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# Returns to education and wage inequality in Namibia

A gendered analysis

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**Abstract:** This paper estimates the returns to education and their implications for wage inequality using data from the 2015/16 Namibia Income and Expenditure Survey. The paper employs recentred influence function regression to analyse the impact of education across the wage distribution and uses a simulation approach to assess the impact on wage inequality of educational equalization. The results indicate higher returns to education among women compared with men. The quantile results further reveal that within the sample of women, higher returns to education are observed at the lower percentile compared with the upper end of the distribution, while within the sample of men, higher returns to education are observed at the upper end of the wage distribution. Simulation results suggest that educational expansion substantially widens inequality relative to equalization efforts for both men and women. The policy implications include the need to prioritize reducing educational disparities and ensuring equal access to quality education. Targeted interventions should address the specific needs of disadvantaged groups, such as women and individuals from lower-income households, by improving educational quality and supporting vocational training.

**Key words:** returns to education, wage inequality, educational equalization, educational expansion, Namibia

**JEL classification:** J24, C21, I26, D31

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# 1 Introduction

The high level of wage inequality and its implications for economic efficiency and standards of living worldwide have become a growing concern among politicians, economists, and international organizations. Wage inequalities are explained by the human capital theory, which is associated with Becker (1965), Mincer (1958, 1996), and Schultz (1960). According to this theory, wage disparities arise due to variations in individuals' human capital stocks, which ultimately determine their productivity levels. Becker (1964) argued that the acquisition of human capital plays a crucial role in shaping an individual's productive attributes, which are positively correlated with productivity. Differences in the level of human capital attained by workers result in differences in their marginal productivity. Consequently, wage inequalities emerge when workers are remunerated based on their marginal productivity. Hence, the theory of marginal productivity can serve as a potential framework for understanding wage inequalities, as individuals at the lower end of the wage spectrum are believed to have lower productivity levels due to their comparatively lower level of attainment of human capital compared with those at the higher end. As such, insufficient educational endowments can be seen as the underlying cause of poverty and wage inequality in Sub-Saharan African countries, including Namibia (Baye et al. 2023).

Namibia's government has prioritized expanding education to address wage inequality, but challenges persist. Despite efforts to enhance access and infrastructure, wage disparities remain high. Merely expanding educational opportunities is insufficient to tackle the root causes of wage inequality. The Maximally Maintained Inequality (MMI) theory explains this issue by highlighting the unequal distribution of educational opportunities. Higher-income individuals benefit more from educational expansion, thanks to resources like private tutoring and enrolment in top-tier schools. Their extensive social networks provide valuable information and support in navigating the education system and securing better job prospects. This concentration of educational advantages among the wealthier exacerbates wage inequality in Namibia. To make meaningful progress, addressing the unequal distribution of resources and opportunities is crucial in reducing wage disparities.

Addressing the unequal distribution of educational opportunities and its impact on wage inequality in Namibia may require a focus on educational equalization policies. These policies aim to ensure equal access to educational resources and opportunities for all individuals, regardless of their socioeconomic background. Implementing measures such as targeted scholarships, financial assistance programmes, and additional academic support for students from disadvantaged backgrounds is crucial. By providing these resources, educational institutions can help to bridge the gap between disadvantaged individuals and their more affluent counterparts, enabling them to fully benefit from educational expansion initiatives. Such efforts are vital in reducing wage disparities and promoting a more equitable society in Namibia.

In many countries, including Namibia, women have been catching up with men in terms of education enrolment and completion rates. Indeed, according to Lembani (2019), statistics from the three main universities in Namibia—the University of Namibia (UNAM), Namibia University of Science and Technology (NUST), and the International University of Management (IUM)—indicate that the total number of female students consistently exceeds the number of male students. The educational preferences of men and women in Namibia can be attributed to a combination of socio-cultural factors, including the evolving roles of women in various societies, as well as economic motivations. Given the individual heterogeneity in optimal schooling choices resulting from variations in the costs of education and the marginal return on investment in education (as discussed by Card 1999), gender disparities in the returns to education could potentially explain the divergence in educational demand between genders. If the returns to education were higher for women than for men, the increase in women's educational attainment would not only hold significant sociological implications for a society like Namibia but would also be economically justifiable. Consequently, it is highly relevant to examine whether there are gender differences in the returns to education and if so, to explore their implications for wage inequality.

While the labour market and education landscape in Namibia indicate the potential existence of gender disparities in the returns to education and their impact on wage inequality, the current empirical evidence regarding whether these returns are higher or lower for women compared with men and the specific implications for wage inequality in Namibia have not been examined. Some empirical studies have endeavoured to investigate this issue in other countries. Arrazola and José de Hevia (2006) conducted an empirical study in Spain that examined gender differentials in returns to education. They found that the returns to education are greater for women compared with men. Similarly, Lee and Ihm (2020) used data from Korea and observed higher returns to education for women than for men. However, in contrast, Huang et al. (2022) analysed Chinese data and found that the returns to education are higher for men than for women. These contrasting findings suggest that gender differentials in returns to education remain an empirical question that requires further investigation in the context of Namibia.

In the context of examining the implications of returns to education for wage inequality, Tansel and Bodur (2012) conducted a study in Turkey, revealing that education contributes to an increase in wage inequality. Similarly, in Greece, Chletsos and Roupakias (2020) explored the relationship between education and the dispersion of men's earnings, finding that education plays a significant role as a driver of wage inequality. In the case of Cameroon, Baye et al. (2023) investigated the impact of education on wage inequality, uncovering that educational expansion has a widening effect on wage inequality while educational equalization has a reducing effect on wage inequality. However, despite the policy relevance of empirical evidence regarding returns to education along gender lines and its implications for wage inequality, no empirical study has undertaken research in this area in Namibia. To the best of our knowledge, published works on returns to education and the implications of education for wage inequality have not yet emerged in Namibia. This paper aims to address this research gap by estimating the returns to education across the wage distribution, both overall and by gender. Additionally, it seeks to assess the impact of educational expansion versus educational equalization on wage inequality in Namibia.

To achieve our set objectives, we employ a robust methodology to estimate the Mincerian wage-generating function. Specifically, we use the recentred influence function (RIF) regression, proposed by Firpo et al. (2009), which allows us to capture returns to education across the entire wage distribution and examine the effect of education on distributional statistics such as the Gini coefficient of inequality. To address potential endogeneity concerns, we incorporate the Lewbel (2012) identification strategy based on heteroscedasticity in the data. Furthermore, we adopt the simulation approach proposed by Bourguignon et al. (2007) and employed by Baye et al. (2023) to assess the impact of educational equalization on wage inequality. By employing these rigorous methodological approaches, we aim to provide comprehensive insights into the relationship between education and wage inequality in Namibia. This allows us to analyse the returns to education across the wage distribution and evaluate the impact of educational equalization on various measures of wage inequality.

The rest of the paper is structured as follows: Section 2 provides background information on the education system and its development in Namibia. In Section 3, we detail the empirical strategy adopted for our study. Section 4 focuses on the data used in our analysis. The empirical findings are presented and discussed in Section 5. Section 6 concludes, summarizing the key findings and discussing their policy implications.

## 2 Background on the education system in Namibia

Namibia, located in Southern Africa, achieved its independence from apartheid South Africa in 1990. As of the 2021 mid-year population estimates the population of Namibia is approximately 2.59 million people (UNDESA 2021), and the nation state covers an expansive area of approximately 823,000 $km^2$. Following independence, Namibia underwent significant scrutiny in its higher education sector, notably through the establishment of the Presidential Commission on Higher Education in 1991. This led to the creation of two noteworthy state-funded public higher education institutions: UNAM in 1992 and the Polytechnic of Namibia (now University of Science and Technology—NUST) in 1994 (Turner 1991). Before this period, Namibia's higher education system was segregated along apartheid lines, with education being allocated based on racial categories for blacks and whites. As a result, significant transformations were necessary. According to Hangula et al. (2018), the primary objective of these reforms was to ensure that every Namibian had equal opportunities to access higher education. The overarching goals guiding the reforms throughout the entire education system were four-fold: access, equity, quality, and democracy (Ministry of Education and Culture 1993). These goals were further strengthened by the World Declaration on Education For All (UNESCO 1990), which emphasized education as a fundamental right for all individuals. While the strategic objectives of access, equity, quality, and democracy continued to be significant, the focus of Namibian higher education institutions (HEIs) has only recently shifted towards positioning higher education as a key pillar in national development.

The legacy of colonialism and apartheid has left deep-seated inequities in Namibia. Black Namibians and other historically oppressed groups experienced institutional discrimination and few educational options. Particularly in rural and remote places, access to education is still unequal. Disparities in access are caused by things like travel time to schools, inadequate infrastructure, and socioeconomic limitations. Children from low-income families and marginalized areas disproportionately suffer the consequences. Namibian education still has gender inequities; although there has been progress in closing the gender gap, girls continue to encounter obstacles such as early marriage, teenage pregnancies, and cultural biases that prevent them from enrolling in and finishing their education. Girls are frequently under-represented in certain academic subjects, which restricts their career options (Hailombe 2011).

Namibia does not consistently provide access to high-quality education. Resource limitations, such as poor infrastructure, a lack of instructional resources, and a paucity of trained teachers, frequently affect schools in underprivileged communities. These elements may influence educational quality, which contributes to disparities in learning outcomes. There are differences in the learning outcomes of various socioeconomic groups. Academic performance can be impacted by factors including poverty, a lack of parental support for education, and language obstacles, all of which contribute to educational attainment disparities.

Although Namibia has made progress in increasing access to higher education, there are still gaps in coverage. Students from underprivileged backgrounds find it difficult to pursue higher education due to affordability issues, university capacity issues, and regional differences. Wider socioeconomic inequities in Namibia are intimately related to educational inequalities. Access to high-quality education is hampered by poverty, unemployment, and income inequality because families find it difficult to pay for school supplies, transportation, and other related costs. Education inequalities persist because educational resources, such as funding, teachers, and facilities, are not equitably distributed. Compared with schools in underprivileged rural areas, schools in urban centres frequently have better resources and facilities.

Over the past decade, the higher education sector in Namibia has witnessed a substantial increase in student enrolment. The growth in the sector has been notable, with the registration of over 40 private institutions by the National Qualification Authority (NQA) in 2016 (Hangula et al. 2018). Several of these institutions were established not solely in response to market demands, but rather due to the significant demand from young individuals who may not meet the minimum requirements for admission into public HEIs but still seek to attain some form of post-secondary qualification. The enrolment figures from 2017 reveal a significant number of students enrolled in prominent public universities. At UNAM, total student enrolment reached 25,936, with 16,920 (65.2 per cent) being female and 9,016 (34.8 per cent) being male (UNAM 2017). Similarly, at NUST, total enrolment stood at 11,226, with 5,643 (50.3 per cent) female students and 5,583 (49.7 per cent) male students (NUST 2017).

## 3 Empirical strategy

### 3.1 Modelling the Mincerian wage equation

The Mincerian earnings function is a widely used model that links logarithmic earnings to factors such as education, experience, and a quadratic term for experience. However, researchers often incorporate additional variables into the equation to account for various factors that affect earnings, including gender, race, place of residence, and marital status (Salisbury 2016). The Mincerian equation is therefore given as:

$$LnW = X'\delta + \beta S \, \nu \tag{1}$$

In the context of the Mincerian earnings function, the variable $LnW$ represents the natural logarithm of an individual's monthly earnings within the 30-day period before the survey. The variable $S$ represents the number of years of education accumulated by the individual. The vector $X$ refers to a collection of other variables or arguments that are included in the earnings function. The coefficient $\beta$ in the earnings function represents the average rate of returns to education. It quantifies the percentage increase in earnings that can be attributed to each additional year of education. $\delta$ represents a vector of parameters for other variables considered in the earnings function, apart from education, while $\nu$ represents the error term.

However, the estimates of returns to education obtained from Equation 1 may suffer from bias due to the omission of individual ability. The ordinary least squares (OLS) equation presented above does not account for individual ability, which is correlated with education. Workers with higher ability tend to acquire more years of schooling compared with those with lower ability. This positive correlation between unobserved individual ability and years of education creates endogeneity in Equation 1 (Kaymak 2009).

Estimating returns to education using the traditional instrumental variable (IV) approach faces two common methodological challenges. The first is the absence of conventional IVs for education. The second arises when conventional IVs are available, but there are concerns regarding their compliance with the exclusion restriction assumption. In order to overcome these issues, Lewbel (2012) suggests a novel identification strategy that leverages the presence of heteroscedasticity in the model to construct internal IVs. This innovative approach provides an alternative solution for estimating returns to education when conventional IVs are lacking. Therefore, in this paper, we utilize this 'Het-IV' approach to estimate the returns to education.

$$LnW = X'\gamma_1 + \beta S + \varepsilon_1 \, , \varepsilon_1 = \alpha_1 U + V_1 \tag{2}$$

$$S = X'\gamma_2 + \varepsilon_2 , \varepsilon_2 = \alpha_2 U + V_2 \quad (3)$$

In Equations 2 and 3, $LnW$ and $S$ represent log wages and years of schooling, respectively. $U$ represents the individual's unobserved ability, which influences both schooling and productivity. $V_1$ and $V_2$ denote the idiosyncratic errors. Lewbel (2012) suggests that by exploiting heteroscedasticity in the error terms, it is possible to address endogeneity concerns using internally generated instruments $[Z - E(Z)]\hat{\varepsilon}_2$, where $Z$ is a vector of observed exogenous variables $X$ or a subset of $X$. This identification strategy relies on satisfying the following moment conditions:

$$E(X\,\varepsilon_1) = 0, E(X\,\varepsilon_2) = 0, Cov(Z, \varepsilon_1, \varepsilon_2) = 0 \quad (4)$$

The rationale for using $[Z - E(Z)]\hat{\varepsilon}_2$ as a vector of instruments is grounded in the concept that identification can be achieved by having regressors that are not correlated with the heteroscedastic disturbance term. By incorporating these instruments, it is possible to address the endogeneity issue and obtain consistent estimates in the presence of heteroscedasticity.

### 4.2 Testing for heterogeneity in returns to education across the wage distribution: recentred influence function (RIF) regression

To examine the returns to education across the unconditional wage distribution, we employ the RIF regression method proposed by Firpo et al. (2009). This approach involves regressing the RIF of the dependent variable (logwage) on a set of independent variables. The RIF is derived by substituting the original dependent variable with a linear approximation using the influence function (IF). It is important to note that the RIF satisfies the property that the integral of the IF with respect to the wage distribution is equal to zero. By definition, $\int_{-\infty}^{\infty} IF(w, \tau) dF(w) = 0$.

Specifically, for quantiles, the IF is given as below:

$$IF(w;\, Q_\tau) = \frac{\tau - \mathbb{I}\{w \le q_\tau\}}{f_{W(q_\tau)}} \quad (5)$$

Where $IF$ stands for the influence function, $W$ denotes the log wage and $f_{W(q_\tau)}$ denotes the probability density function of $W$ evaluated at $q_\tau$. $\mathbb{I}\{w \le q_\tau\}$ represents an indicator function which takes the value 1 when the outcome variable is less than $q_\tau$ and 0 otherwise. $\tau$ denotes the quantile in question. Based on this information, the RIF is computed as:

$$RIF(w;\, q_\tau) = q_\tau + \frac{\tau - \mathbb{I}\{w \le q_\tau\}}{f_{W(q_\tau)}} \quad (6)$$

An essential characteristic of the unconditional quantile (RIF) over the conditional quantile is that RIF regression is linear in expectation—that is, $\mathbb{E}[RIF(w;\, q_\tau)] = q_\tau$. Thus, the expectation of the RIF is basically the specific quantile in question. In keeping with the consideration that $X$ is a vector of independent variables and $S$ represents the years of schooling, and that the expected value of the RIF is linear in $X$ and $S$:

$$\mathbb{E}[RIF(w;\, q_\tau)|\, X, S] = X'\gamma_1 + \beta S \quad (7)$$

An important advantage of quantile regression over mean regression is that it captures heterogeneity in returns to education across the wage distribution. This allows for testing the hypothesis whether the returns to education for the poorest persons (those at the bottom of the distribution) are different from that of richer households (those at the middle or top of the distribution).

Another intriguing aspect of the RIF regression is its versatility in estimating the effect of independent variables on various distributional statistics, such as the Gini coefficient. In this paper, we further employ the RIF regression to estimate the effect of education on inequality measures, specifically the Gini coefficient and inter-quantile wage gaps. This involves a straightforward substitution of the quantile $q_\tau$ in Equation 7 with the desired distributional statistic, such as the Gini coefficient.

### 4.3 Evaluating the impact on inequality of educational equalization versus expansion

Our objective was to assess the impact on wage inequality of educational equalization, specifically the effect of reducing disparities in access to quality education among different social groups or regions. To achieve this, we followed the simulation technique proposed by Bourguignon et al. (2007) and used by Baye et al. (2023). The factual log wage distribution can be obtained by expressing it as $LnW = Ln\widehat{W} + \hat{\varepsilon}_1$ and then taking the antilog, resulting in $W = exp(Ln\widehat{W} + \hat{\varepsilon}_1)$. This equation represents the factual wage distribution and is fully presented in Equation 8.

$$W = exp(X'\hat{\gamma}_1 + \hat{\beta}S + \hat{\varepsilon}_1) \quad (8)$$

By allocating workers the mean years of schooling while keeping other variables observed, we can obtain the corresponding counterfactual education-equalizing benchmark. This gives rise to the counterfactual distribution of wages denoted by $W_{\bar{S}}$ and defined as:

$$W_{\bar{S}} = exp(X'\hat{\gamma}_1 + \hat{\beta}\bar{S} + \hat{\varepsilon}_1) \quad (9)$$

In this simulation, the measured wage inequality can be attributed to unobservable factors and other observed variables, except for years of schooling. This is because the variations in education have been eliminated by allocating the mean value to all workers, ensuring equal educational attainment across the population.

If we denote the distribution with policy as $W_{\bar{S}}$, which represents the counterfactual distribution, and the distribution without policy as $W$, along with an inequality index denoted as $I$, we can define the impact of the policy on wage inequality as $\Theta_I$:

$$\Theta_I = \frac{I(W) - I(W_{\bar{S}})}{I(W)} \quad (10)$$

If $\Theta_I > 0$, it indicates that education is increasing inequality in the factual distribution.

If $\Theta_I = 0$, it suggests that education has a neutral effect on inequality in the factual distribution.

If $\Theta_I < 0$, it signifies that education is reducing inequality in the factual distribution.

The notation $\Theta_I$ represents the contribution of education to wage inequality and is dependent on the specific inequality index used. In this study, we employ the Gini index, the Atkinson measure of inequality, and the inter-quantile wage gaps to assess the role of education in wage inequality.

## 4 Data description

The paper uses data from the Namibia Household Income and Expenditure Survey (NHIES) conducted in 2015/16. The NHIES is a nationally representative survey conducted by the Namibia Statistics Agency (NSA) every five years. It provides data on the income and expenditure patterns of households and individuals, as well as demographic characteristics and data on educational attainment. The NHIES aims to support planning, policy formulation, decision-making, and research and development for a knowledge-based economy, with the ultimate goal of eradicating poverty and reducing income inequalities in Namibia. The survey collects information on various variables including year of birth, age, gender, marital status, region of residence (urban or rural), income, and education level (NSA 2015/16).

The survey employs a stratified two-stage cluster sampling method. Geographical areas, designated as Primary Sampling Units (PSUs), were selected as the first-stage units, while households were selected as the second-stage units. The original sample consisted of 41,581 individuals, of whom 22,677 were part of the labour force. Among them, 14,026 individuals were employed, and 8,060 had complete wage data. Individuals with missing data on race or language, as well as non-Namibians, were excluded from the analysis. Thus, the final sample included 7,274 individuals. The main variables used for the analysis in this paper include earnings, education, age, the square of age, sex, race, marital status, and geographical location represented by the region of residency.

Table 1 presents the descriptive statistics of various variables, including the overall statistics as well as those disaggregated by gender. The variables include log wage, quantiles (q10, q25, q50, q75, q90), inequality measures, education years, age, urban residency, black African ethnicity, and marital status.

Table 1: Descriptive statistics of variables

| *Variable* | | Overall | | Men | | Women |
|---|---|---|---|---|---|---|
| | Mean | Std dev. | Mean | Std dev. | Mean | Std dev. |
| *Log wage* | 9.692 | 1.772 | 9.758 | 1.714 | 9.609 | 1.839 |
| q10 | 7.194 | 3.33 | 7.362 | 3.077 | 6.985 | 3.609 |
| q25 | 8.602 | 3.003 | 8.722 | 2.93 | 8.453 | 3.085 |
| q50 | 9.847 | 2.108 | 9.906 | 2.106 | 9.773 | 2.109 |
| q75 | 10.995 | 2.676 | 10.992 | 2.674 | 11.00 | 2.679 |
| q90 | 12.047 | 2.227 | 12.034 | 2.210 | 12.063 | 2.249 |
| *Inequality measures* | | | | | | |
| lnq90–lnq10 | 4.853 | 3.804 | 4.672 | 3.612 | 5.078 | 4.019 |
| lnq75–lnq25 | 2.393 | 3.291 | 2.269 | 3.281 | 2.548 | 3.296 |
| lnq50–lnq10 | 2.653 | 3.313 | 2.544 | 3.158 | 2.788 | 3.493 |
| lnq90–lnq50 | 2.200 | 2.518 | 2.128 | 2.528 | 2.290 | 2.502 |
| Education years | 9.15 | 4.314 | 8.443 | 4.491 | 10.032 | 3.908 |
| Age | 36.044 | 10.542 | 35.862 | 10.597 | 36.272 | 10.472 |
| Urban residency | 0.622 | 0.485 | 0.592 | 0.492 | 0.660 | 0.474 |
| Black African | 0.877 | 0.329 | 0.883 | 0.322 | 0.869 | 0.337 |
| Married | 0.289 | 0.453 | 0.302 | 0.459 | 0.273 | 0.445 |
| Male | 0.555 | 0.497 | | | | |

Source: authors' construction using Stata 15, based on NSA (2015/16).

Looking at the log wage, the mean values indicate that the overall average log wage is 9.692. Men have a slightly higher mean log wage (9.758) compared with women (9.609). Comparing the inequality measures by gender, it can be observed that women generally experience slightly higher levels of wage inequality than men in Namibia, as indicated by the higher mean values for the inequality measures. Regarding education years, the overall mean is 9.15, indicating an average of 9.15 years of education for the sample. Men have a lower mean value of 8.443 years, while women have a higher mean value of 10.032 years, indicating a gender disparity in educational attainment.

## 5 Empirical results and discussion

### 5.1 Estimates of the returns to education overall and by gender

The objective of this subsection is to estimate the returns to education overall and disaggregated by gender, while considering other correlates. Table 2 presents estimates of the structural forms of the Mincerian wage-generating function based on various assumptions. Column 1 presents the OLS estimate of the returns to education. However, this estimate is likely to be affected by potential endogeneity bias, primarily stemming from unobservable abilities that simultaneously influence both the decision to pursue higher education and labour market outcomes, such as wages. Column 2 displays the two-stage least squares (2SLS) estimates using internally generated instruments, as proposed by Lewbel (2012). Prior to conducting the estimation using the Lewbel (2012) heteroscedasticity-based identification strategy, we perform the Breusch-Pagan test as outlined in Appendix Table A2. This test serves as a prerequisite for applying the method. The results of the Breusch-Pagan test indicate that the null hypothesis of constant variance is rejected, suggesting the presence of heteroscedasticity in the data.

The results of the estimation using the OLS method reveal that the returns to education in Namibia are estimated to be around 17.8 per cent. This implies that, on average, each additional year of education is associated with an approximately 17.8 per cent increase in earnings. After accounting for endogeneity using the 2SLS approach, the estimated returns to education decrease to approximately 13.9 per cent. This adjustment suggests that the initial OLS estimate may have been influenced by the presence of endogeneity bias. The decline in the estimated returns to education from 17.8 per cent in the OLS model to 13.9 per cent in the 2SLS model indicates that the true causal effect of education on earnings may be lower than initially estimated.

Disaggregating the returns to education by gender reveals an intriguing observation: they are higher among women compared with men. Indeed, the returns to education are 14.3 per cent for men and 22 per cent for women, indicating that women tend to experience a greater increase in earnings for each additional year of education completed compared with their male counterparts. The higher returns to education for women compared with men in Namibia are likely due to the fact that women, on average, have higher educational attainment levels than men, as indicated by the descriptive statistics in Table 1. When women have more years of schooling or pursue higher levels of education, they are more likely to secure higher-paying jobs that require advanced qualifications. As a result, women in Namibia experience higher returns to education, as they can access better employment opportunities that offer higher wages and greater career prospects. These findings are consistent with the results obtained by Lee and Ihm (2020) using data from Korea. Their study also found that women experience higher returns to education than men. This suggests that the pattern of higher returns to education for women is not unique to Namibia but can be observed in other countries as well. However, it is important to note that results obtained by Huang et al. (2022) in China contradict the findings observed in Namibia and Korea. Their study found lower returns to education for women than for men. This discrepancy highlights the importance of considering the specific context and country-level factors when examining the returns to education by gender.

Table 2: Returns to education overall and by gender (dependent variable: log wage)

| *Variables* | Overall<br>(1)<br>OLS | <br>(2)<br>2SLS | By gender<br>(3)<br>Men | <br>(4)<br>Women |
|---|---|---|---|---|
| Education | 0.1775*** | 0.1389*** | 0.1426*** | 0.2204*** |
| | (0.0043) | (0.0132) | (0.0192) | (0.0188) |
| Age | 0.0638*** | 0.0707*** | 0.0699*** | 0.0595*** |
| | (0.0114) | (0.0117) | (0.0155) | (0.0176) |
| Age squared | −0.0003** | −0.0005*** | −0.0004** | −0.0002 |
| | (0.0001) | (0.0002) | (0.0002) | (0.0002) |
| Urban residency | 0.5672*** | 0.6621*** | 0.7000*** | 0.4372*** |
| | (0.0377) | (0.0488) | (0.0717) | (0.0667) |
| Black African | −0.8139*** | −0.8945*** | −0.8124*** | −0.8170*** |
| | (0.0521) | (0.0584) | (0.0832) | (0.0816) |
| Married | 0.1827*** | 0.2322*** | 0.1163* | 0.2672*** |
| | (0.0414) | (0.0446) | (0.0615) | (0.0648) |
| Male | 0.4541*** | 0.4022*** | | |
| | (0.0346) | (0.0387) | | |
| Constant | 6.2470*** | 6.5535*** | 6.8747*** | 5.9101*** |
| | (0.2148) | (0.2374) | (0.2991) | (0.3610) |
| Observations | 7,274 | 7,274 | 4,038 | 3,236 |
| R-squared | 0.3570 | 0.3500 | 0.3284 | 0.4086 |

Note: standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Source: authors' construction based on 2015/16 NHIES and Stata 15.

In addition, the preferred 2SLS results indicate that age has an inverted U-shaped effect on the log of overall wages, for men and for women. This finding suggests that the relationship between age and wages follows a curvilinear pattern. For the overall sample, the inverted U-shaped effect of age on wages suggests that individuals experience an initial increase in wages as they age, reaching a peak at around 71 years, and then a subsequent decline. This pattern can be attributed to several factors. In the early stages of their career, individuals typically accumulate experience and skills, which can lead to higher wages. As they gain more knowledge and expertise, their productivity and value in the labour market tend to increase, resulting in higher wages. However, as individuals approach a certain age, other factors come into play. Physical abilities may start to decline, impacting job performance and productivity. Additionally, labour market dynamics, such as changing demand for certain skills or preferences for younger workers, may contribute to a decrease in wages for older people.

The results indicate that there is a wage premium associated with urban residency for the overall sample and for both the men and women subsamples. This suggests that individuals residing in urban areas tend to earn higher wages compared with those living in rural areas. The urban wage premium can be attributed to several factors. Urban areas typically offer a wider range of job opportunities, including higher-skilled and higher-paying jobs, due to the concentration of industries, businesses, and economic activities. Additionally, urban areas often provide better access to educational and training institutions, which can lead to the acquisition of specialized skills and qualifications that are in demand in the labour market.

The results indicate that individuals belonging to the black African Namibian population experience a wage penalty. This means that, on average, individuals of African descent earn lower wages than those of other ethnic groups in the country. The results also show that being married and being male are associated with a wage premium in Namibia. This suggests that, on average, married individuals and men tend to earn higher wages than their unmarried counterparts and women, respectively.

### 5.2 Returns to education across the unconditional wage distribution, overall and by gender

Table 3 presents a summary of the 2SLS unconditional quantile regression (UQR) estimates of the returns to education at different percentiles. The ordinary UQR results are presented in Appendix Table A1. The results are reported for the overall sample, as well as separately for men and women. For the overall sample, the estimated returns to education increase as we move from lower to higher percentiles. At the 10th percentile, the returns to education are estimated at about 15.9 per cent, indicating that individuals at the lower end of the wage distribution experience a smaller increase in earnings for each additional year of education. However, as we move towards higher percentiles, i.e. the 50th percentile (median) and beyond, the returns to education become more substantial. At the 50th percentile, the estimated returns to education are 17.4 per cent, indicating a greater increase in earnings for individuals at the median wage level. Similarly, at the 75th and 90th percentiles, the estimated returns to education are 19.5 per cent and 18.4 per cent respectively, suggesting even greater increases in earnings for individuals at the upper end of the wage distribution.

Table 3: UQR estimates of the returns to education

| | | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|---|
| *Variables* | | 10th_per | 25th_per | 50th_per | 75th_per | 90th_per |
| Overall | Returns to education | 0.1589*** | 0.1662*** | 0.1743*** | 0.1950*** | 0.1844*** |
| | | (0.0261) | (0.0229) | (0.0149) | (0.0181) | (0.0160) |
| | Pseudo R-squared | 0.0663 | 0.1099 | 0.2410 | 0.3026 | 0.2153 |
| Men | Returns to education | 0.0806** | 0.1321*** | 0.1719*** | 0.1661*** | 0.1860*** |
| | | (0.0385) | (0.0360) | (0.0236) | (0.0288) | (0.0251) |
| | Pseudo R-squared | 0.0591 | 0.0900 | 0.2426 | 0.3033 | 0.2248 |
| Women | Returns to education | 0.2892*** | 0.2433*** | 0.1841*** | 0.2534*** | 0.2207*** |
| | | (0.0387) | (0.0319) | (0.0204) | (0.0249) | (0.0225) |
| | Pseudo R-squared | 0.1054 | 0.1680 | 0.2739 | 0.3278 | 0.2198 |

Note: standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Source: authors' construction using Stata 15, based on NSA (2015/16).

For men, the estimated returns to education are positive and statistically significant at the 10th, 25th, 50th, 75th, and 90th percentiles. This suggests that regardless of their position in the wage distribution, men experience an increase in earnings for each additional year of education completed. However, it is worth noting that the magnitude of the estimated returns varies across percentiles. The returns to education are relatively lower at the 10th and 25th percentiles (0.0806 and 0.1321, respectively) compared with the higher percentiles (0.1719, 0.1661, and 0.1860 at the 50th, 75th, and 90th percentiles, respectively). This indicates that men at median and higher wage levels tend to benefit more from education in terms of increased earnings.

For women, the estimated returns to education are also positive and statistically significant at all percentiles, but the pattern differs from that of men. At the lower percentiles (10th and 25th), women experience higher returns to education compared with men. The estimated returns for women gradually decrease as the percentiles increase, indicating that the impact of education on earnings diminishes for women at the higher wage levels. Specifically, at the 10th percentile the returns to education are 0.2892, while at the 90th percentile they are 0.2207. These findings suggest that women, particularly those at the lower percentiles of the wage distribution, tend to benefit more from education in terms of increased earnings than men. However, the magnitude of the returns diminishes for women at the higher wage levels. This indicates that education may have a more significant impact on narrowing wage gaps and improving earnings outcomes for women in lower-paying jobs.

### 5.3 The effect of education on wage inequality measures

Table 4 presents the results of the effect of education on various wage inequality measures. The analysis considers the Gini coefficient and several inter-quantile ranges of the wage distribution. The Gini coefficient for education is −0.0012, indicating a negative relationship between education and Gini wage inequality. Although the coefficient is statistically significant at the 10 per cent level ($p<0.1$), the magnitude of the effect is relatively small. This suggests that higher levels of education are associated with a slight reduction in overall wage inequality, but the effect may not be substantial. The findings further indicate that the coefficient of education is positive but not significant at the 5 per cent level for the inter-quantile wage ranges lnq90–lnq10, lnq75–lnq25, and lnq50–lnq10.

However, for the inter-quantile wage range lnq90–lnq50, the coefficient of education is positive and statistically significant. This indicates that an additional year of schooling contributes to a widening of the wage gap between individuals at the 90th percentile and those at the median. In other words, individuals with higher levels of education are likely to earn significantly higher wages compared with those with lower levels of education, particularly when comparing individuals at the upper end (90th percentile) of the wage distribution with those around the median. The results suggest that education has a limited and mixed effect on wage inequality measures. While there may be a slight reduction in the Gini coefficient associated with higher education levels, the impact on wage differences between specific percentiles is less conclusive.

Table 4: Effect of education on wage inequality measures

| *Variables* | (1) Gini | (2) lnq90–lnq10 | (3) lnq75–lnq25 | (4) lnq50–lnq10 | (5) lnq90–lnq50 |
|---|---|---|---|---|---|
| Education | −0.0012* | 0.0663* | 0.0359 | 0.0176 | 0.0487** |
| | (0.0007) | (0.0346) | (0.0297) | (0.0304) | (0.0228) |
| Age | −0.0029*** | −0.1277*** | −0.0807*** | −0.0141 | −0.1136*** |
| | (0.0006) | (0.0306) | (0.0263) | (0.0269) | (0.0202) |
| Age squared | 0.0000*** | 0.0018*** | 0.0011*** | 0.0003 | 0.0015*** |
| | (0.0000) | (0.0004) | (0.0003) | (0.0004) | (0.0003) |
| Urban residency | −0.0237*** | −1.1009*** | −0.2627** | −0.2598** | −0.8411*** |
| | (0.0025) | (0.1277) | (0.1094) | (0.1119) | (0.0841) |
| Male | −0.0128*** | −0.3859*** | −0.2246*** | −0.2714*** | −0.1145* |
| | (0.0020) | (0.1012) | (0.0867) | (0.0887) | (0.0667) |
| Black African | −0.0007 | −0.1739 | −0.6934*** | −0.1477 | −0.0261 |
| | (0.0030) | (0.1530) | (0.1311) | (0.1341) | (0.1008) |
| Married | 0.0118*** | 0.6409*** | 0.6815*** | 0.4086*** | 0.2323*** |
| | (0.0023) | (0.1167) | (0.1000) | (0.1023) | (0.0769) |
| Constant | 0.1914*** | 7.2265*** | 4.1192*** | 2.9527*** | 4.2738*** |
| | (0.0123) | (0.6217) | (0.5326) | (0.5450) | (0.4095) |
| Observations | 7,274 | 7,274 | 7,274 | 7,274 | 7,274 |
| R-squared | 0.0392 | 0.0313 | 0.0283 | 0.0095 | 0.0256 |

Note: standard errors in parentheses. p<0.01, ** p<0.05, * p<0.1.

Source: authors' construction using Stata 15, based on NSA (2015/16).

### 5.4 The impact of educational equalization on wage inequality

Table 5 provides an analysis of the inequality impacts that would occur if years of education were equalized at their mean values. The table presents various inequality measures and their corresponding impacts on inequality, comparing the factual (actual) distribution with the counterfactual distribution where education levels are equalized. The measures of inequality examined include the Gini index, Atkinson index with different epsilon values (ε = 0.5, ε = 1, ε = 2), and the inter-quantile ranges for various percentile intervals. For the Gini coefficient, the factual value is 0.6716 while the counterfactual value after equalizing education is 0.6046. The difference, 0.0670, is statistically significant at the 1 per cent level. This indicates that equalizing years of education would lead to a substantial reduction in the Gini coefficient, suggesting a significant decrease in overall wage inequality by about 9.98 per cent.

Table 5: Impacts on wage inequality of equalizing years of education

| *Inequality measure* | Factual | Counterfactual | Inequality impact: $\Delta I$ $[\theta_I$ $per\ cent]$ |
|---|---|---|---|
| Gini coefficient | 0.6716 | 0.6046 | 0.0670*** (0.0050) |
| | (0.0065) | (0.0069) | [9.98] |
| Atkinson ($\varepsilon = 0.5$) | 0.3873 | 0.3167 | 0.0706*** (0.0139) |
| | (0.0081) | (0.0074) | [18.23] |
| Atkinson ($\varepsilon = 1$) | 0.6865 | 0.5980 | 0.0885*** (0.0287) |
| | (0.0109) | (0.0110) | [12.89] |
| Atkinson ($\varepsilon = 2$) | 0.9332 | 0.8990 | 0.0342*** (0.00001) |
| | (0.0046) | (0.0058) | [3.66] |
| lnq90–lnq10 | 4.9350 | 4.2905 | 0.6445*** (0.0340) |
| | (0.0445) | (0.0408) | [13.06] |
| lnq75–lnq25 | 2.4225 | 2.2350 | 0.1875*** (0.0109) |
| | (0.0403) | (0.0359) | [7.74] |
| lnq50–lnq10 | 2.6826 | 2.6498 | 0.0328 (0.0310) |
| | (0.0390) | (0.0370) | [1.22] |
| lnq90–lnq50 | 2.2524 | 1.6407 | 0.6117*** (0.0269) |
| | (0.0292) | (0.0247) | [27.16] |

Note: round brackets represent the standard error while square brackets indicate the relative contribution or impact; the counterfactual distribution refers to the wage distribution where years of schooling are equalized to their mean values; ΔI represents the absolute change in inequality.

Source: authors' construction using Stata 15, based on NSA (2015/16).

Similarly, the Atkinson index with $\varepsilon = 0.5$ shows a factual value of 0.3873, which decreases to 0.3167 in the counterfactual scenario. The absolute change in inequality is 0.0706, and it is statistically significant at the 1 per cent level. The Atkinson index with $\varepsilon = 1$ and $\varepsilon = 2$ also exhibit significant reductions in inequality when education is equalized.

Examining the inter-quantile wage ranges, the differences between the factual and counterfactual scenarios provide insights into the impacts of equalizing education. For example, for the lnq90–lnq10 range, the factual value is 4.9350 while the counterfactual value decreases to 4.2905. The absolute change in inequality is 0.6445, and it is statistically significant at the 1 per cent level. This suggests that equalizing education leads to a substantial decrease in wage inequality between individuals at the 90th and 10th percentiles. Similar patterns can be observed for the lnq75–lnq25 and lnq90–lnq50 ranges, where equalizing education results in significant reductions in wage inequality. However, for the lnq50–lnq10 range, the impact is relatively small and statistically insignificant.

Table 6 presents the results of the analysis on the impacts on wage inequality measures by gender of equalizing years of education. The Gini measure of inequality reveals interesting findings. Among men, the factual Gini index stands at 0.659 while the counterfactual Gini index, which equalizes years of education at mean values, is 0.589. This suggests that equalizing education would have a significant relative impact of approximately 10.6 per cent in reducing wage inequality among men. Turning to women, the factual Gini index is 0.687 whereas the counterfactual Gini index is 0.618, indicating a relative impact of around 10 per cent in reducing wage inequality. These results suggest that equalizing education opportunities has a notable impact on reducing wage inequality for both men and women.

Table 6: Impacts on wage inequality measures by gender of equalizing years of education

| *Inequality measure* | Factual | Counterfactual | Men Impact | Factual | Counterfactual | Women Impact |
|---|---|---|---|---|---|---|
| Gini index | 0.659 | 0.589 | 0.070*** (0.007) | 0.687 | 0.618 | 0.069*** (0.004) |
| | (0.008) | (0.008) | [10.62] | (0.0093) | (0.0092) | [10.04] |
| Atkinson (ATK) | | | | | | |
| ATK ($\varepsilon = 0.5$) | 0.370 | 0.300 | 0.070*** (0.018) | 0.409 | 0.330 | 0.079*** (0.014) |
| | (0.009) | (0.008) | [18.92] | (0.011) | (0.010) | [19.32] |
| ATK ($\varepsilon = 1$) | 0.662 | 0.573 | 0.089*** (0.031) | 0.715 | 0.617 | 0.098*** (0.029) |
| | (0.013) | (0.012) | [13.44] | (0.012) | (0.014) | [13.71] |
| ATK ($\varepsilon = 2$) | 0.924 | 0.891 | 0.033*** (0.032) | 0.941 | 0.901 | 0.040*** (0.050) |
| | (0.006) | (0.007) | [3.57] | (0.005) | (0.007) | [4.25] |
| Inter-quantile ranges | | | | | | |
| q90 – q10 | 4.745 | 4.162 | 0.583*** (0.046) | 5.172 | 4.451 | 0.721*** (0.034) |
| | (0.057) | (0.051) | [12.29] | (0.071) | (0.066) | [13.94] |
| q75 – q25 | 2.290 | 2.099 | 0.191*** (0.038) | 2.588 | 2.405 | 0.183*** (0.043) |
| | (0.054) | (0.047) | [8.34] | (0.061) | (0.055) | [7.07] |
| q50 – q10 | 2.560 | 2.593 | −0.033 (0.043) | 2.836 | 2.721 | 0.115*** (0.045) |
| | (0.050) | (0.046) | [1.29] | (0.062) | (0.061) | [4.06] |
| q90 – q50 | 2.185 | 1.569 | 0.616*** (0.037) | 2.336 | 1.730 | 0.606*** (0.019) |
| | (0.039) | (0.034) | [28.19] | (0.043) | (0.035) | [25.94] |

Note: round brackets denote standard errors and square brackets denote relative contribution/impact; counterfactual distribution is the wage distribution in which years of schooling are equalized at the mean values; $\Delta I$ is absolute change in inequality.

Source: authors' construction using Stata 15, based on NSA (2015/16).

For the Atkinson measure with ε = 0.5, the factual values for men and women are 0.370 and 0.409 respectively. When education is equalized at the mean values, the counterfactual values decrease to 0.300 for men and 0.330 for women. This indicates that equalizing education opportunities has a significant impact on reducing wage inequality for both genders, with a greater impact of 7.9 per cent for women compared with 7.0 per cent for men. For the Atkinson measure with ε = 1 and ε = 2, the results convey a similar message, with an important inference that the impact is higher for women than for men: the factual and counterfactual values consistently show larger reductions in wage inequality for women. This suggests that equalizing education opportunities has a relatively larger effect on reducing wage inequality among women than among men, regardless of the specific Atkinson measure used.

Examining the inter-quantile range measures, we can observe notable gender differences in the impact on wage inequality of equalizing education. For the lnq90–lnq10 range, the factual values are higher for women than for men. However, when education is equalized at the mean value, the counterfactual values decrease for both genders, indicating a reduction in wage inequality. The impact is more significant for women, with a reduction of 0.721 compared with 0.583 for men. Similarly, for the lnq75–lnq25 range the factual values show higher wage inequality for women. However, equalizing education leads to a decrease in wage inequality for both genders, with a slightly higher impact for women (0.191) than for men (0.183). In the lnq50–lnq10 range, men and women have similar factual values but equalizing education leads to a reduction in wage inequality only for women. In the lnq90–lnq50 range, men initially experience higher wage inequality, but equalizing education results in a decrease in wage inequality for both genders, with a slightly higher impact for women (0.606 reduction for women compared with 0.616 for men). These findings underscore the importance of equalizing education opportunities in reducing wage disparities, particularly benefiting women in various wage distribution ranges.

## 6 Conclusion and policy implications

The primary aim of this paper was to estimate the returns to education and examine their implications for wage inequality, both overall and by gender, using data from the 2015/16 Namibia Income and Expenditure Survey collected by the NSA. The study's specific objectives were:

1. to examine the effect of returns to education across the entire wage distribution, both overall and by gender, in order to provide insights into how education impacts wages at different points of the income distribution and whether there are any disparities between men and women;
2. to assess the impact of educational equalization on wage inequality, in order to shed light on the potential role of educational policies and interventions in reducing income disparities within the Namibian context.

To achieve these objectives, the study employed a robust methodology to estimate the Mincerian wage-generating function. Specifically, the RIF regression proposed by Firpo et al. (2009) was used. This approach allowed for the estimation of returns to education across the entire wage distribution and for the examination of the effect of education on distributional statistics such as the Gini coefficient and inter-quantile wage gaps. To address potential endogeneity concerns and ensure the validity of the results, the study incorporated the Lewbel (2012) identification strategy, which relies on exploiting heteroscedasticity in the data. This strategy helped to mitigate any biases that could arise due to endogeneity and provided more credible estimates of the returns to education.

Furthermore, the study adopted a simulation approach proposed by Bourguignon et al. (2007) and employed by Baye et al. (2023) to assess the impact on wage inequality of educational equalization. This simulation approach allowed for the exploration of a potential scenario in which educational disparities are eliminated, enabling a better understanding of how such changes might affect wage inequality in Namibia.

The findings from the mean regression estimates indicated that returns to education are generally higher among women than men. This suggests that women, on average, experience a greater increase in wages for each additional level of education attained. Moreover, the results from the UQR analysis revealed that the estimated returns to education vary across different percentiles of the wage distribution. Specifically, women tend to experience higher returns at lower percentiles, indicating that education has a more substantial impact on narrowing wage gaps for women at the lower end of the income distribution. Conversely, for men, higher returns were observed at the upper tail of the distribution compared with the lower tail. The simulation exercise indicated that education has an increasing effect on the factual distribution and a reducing effect on the counterfactual distribution. This suggests that educational expansion, relative to educational equalization, widens inequality substantially for both men and women.

These findings have important policy implications for addressing wage inequality in Namibia. First, policymakers should prioritize efforts to reduce educational disparities and ensure equal access to quality education for all individuals, regardless of gender or socioeconomic background. Additionally, the study highlights the importance of targeted interventions that address the specific needs of different groups within the population. By improving educational opportunities for disadvantaged groups, such as women and individuals from lower-income households, it is possible to narrow wage gaps and promote more equitable outcomes. For instance, policies aimed at improving the quality of education and providing support for vocational training may be particularly beneficial for individuals at the lower end of the income distribution, including women. By enhancing their skills and qualifications, these individuals can increase their earning potential, contributing to reducing wage inequality.

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## Appendix

Table A1: Ordinary UQR

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| *variables* | 10th_per | 25th_per | 50th_per | 75th_per | 90th_per |
| Education | 0.1296*** | 0.1630*** | 0.1672*** | 0.2567*** | 0.2009*** |
| | (0.0096) | (0.0084) | (0.0054) | (0.0066) | (0.0059) |
| Age | 0.1168*** | 0.1371*** | 0.0869*** | 0.0200 | −0.0434*** |
| | (0.0256) | (0.0226) | (0.0146) | (0.0177) | (0.0158) |
| Age squared | −0.0012*** | −0.0012*** | −0.0006*** | 0.0005** | 0.0011*** |
| | (0.0003) | (0.0003) | (0.0002) | (0.0002) | (0.0002) |
| Urban | 0.9657*** | 0.6204*** | 0.7828*** | 0.2759*** | −0.1070** |
| | (0.0817) | (0.0720) | (0.0466) | (0.0566) | (0.0503) |
| Male | 0.6751*** | 0.6048*** | 0.4706*** | 0.4391*** | 0.2888*** |
| | (0.0770) | (0.0679) | (0.0439) | (0.0534) | (0.0474) |
| Black African | −0.4855*** | −0.7612*** | −0.5887*** | −1.3472*** | −0.6121*** |
| | (0.1187) | (0.1046) | (0.0676) | (0.0823) | (0.0731) |
| Married | −0.1022 | 0.0141 | 0.2466*** | 0.6432*** | 0.4762*** |
| | (0.0921) | (0.0812) | (0.0525) | (0.0639) | (0.0567) |
| Constant | 2.9188*** | 3.8286*** | 5.7962*** | 7.8586*** | 10.5292*** |
| | (0.4806) | (0.4236) | (0.2739) | (0.3331) | (0.2959) |
| Observations | 7,274 | 7,274 | 7,274 | 7,274 | 7,274 |
| R-squared | 0.0789 | 0.1200 | 0.2539 | 0.3150 | 0.2196 |

Note: standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Source: authors' construction using Stata 15, based on NSA (2015/16).

Table A2: Breusch-Pagan/Cook-Weisberg test for heteroscedasticity in first-stage regressions with education as dependent variable

| | Overall | Men subsample | Women subsample |
|---|---|---|---|
| Chi2(1) test statistics | 275.75*** | 69.54*** | 188.32*** |
| P-value | 0.0000 | 0.0000 | 0.0000 |

Note: homoscedasticity (constant variance); standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

Source: authors' construction using Stata 15, based on NSA (2015/16).