy to try to isolate the pure trade effects from the potential skilled-biased technological change (SBTC)-type effects of trade opening.

inequality is still scant. Moreover in the absence of good firm-level surveys for the 1990s in Uganda, we can only speculate the extent to which this channel may be at work in Uganda. One problem with this explanation in our case is that Ugandan exports are highly concentrated in unprocessed agricultural products (coffee, fish, maize, beans), the production of which is likely to be intensive in unskilled labour even for export.

This discussion suggests that a country like Uganda is not expected ex ante to depart from the Stolper-Samelsuon prediction that increased international trade should reduce labour market inequality. This should also be the case for the impact of the likely rise in inter-district trade within the country described above. Similarly to the case of international trade, according to the Stolper-Samuelson theorem, districts relatively endowed with skilled labour are supposed to experience an increase in the relative wage of skilled employees following the increase in inter-district trade.

The subsequent district-level analysis offers one way to test for these hypotheses, employing a fairly novel identification strategy of the international trade effects on the wage premia based on the location of the districts rather than on their exposure to tariff reductions based on their production structure.

4. Data

The analysis is based on data from two Household Surveys carried out in 1992 and 1999/2000 through the LSMS. These are the only nationally representative surveys available in Uganda covering economic and social characteristics over the period of time of interest. These contain data on about 10,000 households for 1992 and 1999/2000 (2000 henceforth). In particular, after excluding pensioners and students, data is available on about 3,600 and 2,600 wage earners respectively. In cases when more than one of these individuals belongs to the same household, their labour market participation may potentially affect the decision of other members of the family to enter the labour market. I control for this by clustering the results of the analyses at the household level.[7]

The households are located in districts, which are a third layer of administrative division in Uganda (regions and sub-regions being the first two). According to the Uganda Bureau of

[7] Data is also available for about 5,600 and 5,700 crop farming enterprises in the two periods. These enterprises are run by self-employed labourers who represent most of the income earning population in Uganda. Despite their importance, I cannot include them in the analysis due to missing data on land used in the production for 1999/2000.

Statistics' (UBOS) geographical division there were 38 districts in 1992 and 45 districts in 1999/2000 as a number of districts have been split into two or more districts between 1992 and 1999. For comparability reasons I keep the division of 1992 throughout the analysis. However data is available only on 41 of the 45 districts in the 2000 survey. Due to the insecure situation in the region bordering with Sudan, the data in 2000 is not available in four districts which were included in the 1992 survey: Kitgum and Gulu (Northern region), Bundibugyo and Kasese (Western region).

Most of the district-level regressors used in the analysis are reported in the LSMS surveys at either the individual or the household level so that I can use micro information to construct labour market variables for each of the districts. Hence, these district-level factors can be treated as exogenous to individual decisions, but relevant to the determination of returns to education in local labour markets.[8]

I also construct some of the district-level variables from the part of the surveys containing information on the communities of residence of the individuals surveyed. There are 1,216 and 1,086 communities surveyed in 1992 and 2000 respectively, although not all of them are represented in the wage employees' sample I use. The surveys gather various types of information, including on infrastructure, prices, health and education. I average this community data to generate the district-level variables.

As noted by Goldberg and Pavnick (2007) there is a clear advantage in using households rather than firms' surveys for this type of analysis as the latter contains much more limited information on worker and job characteristics than in household surveys. Therefore researchers have usually to resort to the familiar dichotomy between production and non-production, or white- and blue-collar workers, which cannot capture the labour inequality concept in an entirely satisfactory way.

5. Empirical framework

The empirical strategy consists of a two stage analysis similar in spirit to an increasingly large empirical literature (see for instance Guiso et al., 2004; Mattoo et al., 2008), which uses estimated coefficients from first stage regressions as dependent variables in a second stage

[8] In fact there may be a problem of endogeneity if for instance those district-level variables determine the location of individuals or households with certain characteristics into a specific district. I try to control for this potential source of endogeneity below via instrumental variable estimation.

analysis. In the first stage I estimate standard wage regressions for employees at the district level separately for 1992 and 1999/2000. In the second stage I use returns to education from the district-level regressions as dependent variables to investigate the determinants of returns to education and the effects of pro-market reforms in the nineties. Such analysis is based on the assumption that districts define separate labour markets. This is a fair approximation in the case of Uganda as labour mobility between districts is fairly limited: according to the LSMS data inter-district migrants between 1992 and 2000 represent on average only 6% of the total district's labour force. As mentioned above this is a very low figure by international standards. The idea that districts may represent different labour markets is also supported by the statistically different returns to education across districts found in the analysis below.

5.1 First stage

For each year I run separate Mincerian-type regressions (Mincer, 1974) at the district level in order to compute the variation of returns to education across space. The basic equation in each year reads as follows:

$$wage_{ic} = \sum_{k=1}^{N} \alpha_k d_{ik} + \sum_{k=1}^{N} \beta_k d_{ik} S_i + \Pi W_i + \gamma_c + \varepsilon_i \qquad (1)$$

where *wage* is the (log of) nominal wage earned by wage earner *i* in community *c*, *S* is the number of years of formal education, N is the number of districts, d_{ik} is a dummy that takes the value of 1 if *i* resides in district *k*, and the value of zero otherwise; B_k is a column vector of coefficients associated to the interaction between the vector D_{ik} of d_{ik} dummies and that of the standard Mincerian controls W_i (i.e. gender dummy, experience and its squared term); and γ are community-level fixed effects to control for highly localised time invariant factors that may have an independent impact on wages.[9] Community fixed effects represent a severe control set given the individual data available: the employees covered in the surveys belong to 820 and 796 communities in 1992 and 2000 respectively (with only a handful of communities being the same in both surveys). The effects of the covariates in *W* are not allowed to vary across districts, thus

[9] This estimation computes a wage effect, i.e. the change in income earned by a labourer through an additional year of education (all else constant). This wage effect is technically not the same measure as returns to education, as the computation of (private) returns would require discounting the wage effects by the (private) cost of education. However the literature usually utilises the two terms interchangeably and that is also the approach followed here.

allowing to saving degrees of freedom. In any case in the empirical analysis I test the robustness of the results to allowing W to vary across districts.

One problem with the implementation of such wage regressions is that S in (1) is likely to be endogenous to returns to education as unobserved individual ability may influence both the schooling outcome and the level of income. Data availability would allow us to use only parents' education to instrument for years of education. However family background is likely to influence children's income, thus casting doubts on the exclusion restrictions of such instruments (Card, 1999). Moreover an increasingly popular view holds that OLS estimates appear to be remarkably close to the actual returns calculated using samples of twins and family background to control for ability bias (Ashenfelter and Zimmerman, 1997; Card, 1999). Given the little expected overall bias of the OLS estimates and the substantial difference with IV estimates using parents' background in Uganda (reported in Calì, 2009), I base the following analysis on OLS estimates. In addition since the eventual bias of the estimated β coefficients via OLS is not expected to differ systematically across districts, this should not be a concern in the district-level analysis.

5.2 Second Stage

In the second stage I use the β coefficients derived from (1) and their variants as dependent variables in order to identify the determinants of returns to education across districts and over time. The basic specification reads as follows:

$$\beta_{kt}^{FE} = a_k + b_1 q_{kt} + \Gamma X_{kt} + d2000 + \varepsilon_{kt} \qquad (2)$$

where β_{kt}^{FE} are the schooling coefficients estimated through (1) including community fixed effects, a_k are district fixed effects, q is the relative quantity of skilled labour in district k at time t; X is a vector of demand for skilled labour shifters, and *d2000* is a time dummy. The hypotheses are that the relative price of skilled labour would decrease in its relative supply and would increase in its relative demand (Katz and Autor, 1999).[10]

The fact that the dependent variable in (2) is estimated rather than observed should not present any difficulties for the regressions aside from a loss of efficiency, unless the sampling error in the dependent variable (i.e. the difference between the true and the estimated value of the dependent variable) is not constant across observations. In this case the regression errors will be

[10] The use of district fixed effects in the empirical specification ensures that other relatively fixed factors of production, such as land and to some extent capital, are controlled for.

heteroscedastic and OLS estimation may generate inconsistent standard errors (Lewis and Linzer, 2005). As the samples in the LSMS are stratified at the district level, the estimated β in (3) should not have sampling errors which are systematically different across districts, that is $E(\varepsilon_i \mid \mathrm{Edu}_i, W_i, Z_i) = \hat{\varepsilon}_i$, $\forall\, i$ in (2). On the other hand the large cross-district variation in the number of observations over which the β are estimated (see Table A2 in the Appendix) may represent a source of possible heteroscedasticity in the standard errors in (2). Lewis and Linzer (2005) show that the OLS method using White's (1980) heteroscedastic consistent standard errors is generally the most reliable way of estimating regressions with estimated dependent variable in the presence of heteroscedasticity. That is the case except in two instances. First, when the share of the regression residual due to sampling error in the dependent variable is very high (at least 80%) Weighted Least Square (WLS) estimation is preferred. Second, when information about the sampling errors in the dependent variable is available and highly reliable, the feasible generalized least squares estimator developed by Lewis and Linzer (2005) is a more efficient option. However neither of these cases is likely to apply here: the eventual sampling error in the dependent variable is unknown, and it is likely to be not very high (if it exists at all) as argued above. Therefore I estimate (2) via OLS using White's heteroscedastic consistent standard errors. To be on the safe side I also test the robustness of the results to excluding the district with the lowest number of observations (Kapchorwa), as this may be an important source of the dependent variable's sampling error, and to estimating (2) through WLS. As shown below the results are little affected, suggesting that the standard errors in (2) are robust to the possible concerns related to the use of an estimated dependent variable.

5.2.1 Control variables

I measure the quantity of (relatively) skilled employees (the variable q in (2)) through the average number of years of formal schooling completed by the wage earners (*employees education*).[11] The results of the analysis below are very similar when using the district's share of wage earners who have completed an education equal or higher than the primary level as the quantity of skilled employees' measure (result are available upon request).[12] This control is

[11] Table A1 in the Appendix describes the construction of the variables used in the regressions.

[12] This measure is justified as completion of primary schooling was still far from universal in Uganda in the nineties. The unweighted district-wise average share of adults that completed primary school for wage employees was 58% in 1992 and 63% in 2000).

particularly important as investment in education has been identified as a key policy lever to raise the rate of growth and poverty reduction in Uganda both by the government and by the donor community.[13] However a popular view holds that returns to schooling in poor countries may well fall as educational supply grows unmatched by a proportionate increase in demand (Bennell, 2002).

Admittedly, the supply of skilled employees may be endogenous to returns to education; this is the case for instance if higher relative returns in a district act as an incentive for employees to acquire more education or attract people with more education. As explained below I instrument this variable with distance to school in order to correct for the potential endogeneity bias.

The rest of the control variables in (2), i.e. those included in *X*, include a number of skilled labour demand shifters. First, it is widely recognised that greater access to (and use of) technology increases the relative demand for skilled labour, even in low-income countries (Anderson, 2005). Adapting to a new technology is a difficult task which requires the use of skilled labour; in addition technological progress has often substituted unskilled labour (e.g., automatic assembly lines); finally, greater access to foreign technology allows developing countries to compete internationally in more skill intensive goods, raising their average skill intensity of production, and thus the relative demand for skilled labour. As the data available does not allow for the construction of a direct measure of the use and availability of technology in production, I use the average distance from each community in a district to the closest telephone booth (*telephone distance*) as a technological proxy.[14] In the fixed effects specification in (2) this variable is effectively measuring a district-wise public telephone density. The telephone infrastructure was not developed in the nineties in Uganda (as in most LDCs). The median distance of a community from a telephone boot was 15 Km (and the mean was 27) in 1992 and 12 Km in 2000 (with the mean of 24). This infrastructure can be considered as a good proxy of the technological frontier in such a context. This is confirmed also by the fact that *telephone distance* is statistically associated with the share of households' income spent on electronic goods: the lower the telephone density the lower the consumption of electronic goods.[15]

[13] One of the major steps in this respect has been the introduction of Universal Primary Education (UPE) in 1996.
[14] This supply of technology variable is likely to be more exogenous than demand-related variables, such as the share of households' purchase of electronic goods in total household income is likely to be correlated with the same factors determining also returns to education, including level of education and income.
[15] The result is also robust to the inclusion of the average education and the initial share of electronic assets in total durable assets in the district (results available upon request).

Second, the urban sector is usually one in which the return to human capital is higher than in the rural sector. That is one of the reasons why other things being equal urban dwellers accumulate more human capital than their rural counterpart (Glaeser and Maré, 2001; Lucas, 2004). Once controlling for the average level of human capital in the workforce, the share of employees located in urban areas (*urban employment*) should then be positively associated with returns to education. Finally, I also include the share of males in the population (*male share*) within the vector X to control for gender-based differences in demand for goods and services.[16]

As mentioned above, *employees education* is likely to be endogenous in (2). Therefore I instrument it with the average distance of each community to the nearest primary school (*primary school distance*). In a context like Uganda with poor transport infrastructure and availability of vehicles, distance to school indeed represents a relevant obstacle to school attendance. According to the data in both LSMS surveys over 2% of schooling age population quoted distance from school as the most important reason for never enrolling into school or for dropping out. In 32% of the communities surveyed in 1992 distance from the school was an important or a very important reason in the decision of children not to enrol in one of the three most popular community's primary schools.[17] This is not surprising given that the mean of the *primary school distance* variable in 1992 was 3.7 Km with peaks as high as 48 Km.

A potential problem with this variable is that it refers to the current period while the workforce and the adult population in general was educated before. However this may not be as big of a problem as it first appears. In fact, like most infrastructure, also schools take time to be built and to become operational. Thus there is likely to be a high degree of persistence over time in any measure based on schooling infrastructure, i.e. the average distance from school in 1992 is likely to be highly correlated to the distance during the periods when the current adult population was educated. This fact is indirectly corroborated by the results of the first stage regression presented in Table 3. *Primary school distance* is a negative and significant determinant of

[16] It is worth noting that I do not include any measure of the size of the public sector among the factors influencing the demand for skills although wages in the public sector are usually less linked to the actual marginal product of labour than those in the private sector (Flabbi et al., 2007, Yang, 2005). In fact Calì (2009) documents how the effect of the size of the public sector - defined as the percentage of employees employed in state-owned enterprises and local public firms – on returns to schooling is not significant. The results suggest that the reason behind it is that the *skilled employees* variable captures a substantial share of the public sector effect on returns via its (positive) impact on the stock of educated employees, which appears to complement the traditional effect of public sector in compressing labour market inequalities.

[17] The question was not posed in the 2000 survey.

employees education (column 1). The relatively high F-statistic and partial R-squared suggest that the explanatory power of the instrument is satisfactory.

In column (2) I also add a set of further instruments based on the same rationale, i.e. distance to the nearest secondary school (*secondary school distance*), its squared term (as its effect appears to be non linear) and the interaction between each distance variables with a post-1992 dummy. These additional instruments will be added in the regression framework along with the interaction between *employees education* and a post-1992 dummy, as we will see below. It is worth noting that *secondary school distance* has a non linear U-shaped relationship with *employees education*. When the average distance to the secondary school is high, a further increase in this distance is associated with an increase in the average level of education. A possible explanation for this pattern may be that a high average distance in the absence of motorised transportation is probably associated with schools providing full boarding to the pupils. This may reduce the drop-out rate relative to a situation where the distance is high but not enough to have boarding schools, determining a higher secondary school completion rate. This non-linear pattern seems not to apply to primary schools probably because the average distance to these schools is generally much lower than that to secondary schools.[18]

Other than being good predictors of a district's level of education, the school distance variables need also to be exogenous in (2) in order to be correctly excluded from the second stage. In other words distance from schools needs to be unrelated to other determinants of returns to education not included in (2). This condition would not hold if for instance some unobserved shocks to the labour market which increased returns to education in a district (e.g. a new manufacturing plant requiring specialised labour) induced also the establishment of new schools in the same district (e.g. to supply perspective specialised labour for the plant). Although possible in principle this type of mechanism is not likely to be at work in an essentially public pre-university schooling system as the Ugandan one (in 2000 only 10% of the primary schools were privately run). As the objective of any public education system should be to provide a public good to the population, this implies that the placement of primary and (possibly) secondary schools is likely to be mainly based on population criteria rather than on the basis of the effective

[18] For instance in 1992 there was only one district with *primary school distance* above 15 Km (i.e. a long enough distance to make a daily commute on foot almost impossible, thus requiring the presence of boarding schools) against five for *secondary school distance*.

demand for education.[19] The public good nature of schools is also one of the guiding principles of the Universal Primary Education (UPE) programme launched in 1997.

Also the regressions in Table 3 control for all other skilled labour demand shifters in (2), including *urban employment*. This inclusion is important with respect to the validity of the instrument, as it helps relieve the concern that the distance to school coefficient is not picking up the fact that distance to school is lower in urban areas and that urban areas have a higher skill demand than rural areas.

The validity of the exclusion restriction assumptions is also supported by the test of overidentifying restrictions (Hansen's J) in the 2SLS specifications with more than one instrument, reported in the tables below.

5.2.2 Identification of the trade effects

Ugandan data does not allow the construction of district-wise trade tariffs' changes, which have been used in other studies as proxy for the intensity of trade liberalisation (e.g. Attanasio et al., 2004 and Topalova, 2010). Instead I exploit the districts' location and that of the main border-posts to identify districts' relative exposure to trade. As a proxy for the intensity of trade I construct a dummy variable $Bpost_k$ (named *borderpost*) which takes the value of 1 for districts that host a major border-post or which are close to one (i.e. less than 50 km from its centroid).[20] The identification of the relevant border stations is based on the work of the UBOS, which lists the major road border-posts in 2005 (UBOS, 2006). These border-posts are indicated in Figure 2.

However, not all border-posts were equally operational during the nineties. In particular, those posts bordering Tanzania (to the South and to the East) and Kenya (to the East) were the only ones likely to experience a significant increase in international trade following the liberalisation. This is for three reasons. First, Kenya and Tanzania were among the largest trading partners of Uganda during the nineties (and they still are). Second, the majority of Uganda's international trade transited through Kenya and Tanzania, mainly through the ports of Mombasa

[19] Note that the inclusion of population as a control does not change the results in Table 3. I exclude it from the regressions below on the determinants of returns to education as its coefficient is never significant and reduces somewhat the power of the instruments although does not affect the overall results (results available upon request). This little effect of the population variable seems to be due to a similar pattern of population growth across districts with the fixed effects capturing the cross-district variation in population levels.

[20] The main results of the paper do not change when constructing the dummy assigning the value of 1 to the major border-posts and the adjacent districts.

and Dar-es-Salaam.[21] Importantly, most of this trade was (and still is) formal thus being actually affected by reductions in tariffs.[22] On the other hand, Uganda's trade with Sudan (to the North), DRC (to the West) and Rwanda (to the South-West) was very limited in the nineties, due to civil unrest in these countries or in the Ugandan districts bordering these countries (in the North and North-West). For these reasons, I consider only the border-posts with Kenya and Tanzania for the construction of the variable. Moreover, among those listed by UBOS I exclude those which were not active for most of the nineties.[23] This leaves four major border stations identified by UBOS (2006) whose name is circled in Figure 2: Busia, Malaba, Mutukula and Mirama Hills. I also add Kampala district to this list of road border-posts as it hosts the major airport in Uganda (Entebbe) through which all Uganda's international trade via air transited in the nineties (and still does). Moreover the city of Kampala has also an important border in the railway station for goods transported by train to and from Kampala.

In order to (indirectly) test the appropriateness of the choice of the *borderpost* variable as a proxy for the intensity of trade, let us perform a test using the approach developed by Nicita (2009) for Mexico based on the pass-through literature. The basic idea of this approach is to test the significance of the impact of *borderpost* on changes in prices of traded goods following the liberalisation. Let us take a simplified version of the model developed by Nicita (2009). Prices of imported goods at the district level can be expressed as the product of the international price, the exchange rate, the import tariff and transport costs:

$$P_{gkt} = e_t PX^*_{gt}(1+\tau_{gt})TC_{gkt} \qquad (3)$$

for each traded good g, district k and period t, where the asterisks denote variables expressed in foreign currency. In this framework the international price is given. For each good g, this translates into the empirical specification:

$$P_{kt} = \alpha_k + \rho_1(d_{2000} \times Bpost_k) + \rho_2 d_{2000} + \varepsilon_{dt} \qquad (4)$$

where $Bpost_k$ is *borderpost* and is an inverse measure of transport costs. As tariffs have been decreasing (and transport connections are improving) over the nineties, the time dummy is effectively an indication of trade liberalisation. I test the specification (4) on three different goods

[21] This information is based on personal communication with the UBOS staff.

[22] This naturally applies to imports; however also exports are likely to have been affected as tariffs within COMESA (which Kenya and Tanzania are both part of) were reduced; also there was an increased integration of the East Africa Community (EAC), again which both Kenya and Tanzania are members of.

[23] This identification is based on trade data and information provided directly by UBOS.

(second hand shirt, wheelbarrow and hoe), which were both imported and produced domestically in the nineties, and whose import tariffs were cut between 1994 and 2000. The coefficient of the interaction between the time dummy and the border-post variable $Bpost_k$ should capture the extent to which prices are differently affected across districts according to their proximity to the main border-posts. Following Nicita (2009), ρ_1 is expected to be negative, as the pass-through effect (from international to district-level prices) also depends on transportation costs and local production could become more profitable when transport costs are high. If these are not the major borders through which the goods are imported into Uganda, the interaction effect should not be significant.

As Table 4 shows (columns 1-4) ρ_1 is negative and significant for all goods (for hoe the coefficient is significant only at the 15% level, column 3). This result is unaffected by the inclusion of a set of controls that may capture co-determinants of prices, i.e. the districts' total population and the share of population in urban areas (column 4). To check whether the negative coefficient is not only picking up the location of those districts on the border, I test whether the same result holds even when using a dummy taking the value of 1 if the district is on the border with another country (*Border*). This variable has an insignificant effect on the price of second hand shirt (column 5), confirming that *borderpost* is identifying the major trading districts of Uganda.

Note that by construction this test is applicable only to imports but not to exports. However the *borderpost* variable is likely to capture the intensity of exports as well, as the exports transit through the same major border-posts as the imports (UBOS, 2006). In any instance the effect of the trade liberalisation process is captured mainly through the changes in imports as those are the direct consequence of the liberalisation. Moreover in 1999 exports were still 10% of GDP while imports represented over 20%. Thus it is mainly the impact of changes in imports that I test through the variable $Bpost_k$.

I use *borderpost* to test for the effects of international trade on returns to education through an extended version of equation (2):

$$\beta_{kt}^{FE} = a_k + b_1 q_{kt} + \Gamma X_{kt} + b_2 Bpost_k d_{2000} + d_{2000} + \varepsilon_{kt} \qquad (5)$$

where the interaction between *Bpost* and the post-1992 time dummy d_{2000} identifies the effects of trade on returns to education. This is essentially a difference-in-difference specification, which tests whether changes in returns to education are different in the 'treated districts' (where the treatment is the increase in trade) relative to the control group. As discussed above, both the simple (based on the 2x2x2 model) and the extended version of the H-O model would predict that $b_2<0$ in the case of Uganda.

According to the H-O model, districts relatively endowed with skilled labour are supposed to experience an increase in the relative wage of skilled employees following a rise in inter-district trade. In order to test for this hypothesis, I extend equation (5) as follows:

$$\beta_{kt}^{FE} = a_k + b_1 q_{kt} + \Gamma X_{kt} + b_2 Bpost_k d_{2000} + b_3 S_{kt,1992} d_{2000} + d_{2000} + \varepsilon_{kt} \qquad (6)$$

where $S_{kt,1992}$ is the relative supply of skilled employees as measured by the share of skilled (i.e. at least primary educated) in total employees. The prediction from the H-O model in this case is that $b_3>0$. This test is similar in spirit to that of Michaels (2008), who examines the effect of increased trade between US counties (induced by road infrastructure development) on the skilled-unskilled wage gap.

6. Results

The returns to schooling (β coefficients) estimated via equation (1) display a large statistical variation across districts (see Table A2 in the Appendix). This provides further evidence that districts are likely to represent separate labour markets. The F-test strongly rejects the null of equality of returns to education coefficients in 1992 and even more strongly in 2000. Almost all of the coefficients are statistically significant at the standard levels in both years: 29 out of 38 coefficients in 1992 and 33 of 34 in 2000 are significant at the 5% level. The only coefficient which is not significant in any years is that of Kapchorwa district, which is also the only negative coefficient in both periods. While these negative coefficients may be representative of true negative returns to education in that district, their estimation is very imprecise and the standard intervals of confidence include also positive values. Such imprecise estimates may be due to the very low number of employees over which the returns are computed (37 in 1992 and only 14 in 2000).[24] Because of this I exclude Kapchorwa district from some of the specifications in the subsequent analysis as a robustness check.

Figure 3 graphically represents the evolution of returns to education reported in Table A2 on the Ugandan map. Although the increase in returns to education is general across districts,

[24] These are among the lowest numbers of employees used in districts and over four times below the average number used in each district.

some geographical patterns of increases do emerge: interior districts in the Central and Western regions experience particularly significant rise in returns, while increases in the districts on the Eastern border appear to be more limited. I will test for some of these patterns more formally in the following analysis.

Table 5 presents the results based on equations (2) and (5)-(6'). Column (1) runs the fixed effects regression without instrumenting the measure of the supply of education - *employees education* - which has a non significant effect on returns to education. The coefficients of the other variables are consistent with the expectations: negative for *telephone distance* and positive for *urban employment* and *male share*, although the latte is not significant. Instrumenting *employees education* (with primary school distance) reverses its sign, which is now negative (column 2). This result is consistent with the idea that endogeneity may bias the educational coefficients downwards as is the case if higher returns to education generate incentives to invest in education.

Column (3) tests for the effects of international trade opening by adding the $borderpost*d_{2000}$ term (i.e. specification 5). The coefficient of this term is negative and highly significant and indicates that during the 1990s returns to education in districts more exposed to international trade (i.e. districts including a major border-post or located near to one) decreased on average by 5.5 percentage points relative to the others. This is a significant difference considering that the average returns to education were between 13 and 15% in 2000 and that they grew by around 6 percentage points during the 1990s. This result is consistent with the theoretical prediction from the simple H-O framework as well as from its extensions, as discussed above: a higher exposure to international trade seems to have increased the relative demand for unskilled labour in Uganda. The inclusion of the trade variable increases the absolute magnitude and the significance of all the other coefficients. Therefore failure to consider the impact of trade causes an underestimation of the effects of the co-determinants of returns to education. In particular, a rise in the supply of relatively educated labour is now associated with statistically lower returns to education: a 10% increase in the average years of education of wage employees is associated with a reduction of 1.6 percentage points in returns.

The *borderpost* coefficient is almost halved when the other controls are excluded and its significance drops substantially (column 4), suggesting that such controls are mediating a sizable part of the effect of international trade on returns to schooling. Without taking into account the effects of these controls one may mistakenly conclude that increased international trade has had no significant impact on wage inequality in Uganda.

Interestingly, the inclusion of the *borderpost* variable into specification (5) does not substantially affect the time coefficient (d_{2000}), which remains strongly positive. This suggests that the observed increase in returns to education over the nineties is not likely to have been driven by trade opening, but rather by other factors.

The increase in domestic trade may be a good candidate in this respect, as shown in column (5), which presents the results of specification (6). I address the likely endogeneity of the new variable *skilled employees*$_{1992}$*d_{2000} by adding secondary school distance and the interaction between distance to school variables and the post-1992 dummy to the list of instruments. The coefficient of the new variable (b_3 in equation 6) is positive and significant as expected. Returns to schooling in a district with a share of skilled employment 10% larger than the average in 1992 would rise by 3.5 percentage points above the average between 1992 and 2000. This results is consistent with Michaels (2008), who finds that increased inter-county trade induced by the construction of an inter-state highway raised wage inequality in US counties relatively abundant in skilled labour. Interestingly, the inclusion of this interaction term wipes away the positive coefficient of the post-1992 dummy, providing some prima facie evidence that the increased internal trade and market liberalisation may go some way towards explaining the rising wage inequality during the 1990s.

The addition of this interaction term (*employees education*$_{1992}$*d_{2000}) makes the *skilled employees* coefficient not significant, suggesting that the positive effect of education supply on inequality is entirely driven by the positive effect in districts with relatively educated employees.

Along similar lines of specification (6), I also examine whether there is a differential effect of increased international trade on returns to education according to the initial level of skills. I add a double interaction term to (6):

$$\beta_{kt}^{FE} = a_k + b_1 q_{kt} + \Gamma X_{kt} + b_2 Bpost_k d_{2000} + b_3 S_{kt,1992} d_{2000} + b_4 (S_{kt,1992} Bpost_k d_{2000}) + d_{2000} + \varepsilon_{kt} \qquad (6')$$

with the H-O hypothesis being that $b_4>0$ (i.e. the negative effect of international trade on returns to education is smaller in districts relatively endowed with skilled employees than in the others). In column (6) I first test for this effect without accounting for the general positive effect of initial education on returns to schooling. The b_4 coefficient is positive and significant, i.e. among those districts more exposed to trade more skilled districts experience a lower decline in returns relatively to the others. However the result does not hold when I put *empl.* $edu_{1992}*d_{2000}$ back in the regression, as the b_4 coefficient becomes not significant although it remains positive (column 7). This result suggests that there is no statistically differential impact of trade across districts on the basis of skills, i.e. b_4 in (6') is not significantly different from zero. The addition of this triple interaction inflates both the $borderpost*d_{2000}$ coefficient and its standard errors probably due to the high collinearity with *Skilled* $empl_{1992}*borderpost*d_{2000}$.

An interesting further question is whether any other trade theory than the traditional H-O framework can offer an alternative (or complementary) explanation for the effects of trade reforms on the skilled-unskilled wage gap observed in Uganda. The New Economic Geography core-periphery model (Krugman, 1991) predicts that a reduction in trade costs gives firms an incentive to move production to regions with relatively good access to foreign markets, such as border areas or port cities. Hanson (1997) shows that following NAFTA production within Mexico relocated towards the Mexico-US border and away from Mexico City. In the case of Uganda the international trade liberalisation was mainly a unilateral liberalisation, as described above. Therefore increased access to markets during the nineties did not happen much along the country's international borders, but it rather occurred within the domestic borders through a reduction in internal trade costs. According to the core-periphery model this should lead to an increased concentration of economic activity near the main market(s). In the case of Uganda Kampala district represents by far the densest area in terms of economic activity in the country. But while the predictions of this framework are clear with respect to the (re-)location of economic activity within a country following a trade shock, they are less clear as far as the districts' skilled-unskilled wage gap is concerned. This will ultimately depend on the skill intensity of the productions the district specialises in. If the domestic goods that Kampala's market demands are relatively skill intensive then a reduction in internal trade barriers may lead to the concentration of relatively skilled economic activity in nearby districts. All else equal, this would generate an increase in the skilled wage premium in those districts and vice-versa in the districts further away.

Empirically I try to identify any such effects through an interaction term between the distance (in log) to Kampala and the post-1992 dummy.[25] I add this variable in column (8), but

[25] This distance is calculated with the method of the great circle distance between the city of Kampala and the centroid of the district.

the coefficient is not significant (while *borderpost* remains unaffected). Taking out borderpost from the regression – which is not shown here - makes distance to Kampala positive and significant. This suggests that after liberalisation districts closer to Kampala experience a decline in returns relative to those further away and this effect is driven entirely by Kampala's role as a major border-post rather than as a major market.

6.1. Robustness

Although the main results of the analysis appear to be quite neat, I test their robustness to a variety of checks. Tables 6 and 7 present the results of these checks. First, one potential concern with these results is that they may be driven by the choice of first stage specification to estimate the β coefficients used as dependent variables in equation (6). In column (1) I use β coefficients estimated through OLS without community fixed effects as dependent variables (β_{OLS}). This way of estimating β allows for more degrees of freedom in the first stage. Both the main coefficients of *borderpost* and $Empl\ edu_{92}*d_{2000}$ are little affected. So far the estimation of β has been based on the other Mincerian coefficients in the first stage (vector W in (1)) being constrained not to vary across districts. In column (2) I use β coefficients estimated by relaxing this constraint, which is analogous to estimate the βs via separate regressions for each district ($\beta_{FE}(sep)$).[26] Again the main results are robust to this change of variable.

The standard method of computing the βs through (1) assumes the same marginal effect across education categories. As it is often the case, however the marginal effect may differ between levels of education. In order to allow for such heterogeneous effects, I re-write (1) as an extended wage equation with the educational variable constructed as a spline function with N nodes at selected levels of education (see Moll, 1996):

$$wage_{ic} = \sum_{k=1}^{N} \alpha_k d_{ik} + \sum_{k=1}^{N}\sum_{j=1}^{J} d_{ik}\beta_{kj}S_{ij} + \Pi W_i + \gamma_c + \varepsilon_i \qquad (1')$$

[26] In particular the first stage specification becomes: $wage_i = \sum_{k=1}^{N}\alpha_k d_{ik} + \sum_{k=1}^{N}\beta_k d_{ik} S_i + \sum_{k=1}^{N} \mathrm{B}_k D_{ik} W_i + \gamma_i + \varepsilon_i$.

where S_{ij} is the number of years attended at the *j-th* level. In particular, I follow the three educational levels in Uganda to mark the nodes, i.e. primary, secondary and post-secondary.[27] The main results are robust also to using returns to primary schooling as the dependent variable (column 3). The absolute magnitude of *borderpost* and especially of *empl* $edu_{92}*d_{2000}$ coefficients increases, suggesting that the impact of both variables on returns to schooling is driven by their effect on returns to primary schooling. In other words higher exposure to international trade has the effect of decreasing the employees' wage gap essentially between primary educated and less educated employees.

Following Flabbi et al (2008), I also estimate specification (6) through WLS using the inverse of the standard errors of the returns to education coefficients as weights. This should account for the fact that the β are estimated with different levels of precision in (1). Again, estimates obtained through this method change little compared to the baseline estimation (column 4). The results are also robust to the exclusion of the outlier district Kapchorwa (column 5).

As mentioned above, to the extent that a higher premium for educated labour attract higher skilled employees, in a context of unrestricted labour mobility the results may be biased by reverse causality. Another way to tackle this potential problem other than IV estimation is to control for inter-district migration. Data on inter-district migrants between 1992 and 2000 is available in the 2000 survey. Therefore I add to the expression (6) the interaction between the share of skilled migrants (i.e. at least primary educated) in the district's total adult population and the post-1992 dummy. I also include its squared term due to the non linear effect of migration on returns:

$$\beta_{kt}^{FE} = a_k + b_1 q_{kt} + \Gamma X_{kt} + b_2 Bpost_k d_{2000} + b_3 S_{kt,1992} d_{2000} + b_4 Mig_k d_{2000} + b_5 Mig_k^2 d_{2000} + \varepsilon_{kt} \qquad (7)$$

The inequality reducing effects of trade is robust even to the inclusion of migration controls. The signs of the migration variables are consistent with the idea that a higher skilled immigration rate depresses returns to schooling as it raises the supply of skills up to a certain threshold, after which returns benefit from higher skilled migration. The latter phenomenon is in

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[27] Thus the variable S_{ij} is constructed as follows: $S_{i1}=\begin{cases} S_i & if\ 0 \le S_i \le 7 \\ 7 & if\ S_i \le 7 \end{cases}$ $S_{i2}=\begin{cases} 0 & if\ S_i \le 7 \\ S_i - 7 & if\ 7 < S_i \le 13 \\ 6 & if\ S_i > 13 \end{cases}$

$S_{i3}=\begin{cases} 0 & if\ S_i < 13 \\ S_i - 13 & if\ S_i \ge 13 \end{cases}$

line with the hypothesis of increasing returns to skills in situation of relatively high concentration of skilled workers. The inclusion of the migration controls reduces somewhat the *employees edu* interaction coefficient making it only mildly significant. This suggests that some of the greater rise in returns to education in relatively highly skilled vis-à-vis the less skilled districts following the domestic liberalisation is explained by inter-district migration.

6.2. Robustness for *borderpost*

So far the results for the effects of international trade on returns to schooling have implicitly relied on the assumption that the 'treatment' (i.e. increased international trade over the 1990s) was exogenous to returns to education as it was determined by districts' location. As all the major border-posts had been in place for some time before the beginning of the liberalisation process, this assumption seems reasonable. But what about if those districts had been selected to host the border-posts on the basis of certain unobserved characteristics (e.g. good infrastructure) which had also an effect on returns to education? The districts' fixed effects should account for these characteristics as long as they are time invariant. However to the extent that some unobserved shocks (e.g. an improvement in the infrastructure network to facilitate international trade) may affect only those districts hosting a border-post (or close to one) as well as returns to education, this may generate a bias in the $borderpost*d_{2000}$ coefficient. In order to control for this possibility I instrument $borderpost*d_{2000}$ with a dummy identifying all the districts bordering Kenya and Tanzania interacted with the post-1992 dummy. This variable is highly correlated with *borderpost* due to the way the latter has been constructed (i.e. considering only the active border-posts during the 1990s, which were on the borders with Tanzania and Kenya). This is confirmed by the high F-statistics (14.08) and partial R-squared (0.55) of the first stage regression for *borderpost* (not shown here). Furthermore, the test of overidentifying restrictions (Hansen's J) supports the exclusion restriction assumptions for all IV specifications (Table 7, column 1).

Instrumenting *borderpost* has the effect of raising slightly its coefficient (cf. Table 7, column 1) which remains highly significant, like all other variables as well. The inclusion of *borderpost* in the endogenous variable list reduces the F-statistics. To control that the results are not influenced by the relatively weak predictive power of the instruments, I exclude $Empl.\ edu_{92} * d_{2000}$ from the regression in column (2). This raises the first stage F-statistics again while the

Borderpost coefficient is reduced back to the level of the not instrumented regression (cf. Table 5, column 3). I also run the same regression as in column (1), which includes *Empl.* $edu_{92} * d_{2000}$, using the Limited Information Maximum Likelihood (LIML) method, which performs better than standard IV when instruments are weak (Stock and Yogo, 2005). The results are once again unaffected with the Borderpost coefficient raised further (column 3).

Finally, I check that *borderpost* is not capturing the effects of districts being located on the foreign border of Uganda rather than the actual effects of trade. In order to do that I use a different proxy for the intensity of trade, i.e. a variable identifying the districts on any national borders (i.e. including also those bordering Sudan, DRC and Rwanda) interacted with a post-1992 dummy ($border*d_{2000}$). The fact that this variable is not significant (column 4) adds confidence to the claim that *borderpost* identifies those districts more exposed to Uganda's international trade in the 1990s, which instead did not occur much through the country's Northern and Western borders.

7. Conclusions

Trade liberalisation is often associated with rising wage skill premium in developing countries. That was the case also in Uganda where wage inequality substantially increased during the 1990s, a decade of abrupt trade liberalisation. In order to explicitly test for the effects of this liberalisation on the wage skill premium, I employ a novel two stage district-level analysis, exploiting districts' location to predict their relative exposure to trade and controlling for skills' supply and demand related variables. In this way the analysis can also overcome the lack of firm-level data, which analyses of trade and inequality usually are based on, typical of LDCs during that period.

I find that the rising wage skill premium in Uganda – measured by returns to schooling - is not the product of international trade integration, which on the contrary tends to reduce the wage premium consistently with the predictions of the standard H-O framework. In particular, during the 1990s returns to education in districts more exposed to international trade decreased on average by 5.5 percentage points relative to the others. On the other hand, the intensification of domestic trade across districts during the 1990s (another by-product of the reforms) was associated – again in line with the H-O predictions - with larger increases in returns in those districts relatively endowed with skilled employees at the beginning of the period.

These findings are robust to using a variety of dependent and controls, estimation methods and samples. They also appear to be causal in nature as confirmed by the results of the IV estimation which tackles the potential endogeneity of the skills' supply variables through a set of instruments based on districts' distance to primary and secondary schools.

These results reconcile the widening wage skill gap following trade opening with the standard H-O model predictions of a relative rise in the returns to the (relatively) abundant factor of production, i.e. unskilled labour in the case of Uganda. They suggest that the standard H-O model - and the related Stolper-Samuelson theorem - may well be a suitable framework to analyse the effects of trade integration on wage inequalities in low income unskilled labour abundant countries.

More importantly from a policy angle, these findings support the idea that trade integration tends to reduce wage inequalities in countries endowed with unskilled labour relatively to their trading partners. LDCs as Uganda usually fall in this category. More evidence on the effects of trade integration in other such countries would be needed to substantiate this relationship more robustly. In addition the evidence available is so far confined to the case of wage employees, while little is known about the impact on self-employed, who usually operate in a separate – and very important – labour market than wage employees.

If it wasn't trade liberalisation, what drove the increase in wage inequality in Uganda during the 1990s? While the effect of domestic market reforms and the related increased in cross-district trade seems to account for most of the rise in returns to education in our analysis, the question remains open. Given the data constraints, in particular the inability to measure cross-district trade, this finding can provide only some prima facie evidence of the possible direction of the effects of domestic market reforms on wage inequality. A more precise identification of such effects would be needed in order to properly assess to what extent these reforms may explain rising returns to schooling in the 1990s in Uganda. In a similar vein, more analysis would be needed on the role of SBTC in explaining the rise in inequality. The analysis above provides some evidence that higher access to technology at the household level increases returns to education. However better proxies for access to technology, possibly at the firm level, would be needed before settling the question on the role of SBTC in wage inequality.

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TABLES AND FIGURES

Table 1: Uganda's import tariff rates between 1994 and 2000

	Average	Std. Dev.	Min	Max
WORLD				
Tariff rate 1994	17.34%	9.09	0	60
Tariff rate 2000	9.36%	5.42	0	15
COMESA				
Tariff rate 1994	9.78%	9.12	0	66
Tariff rate 2000	4.25%	2.11	0	6

Source: UNCTAD Trade Analysis and Information System

Table 2: Measures of price variation across Ugandan districts, 1992-2000

	Year	Mean	Std. Dev	Min	Max	range/mean	StDev/mean
Soap	1992	685	58	600	925	0.47	0.09
	2000	818	64	772	1,083	0.38	0.08
Hoe	1992	3,536	518	2,165	4,838	0.76	0.15
	2000	3,277	296	2,981	4,583	0.49	0.09
Wheelbarrow	1992	26,389	9,761	1,500	60,000	2.22	0.37
	2000	42,178	6,717	20,000	54,643	0.82	0.16
Bicycle	1992	94,235	13,745	69,286	131,667	0.66	0.15
	2000	87,632	11,687	76,667	139,375	0.72	0.13

Source: author's elaboration on household survey data

Table 3: First stage regressions for *employees education*

	(1) *Employees education*	(2) *Employees education*
Primary school distance	-0.070***	-0.062***
	(0.016)	(0.014)
Sec. school distance		-0.188**
		(0.075)
Sec. school distance sq.		0.0048***
		(0.0013)
*Primary school dist.*Post-92*		-0.841**
		(0.387)
*Sec. school dist.*Post-92*		0.198***
		(0.048)
Other controls	YES	YES
Observations	68	68
Nr. of districts	34	34
R-sq. (within)	0.249	0.549
Partial R-sq.	0.210	0.499
F-stat	18.46	96.18

*Robust standard errors (Huber-White method) in parentheses; *significant at 10%; ** significant at 5%; *** significant at 1%; all regressions include district fixed effects and year effect; other controls include the other controls in Table 5. Partial R-squared is the R-squared of the school distance variables (i.e. excluded instruments in the Tables 5-7); F-statistics refers to the test for the joint insignificance of the school distance variables.*

Table 4: The effects of distance from border-posts on changes in prices of a unit of traded goods, 1992-2000

	(1) Shirt	(2) Wheelbarrow	(3) Hoe	(4) Shirt	(5) Shirt
Borderpost x d2000	-0.129**	-0.282*	-0.076	-0.129**	
	(0.050)	(0.170)	(0.052)	(0.049)	
d_{2000}	-0.027	0.638***	-0.056	-0.056	-0.058
	(0.044)	(0.162)	(0.050)	(0.047)	(0.073)
Border x d_{2000}					-0.011
					(0.081)
Constant	7.818***	10.079***	8.164***	7.879***	7.819***
Controls	NO	NO	NO	YES	NO
Observations	71	69	72	71	71
Nr. of district	38	38	38	38	38
R-squared	0.201	0.451	0.150	0.215	0.113

*Robust standard errors (Huber-White method) in parentheses; dependent variable is the price (in Ugandan Schelling) of the product in log; * significant at 10%; ** significant at 5%; *** significant at 1%; all regressions include district fixed effects. Controls include the share of population in urban areas and districts' total population.*

Table 5: The effects of trade on returns to education in Uganda, 1992-2000

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	FE	FE IV	FE IV	FE	FE IV	FE IV	FE IV	FE IV
	β_{FE}	β_{FE}	β_{FE}	β_{FE}	β_{FE}	β_{FE}	β_{FE}	β_{FE}
Borderpost x d2000			-0.055**	-0.032	-0.062***	-0.488**	-0.291	-0.059**
			(0.026)	(0.022)	(0.022)	(0.224)	(0.206)	(0.024)
Employees education	0.004	-0.011	-0.016*		0.024	-0.014*	0.017	0.023
	(0.013)	(0.008)	(0.009)		(0.015)	(0.008)	(0.013)	(0.014)
Telephone distance (x100)	-0.069**	-0.078***	-0.090***		-0.078***	-0.092***	-0.080***	-0.075***
	(0.030)	(0.028)	(0.019)		(0.019)	(0.016)	(0.016)	(0.018)
Urban employment	0.159*	0.171**	0.225***		0.284***	0.243***	0.270***	0.281***
	(0.084)	(0.086)	(0.074)		(0.079)	(0.063)	(0.075)	(0.082)
Male share	0.402	0.378	0.511		0.927*	0.539	0.815*	0.931*
	(0.478)	(0.473)	(0.475)		(0.480)	(0.419)	(0.428)	(0.481)
Empl. edu$_{1992}$ x d_{2000}					0.035**		0.024	0.034**
					(0.014)		(0.016)	(0.016)
Empl. edu$_{1992}$ x *Borderpost* x d_{2000}						0.055**	0.030	
						(0.028)	(0.026)	
Log Distance to Kampala x d_{2000}								0.005
								(0.008)
d_{2000}	0.062***	0.068***	0.089***	0.060***	-0.174*	0.091***	-0.092	-0.195
	(0.017)	(0.017)	(0.014)	(0.013)	(0.102)	(0.013)	(0.118)	(0.144)
Observations	72	68	68	68	68	68	68	68
Nr. of districts	38	34	34	34	34	34	34	34
R-sq. (within)	0.516	0.459	0.509	0.421	0.566	0.624	0.635	0.571
1st stage F-stat		17.40	17.42		3.476	17.63	3.227	3.250
Hansen J-stat					3.627		4.130	4.549
Chi-sq(3) P-val.					0.305		0.248	0.208

*Dependent variable is return to education as estimated through equation (3). Robust standard errors (Huber-White method) in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. F-stat is the statistics for the joint significance of the excluded instruments in the first stage; Hansen J-statistic is the over-identification test. Endogenous variables are skilled employees, population edu and skilled employees*$_{1992}$*d_{2000}. *Skilled employees and population edu are instrumented through distance to primary school, distance to secondary school, distance to secondary school squared and distance to secondary school x Post-1992 dummy. In the specifications with Skilled empl*$_{1992}$*d_{2000}, *distance to secondary school squared is replaced by distance to primary school x Post-1992 dummy in the instrument set. The 4 districts with a single observation are dropped in the IV estimations.*

Table 6: The effects of trade on returns to education in Uganda, 1992-2000, robustness checks

	(1)	(2)	(3)	(4)	(5)	(6)
	FE IV	FE IV	FE IV	FE IV (WLS)	FE IV	FE IV
	β_{OLS}	β_{FE} *(sep)*	$\beta_{OLS}^{PRIMARY}$	β_{FE}	β_{FE}	β_{FE}
Borderpost x d_{2000}	-0.048**	-0.063***	-0.086**	-0.052**	-0.044**	-0.049**
	(0.019)	(0.023)	(0.043)	(0.021)	(0.018)	(0.022)
Employees edu_{92} x d_{2000}	0.048***	0.031**	0.136**	0.029**	0.033**	0.027*
	(0.014)	(0.014)	(0.062)	(0.014)	(0.015)	(0.017)
Mig_{92} x d_{2000}						-1.182**
						(0.596)
Mig_{92} x d_{2000} *squared*						4.695*
						(2.623)
Other controls	YES	YES	YES	YES	YES	YES
Observations	68	68	68	68	66	68
Nr. of districts	34	34	34	34	33	34
R-squared	0.632	0.554	0.406	0.626	0.674	0.627
1^{st} stage F-stat	3.476	3.476	3.476	4.822	3.557	2.462

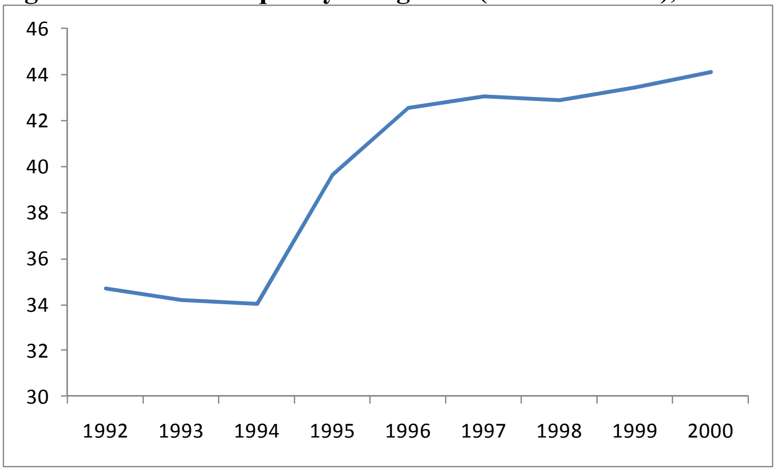
*Dependent variables are return to education as estimated through equation (1) and its variants. Robust standard errors (Huber-White method) in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. F-stat is the statistics for the joint significance of the excluded instruments in the first stage; Hansen J-statistic is the over-identification test. Endogenous variables are employees education and employees edu_{1992} x d_{2000}. Employees education and employees $education_{1992}$ x d_{2000} are instrumented through distance to primary school, distance to secondary school, distance to primary school x Post-1992 dummy and distance to secondary school x Post-1992 dummy. In column 5 Kapchorwa district is excluded.*

Table 7: The determinants of returns to education, robustness for *Borderpost*

	(1) FE IV	(2) FE IV	(3) FE IV LIML	(4) FE IV
Borderpost x d_{2000}	-0.072**	-0.054*	-0.088*	
	(0.033)	(0.029)	(0.049)	
Employees edu_{92} x d2000	0.036**		0.060*	0.038
	(0.015)		(0.032)	(0.024)
Border x d_{2000}				-0.019
				(0.029)
Borderpost instrumented	YES	YES	YES	NO
Other controls	YES	YES	YES	YES
Observations	68	68	68	68
Number of districts	34	34	34	34
R-squared	0.553	0.590	0.329	0.444
1st stage F-stat	3.099	15.10	3.099	5.555
Hansen J-stat	4.725	4.593	3.323	1.848
Chi-sq(3) P-val.	0.193	0.332	0.344	0.605

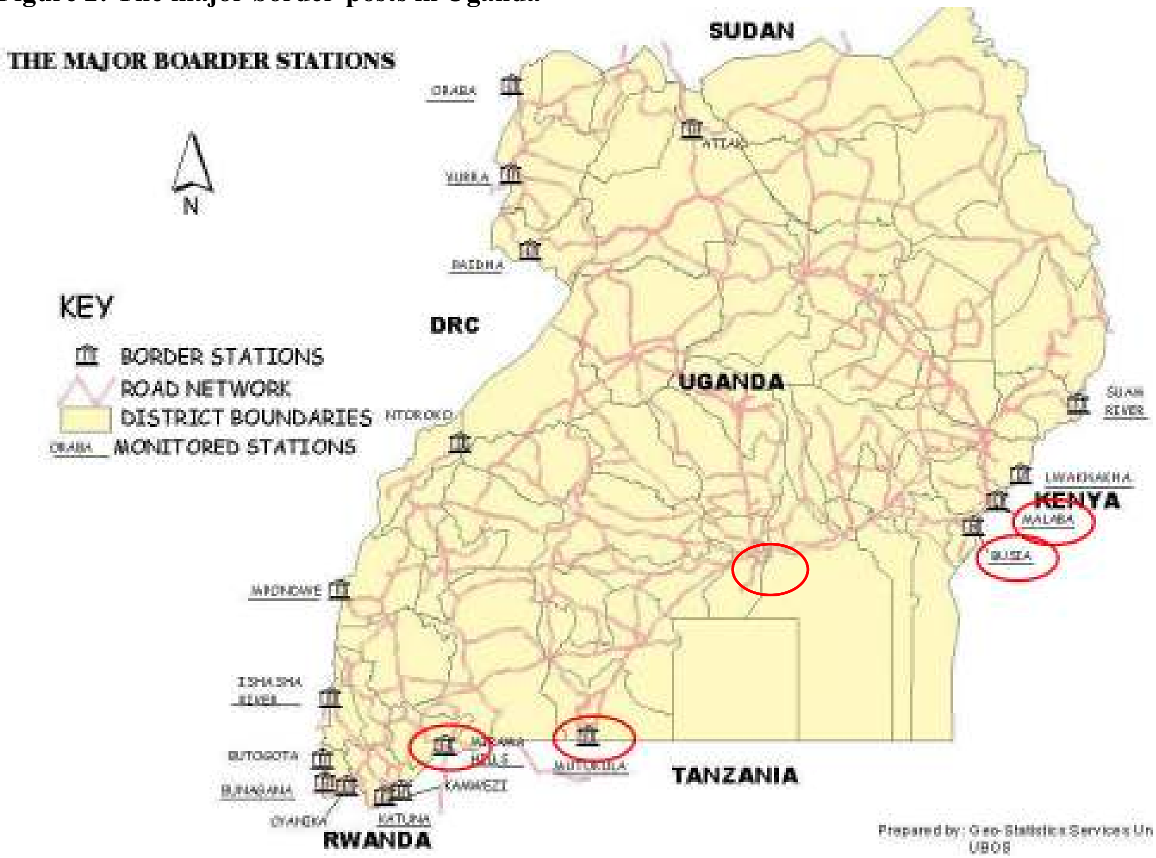
*Dependent variables are return to education as estimated through equation (1). Robust standard errors (Huber-White method) in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%. F-stat is the statistics for the joint significance of the excluded instruments in the first stage; Hansen J-statistic is the over-identification test. Endogenous variables are employees education, employees edu_{1992} x d_{2000} and Borderpost in those columns in which it is indicated. Employees education, employees edu_{1992} x d_{2000} are instrumented through distance to primary school, distance to secondary school, distance to primary school x Post-1992 dummy and distance to secondary school x Post-1992 dummy. Borderpost*d_{2000} is instrumented through a dummy for districts bordering Kenya and Tanzania.*

Figure 1: Income inequality in Uganda (EHII measure), 1992-2000



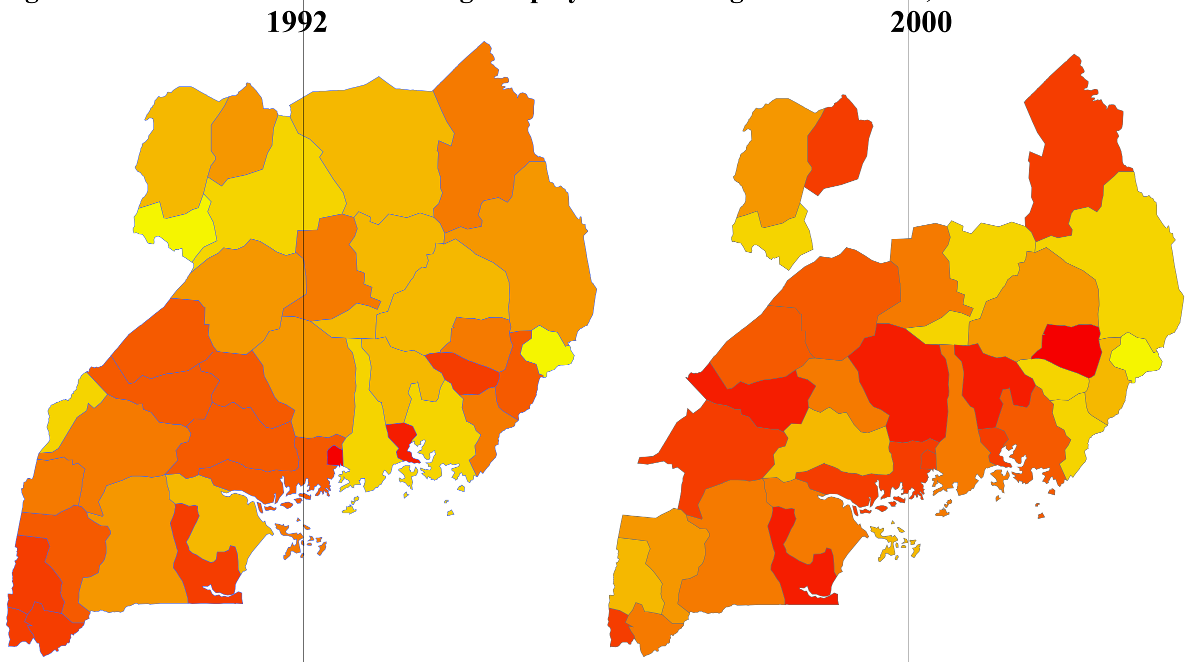
Source: University of Texas Inequality Project (UTIP) database

Figure 2: The major border-posts in Uganda



Source: UBOS (2006)

Figure 3: Returns to education for wage employees across Ugandan districts, 1992-2000



Note: returns to education are increasing in the tonalities of red (i.e. the higher the returns, the more red the district)

Appendix

Table A1: Description of district-level variables

Variable's name	Units	Description
Employees education	Years	District-wise average number of years of formal schooling completed by the wage earners
Telephone distance	Km	Average distance from the centre of each community within a district to the closest public telephone booth
Urban employment	Percentage	District-wise number of wage earners located in urban areas over total district's number of wage earners
Male share	Percentage	District-wise share of males in the adult population
Primary school distance	Km	Average distance from the centre of each community within a district to the closest primary school
Secondary school distance	Km	Average distance from the centre of each community within a district to the closest secondary school
Borderpost	Dummy	Value 1 if the district hosts a major border-post or is close to one (i.e. less than 50 km from its centroid); 0 otherwise.
Border	Dummy	Value 1 if the district is bordering a foreign country; 0 otherwise.
Migration$_{92}$	Percentage	District-wise population with primary education or above that migrated into the district after 1992 over total adult population of the district
Skilled employees	Percentage	District-wise share of wage earners who have completed an education equal or higher than the primary level (i.e. 7 years or more)

Source: World Bank LSMS on Uganda for 1992 and 1999/2000.

Table A2: Returns to education across Ugandan districts, 1992-2000

District No.	District name	1992			2000		
		β_{1992}	S.E.	Nr obs	β_{2000}	S.E.	Nr obs
1	Kalangala	0.082	0.034	56	0.133	0.023	32
2	Kampala	0.105	0.010	430	0.173	0.013	235
3	Kiboga	0.082	0.026	44	0.158	0.068	23
4	Luwero	0.081	0.019	64	0.193	0.031	74
5	Masaka	0.079	0.015	169	0.159	0.014	128
6	Mpigi	0.146	0.014	237	0.174	0.016	199
7	Mubende	0.141	0.019	73	0.136	0.017	69
8	Mukono	0.070	0.019	161	0.157	0.023	161
9	Rakai	0.208	0.034	58	0.203	0.032	37
11	Iganga	0.072	0.023	93	0.165	0.028	87
12	Jinja	0.130	0.015	242	0.178	0.017	125
13	Kamuli	0.032	0.032	44	0.202	0.023	64
14	Kapchorwa	-0.025	0.019	37	-0.058	0.113	14
15	Kumi	0.048	0.020	49	0.245	0.024	24
16	Mbale	0.072	0.018	126	0.131	0.019	129
17	Pallisa	0.096	0.032	34	0.117	0.037	32
18	Soroti	0.073	0.031	62	0.148	0.032	61
19	Totoro	0.048	0.021	90	0.114	0.019	104
21	Bundibugyo	0.010	0.035	32			0
22	Bushenyi	0.133	0.022	107	0.144	0.017	97
23	Hoima	0.147	0.035	39	0.164	0.023	61
24	Kabale	0.084	0.024	63	0.160	0.023	71
25	Kabarole	0.114	0.013	181	0.182	0.014	115
26	Kasese	0.049	0.022	110			0
27	Kibaale	0.128	0.022	38	0.191	0.036	38
28	Kisoro	0.027	0.041	22	0.184	0.038	41
29	Masindi	0.090	0.029	39	0.162	0.026	45
31	Mbarara	0.098	0.013	240	0.155	0.017	156
32	Rukungiri	0.093	0.032	55	0.140	0.017	58
41	Apac	0.116	0.022	44	0.161	0.031	26
42	Arua	0.061	0.020	92	0.149	0.027	87
43	Gulu	0.072	0.019	93			0
44	Kitgum	0.117	0.029	64			0
45	Kotido	0.175	0.056	65	0.172	0.048	42
46	Lira	0.059	0.023	78	0.113	0.029	72
47	Moroto	0.149	0.030	80	0.115	0.021	29
48	Moyo	0.047	0.025	53	0.180	0.034	45
49	Nebbi	-0.019	0.034	21	0.116	0.034	27
	Mean	0.087	0.025	94	0.153	0.029	77
	S.D.	0.049	0.009	82	0.048	0.018	53
	Signif. (5%)	29/38			33/34		
	Signif. (1%)	24/38			30/34		
	F-stat	15.99			37.70		

Note: βs are estimated through the FE method in equation (1); the F-stat is the test for the joint significance of the β coefficients; Nr obs. refers to the number of observations over which β is estimated.

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