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EDUCATION EXPENDITURE, TERTIARY ENROLLMENT AND ECONOMIC GROWTH IN UZBEKISTAN: AN AUGMENTED SOLOW AND ARDL APPROACH

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Abstract. This paper estimates the effects of education expenditure and tertiary enrollment on GDP per capita in Uzbekistan over 1999–2025. The augmented Solow model of Mankiw, Romer and Weil [15] provides the theoretical framework, and the empirical analysis combines ordinary least squares (OLS), the autoregressive distributed lag (ARDL) bounds test and distributed-lag models (unrestricted DL and Almon polynomial distributed lags, PDL). Gross tertiary enrollment rose from 8.6% in 2017 to 59.8% in 2025, an almost sevenfold increase, yet its effect on growth is lagged rather than immediate: the short-run effect is negative ($\gamma_1 = -6.45$; $p = 0.034$), a positive reversal appears in the second year ($\gamma_2 = +3.81$; $p = 0.096$), and the cumulative multiplier is still negative. The bounds test does not confirm long-run cointegration. Education spending has a positive effect on growth with a one-year lag (+3.40; $p = 0.078$). The results suggest that the expansion of enrollment must be matched by gains in education quality and spending efficiency.

Keywords: human capital, tertiary enrollment, education expenditure, economic growth, Mankiw–Romer–Weil model, ARDL, bounds test, distributed lags, Almon polynomial, Uzbekistan.

JEL classification: C22, I25, O15, O47.

1. Introduction

In modern growth theory, human capital, together with physical capital, is regarded as a key determinant of long-run prosperity. Since the reforms launched in 2017, Uzbekistan's higher education system has expanded at an unprecedented pace:

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according to World Bank data [25], the gross tertiary enrollment ratio rose from 8.6% in 2017 to 59.8% in 2025, an almost sevenfold increase (and a 4.6-fold increase relative to 2019). This expansion is being carried out within the Concept for the Development of the Higher Education System of the Republic of Uzbekistan until 2030, approved by Presidential Decree No. PF-5847 of 8 October 2019 [26]. At the same time, public expenditure on education has fluctuated between 4.1% and 6.3% of GDP and has not increased in line with enrollment. Under these conditions, the question of when, and how strongly, a sharp rise in enrollment affects economic growth is of considerable practical importance for the state budget, education policy and the labor market.

In theory, higher education should accelerate growth by enlarging the stock of human capital. In practice, however, the effect is not instantaneous: students partly withdraw from the labor market while they study, and the impact of graduates on productivity becomes visible only after three to five years. Moreover, a rapid increase in enrollment may be accompanied by a decline in the quality of education [9, 20]. It is therefore essential to identify the lag structure of the effect.

The aim of this study is to estimate econometrically the effects of tertiary enrollment and education expenditure on GDP per capita and its growth rate in Uzbekistan, with particular attention to the lag structure of these effects. The specific objectives are (1) to derive an empirical specification from the Mankiw–Romer–Weil (MRW) model [15]; (2) to test the stationarity of the series; (3) to estimate OLS and ARDL models and obtain the long- and short-run elasticities; (4) to trace the year-by-year profile of the enrollment effect and its cumulative multiplier with distributed-lag models; and (5) to derive policy recommendations from the results. The object of the study is the national economy and the higher education system of the Republic of Uzbekistan; its subject is the quantitative, and in particular lagged, cause-and-effect relationships among education expenditure, tertiary enrollment and GDP per capita.

The novelty of the study is twofold. To the best of our knowledge, it is the first to estimate the lagged effect of enrollment on growth over a period that covers the sharp (structural) rise in enrollment in 2017–2025, using distributed-lag and Almon polynomial distributed-lag models, and the first to document empirically a short-run “transition cost” in the form of a J-shaped lag profile. In practical terms, the results can inform education budget planning, quality assurance in higher education and policies for integrating graduates into the labor market.

The remainder of the paper is organized as follows. Section 2 reviews the literature. Section 3 presents the theoretical framework, the econometric methodology and the data.

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Section 4 reports the empirical results, Section 5 discusses them, and Section 6 concludes.

2. Literature Review

The neoclassical theory of economic growth builds on the work of Solow [22], in which long-run growth is determined by exogenous technological progress, whereas capital accumulation contributes to growth only during the transition to the steady state. Mankiw, Romer and Weil [15] augmented the Solow model with human capital and showed that the accumulation of physical and human capital, together with population growth, can explain about 80% of cross-country differences in income. In their model, the rate of investment in human capital is measured by secondary school enrollment, an approach that subsequently motivated the widespread use of enrollment indicators as proxies in empirical research.

Proponents of endogenous growth theory, Lucas [14] and Romer [21], interpreted human capital as a source that continuously drives growth itself: human capital affects the growth rate through externalities and through the production of new ideas. Using cross-country regressions, Barro [2] found a positive effect of primary school enrollment on subsequent growth. Benhabib and Spiegel [3], in turn, showed that it is the level of the human capital stock, rather than its change, that affects growth through technology adoption.

The empirical evidence, however, is mixed. Asking “where has all the education gone?”, Pritchett [20] showed that the rapid expansion of schooling in many developing countries did not accelerate growth and pointed to low education quality and the institutional environment as the reasons. Krueger and Lindahl [12] argued that such weak results stem largely from measurement error in education data. Bils and Klenow [4] showed that the direction of causality may be reversed, with expected growth raising the demand for schooling. Hanushek and Woessmann [10] demonstrated that what matters for growth is the quality of skills, as measured by international test scores, rather than years of schooling. Cohen and Soto [5] found that the effect of human capital on growth becomes statistically significant once the quality of the education data improves. Studies of higher education have likewise found that its effect is lagged and conditional. Gyimah-Brempong, Paddison and Mitiku [8] found a positive and significant effect of higher-education human capital on the growth of per capita income in African countries. Holmes [11], by contrast, concluded that the mass expansion of higher education does not automatically lead to faster growth, and Hanushek [9] stressed that raising tertiary

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enrollment alone is not a sufficient growth strategy, because quality and the skills acquired at school are of primary importance.

Methodologically, when the variables of a small-sample time series have different orders of integration, the ARDL bounds test proposed by Pesaran, Shin and Smith [19] is widely used; Narayan [16] adapted it to small samples. Econometric modelling methods and their methodological application are discussed in [27]. For estimating the lag structure of an effect with a limited number of parameters, the polynomial distributed-lag approach of Almon [1] is the classical method. For Uzbekistan, the World Bank's public expenditure review [24] noted the rapid growth of tertiary enrollment but also the relatively small share of higher education in education spending (about 10% in 2019).

In sum, the literature shows that (i) although the link between human capital and growth is theoretically well founded, (ii) its empirical effect is lagged and depends on quality, and (iii) single-country studies that examine the structure of this effect during periods of rapid enrollment growth are scarce. This paper seeks to fill precisely this gap.

3. Methodology

3.1. Theoretical framework

The study builds on the augmented Solow model of Mankiw, Romer and Weil [15]. Output is produced according to a Cobb–Douglas production function:

$$Y(t) = K(t)^{\alpha} H(t)^{\beta} [A(t) L(t)]^{1-\alpha-\beta}, \quad \alpha > 0,\ \beta > 0,\ \alpha + \beta < 1 \qquad (1)$$

where Y is output, K is physical capital, H is human capital, L is labor and A is the level of technology; α and β are the output elasticities of physical and human capital, respectively. Labor and technology grow exogenously:

$$L(t) = L(0)\, e^{nt}, \qquad A(t) = A(0)\, e^{gt} \qquad (2)$$

Denoting quantities per effective unit of labor by $\hat{k} = K/(AL)$, $\hat{h} = H/(AL)$ and $\hat{y} = Y/(AL)$, the dynamics of the two capital stocks are described by the differential equations

$$\dot{\hat{k}}(t) = s_K\, \hat{y}(t) - (n + g + \delta)\, \hat{k}(t) \qquad (3)$$

$$\dot{\hat{h}}(t) = s_H\, \hat{y}(t) - (n + g + \delta)\, \hat{h}(t) \qquad (4)$$

where s_K and s_H are the fractions of income invested in physical and human capital, respectively, and δ is the depreciation rate, assumed to be the same for both types of capital. The conditions $\dot{\hat{k}} = \dot{\hat{h}} = 0$ yield the steady-state values

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$$\hat{k}^* = \left(\frac{s_K^{1-\beta}\, s_H^{\beta}}{n+g+\delta}\right)^{\frac{1}{1-\alpha-\beta}}, \qquad \hat{h}^* = \left(\frac{s_K^{\alpha}\, s_H^{1-\alpha}}{n+g+\delta}\right)^{\frac{1}{1-\alpha-\beta}} \tag{5}$$

Substituting these values into the production function and taking logarithms gives the steady-state equation for income per capita:

$$\ln\frac{Y(t)}{L(t)} = \ln A(0) + gt + \frac{\alpha}{1-\alpha-\beta}\ln s_K + \frac{\beta}{1-\alpha-\beta}\ln s_H - \frac{\alpha+\beta}{1-\alpha-\beta}\ln(n+g+\delta) \tag{6}$$

3.2. Econometric specification

On the basis of equation (6), the baseline regression model for the time series of a single country is written as

$$\ln y_t = \beta_0 + \beta_1 \ln s_{K,t} + \beta_2 \ln s_{H,t} - \beta_3 \ln(n_t + g + \delta) + \varepsilon_t, \quad \varepsilon_t \sim \text{iid}\,(0, \sigma^2) \tag{7}$$

where y_t is real GDP per capita. Theory implies $\beta_1 = \alpha/(1-\alpha-\beta)$, $\beta_2 = \beta/(1-\alpha-\beta)$ and $\beta_3 = (\alpha+\beta)/(1-\alpha-\beta)$. To account for exogenous technological progress, the term gt in equation (6) is introduced as a deterministic trend, which yields the extended model

$$\ln y_t = \beta_0 + \beta_1 \ln s_{K,t} + \beta_2 \ln s_{H,t} - \beta_3 \ln(n_t + g + \delta) + \lambda t + \psi_1 \mathrm{D20}_t + \psi_2 \mathrm{S05}_t + \varepsilon_t \tag{8}$$

where $\mathrm{D20}_t$ is an impulse dummy for the 2020 pandemic shock and $\mathrm{S05}_t$ is a step dummy that accounts for the methodological break in the enrollment series from 2005 onwards. The restriction implied by the MRW theory, which is tested with a Wald test, is

$$H_0\colon\ \beta_1 + \beta_2 - \beta_3 = 0 \tag{9}$$

The models are first estimated by ordinary least squares (OLS); in the trend model, standard errors are computed with the heteroskedasticity and autocorrelation consistent (HAC) estimator of Newey and West [17]. Following the standard calibration in the literature [15], we set $g + \delta = 0.05$.

3.3. Unit-root tests

The order of integration of the time series is determined with the augmented Dickey–Fuller (ADF) test [6]:

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$$\Delta x_t = \mu + \tau t + \rho\, x_{t-1} + \sum_{i=1}^{m} \kappa_i\, \Delta x_{t-i} + u_t, \qquad H_0: \rho = 0 \tag{10}$$

where the number of lags m is selected by the Akaike information criterion (AIC). The results are cross-checked with the KPSS test [13], whose null hypothesis is the reverse one, namely stationarity.

3.4. ARDL bounds-testing approach

Because static OLS regressions of non-stationary series may be spurious [7], the main model is the ARDL(p, q_1, q_2, q_3) model of Pesaran and Shin [18] and Pesaran, Shin and Smith [19]:

$$\begin{aligned} \ln y_t = c_0 + \lambda t + \sum_{i=1}^{p} \phi_i \ln y_{t-i} + \sum_{j=0}^{q_1} \beta_{1j} \ln s_{K,t-j} \\ + \sum_{j=0}^{q_2} \beta_{2j} \ln s_{H,t-j} + \sum_{j=0}^{q_3} \beta_{3j} \ln(n+g+\delta)_{t-j} + \varepsilon_t \end{aligned} \tag{11}$$

For simplicity, the vector of dummy variables $\boldsymbol{\psi}'\mathbf{D}_t$ (D20, S05) is not shown in equation (11), although it is included in the estimation. The lag orders were selected by minimizing the AIC over all 54 combinations of $p \in \{1,2\}$ and $q_k \in \{0,1,2\}$, for specifications with a constant and with a constant and a trend. To test for cointegration, the model is rewritten in unrestricted error-correction form (UECM):

$$\begin{aligned} \Delta \ln y_t = c_0 + \lambda t + \pi_y \ln y_{t-1} + \sum_{k=1}^{3} \pi_k\, x_{k,t-1} + \sum_{i=1}^{p-1} \varphi_i\, \Delta \ln y_{t-i} \\ + \sum_{k=1}^{3} \sum_{j=0}^{q_k-1} \omega_{kj}\, \Delta x_{k,t-j} + \boldsymbol{\psi}'\mathbf{D}_t + \varepsilon_t \end{aligned} \tag{12}$$

where $x_1 = \ln s_K$, $x_2 = \ln s_H$ and $x_3 = \ln(n+g+\delta)$. The hypotheses of the bounds test are

$$H_0:\ \pi_y = \pi_1 = \pi_2 = \pi_3 = 0 \quad \text{against} \quad H_1:\ \pi_y \neq 0 \ \vee\ \pi_k \neq 0 \tag{13}$$

The F-statistic is compared with the $I(0)$ and $I(1)$ critical bounds tabulated by Pesaran et al. [19] (Case V: unrestricted intercept and unrestricted trend); in addition, the t-bounds test is applied to π_y. The long-run coefficients are computed as

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$$\theta_k = -\frac{\pi_k}{\pi_y} = \frac{\sum_{j=0}^{q_k} \beta_{kj}}{1 - \sum_{i=1}^{p} \phi_i}, \qquad k = 1,2,3 \tag{14}$$

and their standard errors are obtained by the delta method, $\mathrm{Var}(\hat{\theta}_k) \approx \nabla\theta_k{}' \, \hat{\mathbf{V}} \, \nabla\theta_k$, where $\hat{\mathbf{V}}$ is the estimated covariance matrix of the ARDL coefficients. The short-run dynamics and the speed of adjustment towards equilibrium are described by the error-correction model (ECM):

$$\Delta \ln y_t = c + \eta \, \mathrm{ECT}_{t-1} + \sum_{k=1}^{3} \sum_{j=0}^{q_k - 1} \omega_{kj} \, \Delta x_{k,t-j} + \boldsymbol{\psi}' \mathbf{D}_t + \varepsilon_t \tag{15}$$

where the error-correction term is defined as $\mathrm{ECT}_t = \ln y_t - \sum_{k=1}^{3} \hat{\theta}_k \, x_{k,t} - \hat{\theta}_4 \, t$ and $\eta = \pi_y = \sum \phi_i - 1$ is the speed-of-adjustment coefficient. If cointegration exists, η must satisfy $-1 < \eta < 0$ and be statistically significant.

3.5. Distributed-lag models

The main hypothesis of the study, namely that enrollment growth affects economic growth with a lag, is tested with a finite distributed-lag model:

$$g_{y,t} = \alpha + \sum_{j=0}^{L} \gamma_j \, \Delta \ln s_{H,t-j} + \beta \, \Delta \ln s_{K,t} + \psi \, \mathrm{D20}_t + u_t \tag{16}$$

where $g_{y,t} = 100 \cdot \Delta \ln y_t$ is the growth rate of GDP per capita in percent and γ_j is the effect of the change in enrollment j years earlier on current growth (the lag coefficient). The overall strength of the effect is measured by the interim and cumulative multipliers

$$M_k = \sum_{j=0}^{k} \gamma_j \,, \qquad M_L = \sum_{j=0}^{L} \gamma_j \,, \qquad \mathrm{se}(\widehat{M}_k) = \sqrt{\mathbf{r}_k{}' \, \hat{\mathbf{V}}_{\mathrm{HAC}} \, \mathbf{r}_k} \tag{17}$$

where $\mathbf{r}_k$ is a selection vector with ones in the positions of $\gamma_0, \ldots, \gamma_k$ and zeros elsewhere, and $\hat{\mathbf{V}}_{\mathrm{HAC}}$ is the Newey–West covariance matrix of the estimates. To reduce multicollinearity among the lag coefficients in a small sample, the polynomial distributed-lag (PDL) restriction of Almon [1] is imposed:

$$\gamma_j = a_0 + a_1 j + a_2 j^2, \qquad j = 0,1, \ldots, L \tag{18}$$

Substituting this restriction into equation (16) yields a model in new variables that is estimated by OLS:

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$$g_{y,t} = \alpha + a_0 Z_{0t} + a_1 Z_{1t} + a_2 Z_{2t} + \beta\, \Delta\ln s_{K,t} + \psi\, \mathrm{D20}_t + u_t,$$
$$Z_{rt} = \sum_{j=0}^{L} j^r\ \Delta\ln s_{H,t-j} \quad (19)$$

Because the expenditure data cover a short period (2011–2023), the effect of education expenditure is estimated separately by a regression with a one-year lag:

$$g_{y,t} = \alpha + \zeta \ln \mathrm{EDU}_{t-1} + \psi\, \mathrm{D20}_t + u_t \quad (20)$$

3.6. Diagnostic tests

The residuals of the models were examined with the Breusch–Godfrey test (autocorrelation), the Breusch–Pagan and ARCH tests (heteroskedasticity), the Jarque–Bera test (normality) and the Ramsey RESET test (functional form). All computations were carried out in Python 3 with the statsmodels library.

3.7. Data

The study uses annual data for 1999–2025 from the World Bank's World Development Indicators (WDI) database [25] (Table 1). The WDI series on education expenditure covers 2013–2023; the values for 2011–2012 were taken from the UNESCO Institute for Statistics (UIS) [23].

Table 1. Variables, definitions and sources

Symbol	Indicator	WDI code	Period
y	GDP per capita, constant 2015 US$	NY.GDP.PCAP.KD	1999–2025
s_H	Gross tertiary enrollment ratio, %	SE.TER.ENRR	1999–2025
EDU	Government expenditure on education, % of GDP	SE.XPD.TOTL.GD.ZS	2011–2023
s_K	Gross capital formation (capital investment), % of GDP	NE.GDI.TOTL.ZS	1999–2025
n	Population growth, annual %	SP.POP.GROW	1999–2025
$g + \delta$	Sum of the rates of technological progress and depreciation	MRW [15] calibration	0.05

Source: compiled by the authors from World Bank WDI [25] and UNESCO UIS [23].

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4. Results

4.1. Descriptive analysis

Figure 1 shows the dynamics of the main indicators. Real GDP per capita rose from US$1,363 in 1999 to US$4,733 in 2025, a 3.5-fold increase, with average annual growth of 4.7%. Tertiary enrollment, by contrast, followed a two-phase pattern: in 1999–2016 it declined from 13% to about 8% (the sharp drop in 2005 is partly due to a change in the UIS methodology), whereas after 2017 it grew exponentially, from 8.6% to 59.8%.

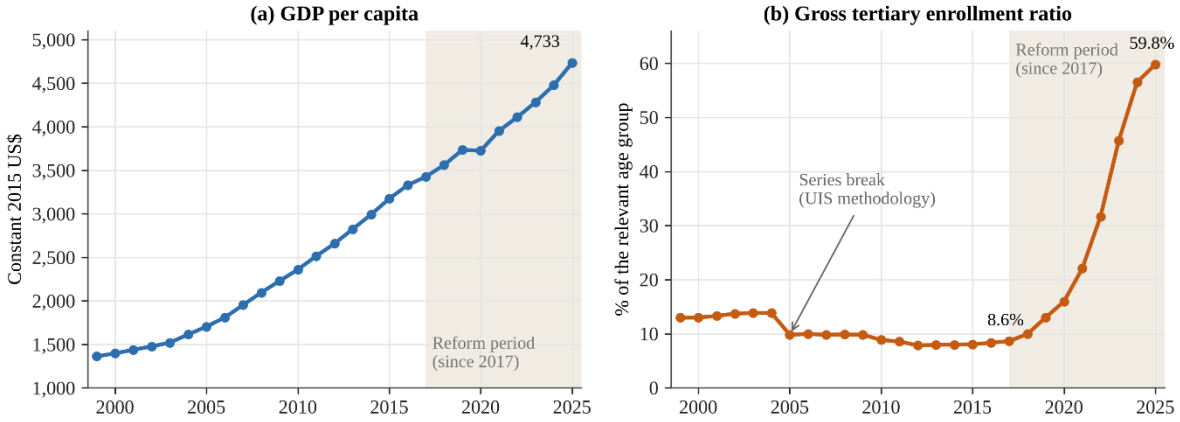


Figure 1. Dynamics of GDP per capita and gross tertiary enrollment in Uzbekistan, 1999–2025

Figure 2 shows that the share of gross capital formation rose to 33–39% of GDP after 2018, whereas education expenditure fluctuated between 4.1% and 6.3% of GDP in 2011–2023 without a clear upward trend. Thus, while enrollment increased sevenfold, the resources allocated to education remained relatively stable, which implies a decline in expenditure per student.

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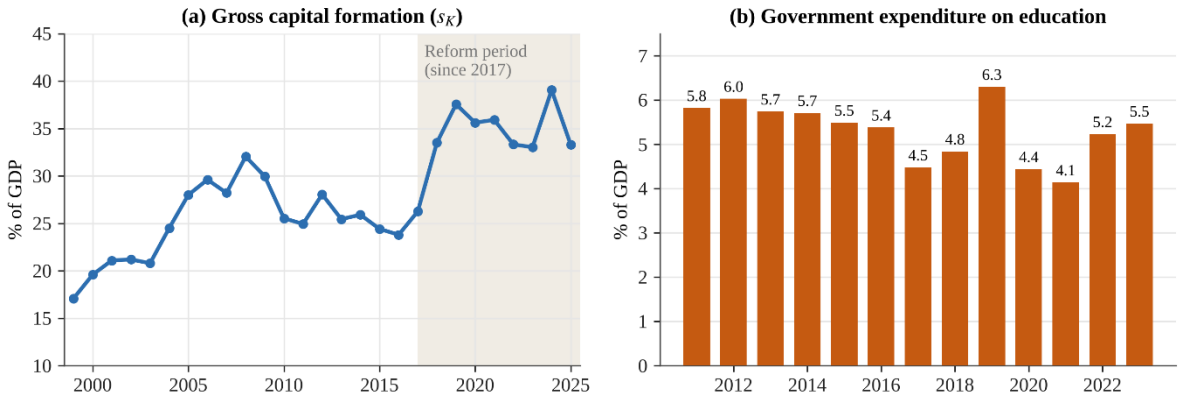


Figure 2. Gross capital formation and government expenditure on education, % of GDP

Source: World Bank WDI [25] (NE.GDI.TOTL.ZS, SE.XPD.TOTL.GD.ZS); 2011–2012: UNESCO UIS [23].

Table 2. Descriptive statistics of the variables (1999–2025)

Variable	Obs.	Mean	Std. dev.	Min.	Max.
y — GDP per capita, US$ (constant 2015)	27	2,756.83	1,066.92	1,362.51	4,732.68
s_K — gross capital formation, % of GDP	27	28.07	5.80	17.10	39.05
s_H — gross tertiary enrollment, %	27	16.71	14.50	7.85	59.81
n — population growth, %	27	1.59	0.27	1.22	2.02
EDU — education expenditure, % of GDP	13	5.31	0.66	4.14	6.29
g_y — growth of GDP per capita, %	27	4.72	1.75	−0.26	7.65

4.2. Unit-root tests

According to Table 3, $\ln y$, $\ln s_K$ and $\ln s_H$ are non-stationary in levels (the ADF null hypothesis is not rejected at the 5% level), and $\ln s_K$ is stationary in first differences, whereas $\ln(n + g + \delta)$ is stationary around a trend. The variables are thus a mixture of $I(0)$ and $I(1)$ processes, and the KPSS test confirms that none of them is $I(2)$; this justifies the use of the ARDL bounds approach.

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Table 3. Unit-root test results

Variable	ADF, level (c + t)	ADF, Δ (c)	KPSS, level (c + t)	KPSS, Δ (c)	Conclusion
$\ln y$	−3.149*	−2.165	0.144*	0.142	I(1)/trend-stationary
$\ln s_K$	−2.472	−4.302***	0.081	0.156	I(1)
$\ln s_H$	−2.600	−2.063	0.189**	0.491**	I(1)
$\ln(n+g+\delta)$	−3.538**	−2.571*	0.063	0.138	I(0)/I(1)

Note: ADF test statistics are reported; *, ** and *** denote rejection of the null hypothesis at the 10%, 5% and 1% levels, respectively. For the KPSS test, the asterisks denote rejection of the null hypothesis of stationarity. c — constant; t — trend.

4.3. OLS estimates of the MRW model

Table 4 reports the estimates of equations (7) and (8). In Model 1a, the coefficient on $\ln s_K$ is positive, in line with theory (0.324; p = 0.049), whereas the coefficient on $\ln s_H$ is insignificant (−0.043; p = 0.347). However, the coefficient on $\ln(n+g+\delta)$ is strongly positive (8.70), contrary to theoretical expectations, and the Durbin–Watson statistic (0.60) is much lower than R^2 (0.93). By the Granger–Newbold criterion [7] (R^2 > DW), this is a symptom of spurious regression: the variables are correlated because of a common trend. In Model 1b, which includes a trend, the coefficients on $\ln s_K$ and $\ln s_H$ turn negative, and the MRW restriction (9) is rejected by the Wald test (χ^2 = 6.76; p = 0.009). Hence, the static levels regression does not support the theoretical restrictions of the augmented Solow model in Uzbekistan's short time series, and a dynamic (ARDL) approach is required.

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Table 4. OLS estimates of the augmented Solow (MRW) model (dependent variable: $\ln y_t$)

Regressor	**Model 1a: OLS**	**Model 1b: OLS + trend, HAC**
Constant (β_0)	30.541*** (2.708)	10.301*** (0.814)
$\ln s_K$ (β_1)	0.324** (0.156)	−0.112*** (0.026)
$\ln s_H$ (β_2)	−0.043 (0.044)	−0.058*** (0.009)
$\ln(n+g+\delta)$ ($-\beta_3$)	8.696*** (0.851)	0.973*** (0.301)
Trend t (g)	—	0.051*** (0.001)
S05 (series break)	—	0.007 (0.015)
R^2 / adjusted R^2	0.929 / 0.920	0.999 / 0.998
Durbin–Watson	0.602	1.327
AIC / BIC	−36.31 / −31.13	−137.03 / −129.25
Wald: $\beta_1+\beta_2-\beta_3=0$ (χ^2, p)	—	6.758 (0.009)
Observations	27	27

4.4. ARDL estimates and bounds test

The AIC selected ARDL(1,2,2,1) for the specification with a constant (AIC = −157.15) and ARDL(1,2,2,2) for the specification with a trend. Since the term gt of the MRW model requires a trend, the trend-augmented ARDL(1,2,2,2) model was adopted as the main specification (AIC = −158.34; n = 25; 2001–2025). The model describes the dynamics of $\ln y$ with very high accuracy (R^2 = 0.9998; Figure 3), and the residuals remain within the $\pm 2\sigma$ band.

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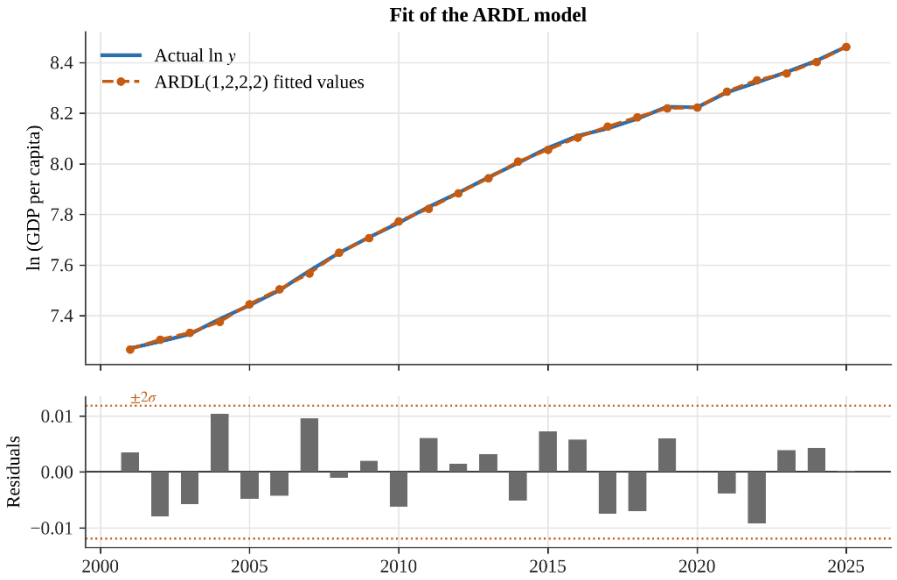


Figure 3. Actual and fitted values of the ARDL(1,2,2,2) model and its residuals

According to the bounds test in Table 5, the statistic F_{PSS} = 3.43 lies below even the lower $I(0)$ bound at the 10% significance level (3.47), and t_{PSS} = −0.81 is far below the critical values in absolute terms. Consequently, a stable long-run (cointegrating) relationship among $\ln y$, $\ln s_K$, $\ln s_H$ and $\ln(n + g + \delta)$ is not supported statistically over the period considered.

Table 5. ARDL bounds test results

Test statistic	Value	10%: I(0) / I(1)	5%: I(0) / I(1)	1%: I(0) / I(1)	Conclusion
F_{PSS} (k = 3, Case V)	3.432	3.47 / 4.45	4.01 / 5.07	5.17 / 6.36	No cointegration
t_{PSS} (ECT)	−0.809	−3.13 / −3.84	−3.41 / −4.16	−3.96 / −4.73	No cointegration

Accordingly, the long-run coefficients in Table 6 are reported for information only: the elasticity with respect to s_K is positive (0.30) and that with respect to s_H is small and negative (−0.046), but neither is statistically significant. The only significant long-run component is the trend (θ_4 = 0.048; $p < 0.001$), that is, exogenous growth of 4.8% per year on average, which corresponds to the parameter g of the MRW model.

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Table 6. Long-run coefficients of the ARDL(1,2,2,2) model (delta method)

Variable	Coefficient	Std. error	t-statistic	p-value
θ_1: $\ln s_K$	0.301	0.581	0.52	0.615
θ_2: $\ln s_H$	−0.046	0.048	−0.95	0.365
θ_3: $\ln(n+g+\delta)$	2.704	1.965	1.38	0.196
θ_4: trend (g)	0.048***	0.008	6.13	0.000

The short-run dynamics (Table 7) reveal an important result: the current change in enrollment ($\Delta\ln s_{H,t}$) has a negative effect on growth that is significant at the 5% level (−0.096; p = 0.039). This means that in a year in which enrollment rises by 10%, the growth of GDP per capita slows by approximately 0.96 percentage points. The 2020 pandemic shock reduced growth by 4.9 percentage points. The error-correction term (ECT) is negative (−0.207) but insignificant, which confirms the conclusion of the bounds test.

Table 7. Short-run coefficients and the error-correction term (UECM; dependent variable: $\Delta\ln y_t$)

Regressor	Coefficient	Std. error	t-statistic	p-value
ECT_{t-1} $(\phi-1)$	−0.207	0.256	−0.81	t-bounds: insignificant
$\Delta\ln s_{K,t}$	0.010	0.029	0.35	0.730
$\Delta\ln s_{K,t-1}$	−0.053	0.033	−1.58	0.142
$\Delta\ln s_{H,t}$	−0.096**	0.041	−2.35	0.039
$\Delta\ln s_{H,t-1}$	−0.063	0.039	−1.62	0.134
$\Delta\ln(n+g+\delta)_t$	1.491	0.865	1.72	0.113
$\Delta\ln(n+g+\delta)_{t-1}$	−0.599	0.494	−1.21	0.250
D20 (COVID-19)	−0.049***	0.015	−3.32	0.007
S05 (series break)	−0.022	0.019	−1.18	0.263
R^2 ($\Delta\ln y$) / adjusted R^2	0.880 / 0.738			

Note: ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

The diagnostic tests in Table 8 confirm that the residuals of the ARDL model are normally distributed and homoskedastic and that the functional form is correctly specified. The Breusch–Godfrey test detects no autocorrelation at one lag, whereas at

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two lags it is marginally significant at the 5% level; given the small sample (n = 25) and the 14 estimated parameters, these results should be interpreted with caution.

Table 8. Diagnostic tests of the ARDL(1,2,2,2) model

Test	Null hypothesis	Statistic	p-value	Conclusion
Breusch–Godfrey LM (1 lag)	No autocorrelation	1.294	0.282	Not rejected
Breusch–Godfrey LM (2 lags)	No autocorrelation	4.596	0.042	Rejected at 5%
Breusch–Pagan	Homoskedasticity	0.822	0.636	Not rejected
ARCH LM (1 lag)	No ARCH effects	0.369	0.550	Not rejected
Jarque–Bera	Normality	1.643	0.440	Not rejected
Ramsey RESET	Correct specification	3.002	0.114	Not rejected

4.5. Distributed-lag models: the lag profile of the effect

The main result of the study is presented in Table 9 and Figure 4. In the DL(2) model (2008–2025; n = 18), the contemporaneous effect of the change in enrollment is insignificant (γ_0 = −0.50; p = 0.746), the effect at a one-year lag is negative and significant (γ_1 = −6.45; p = 0.034), and the effect at a two-year lag turns positive (γ_2 = +3.81; p = 0.096). In other words, a 10% increase in enrollment slows growth by 0.64 percentage points in the following year and accelerates it by 0.38 percentage points in the second year. This is a classic J-shaped lag profile. The cumulative multiplier is M_2 = −3.14 (p = 0.003), which indicates that the short-run negative effect has not yet been fully offset by the subsequent positive effect.

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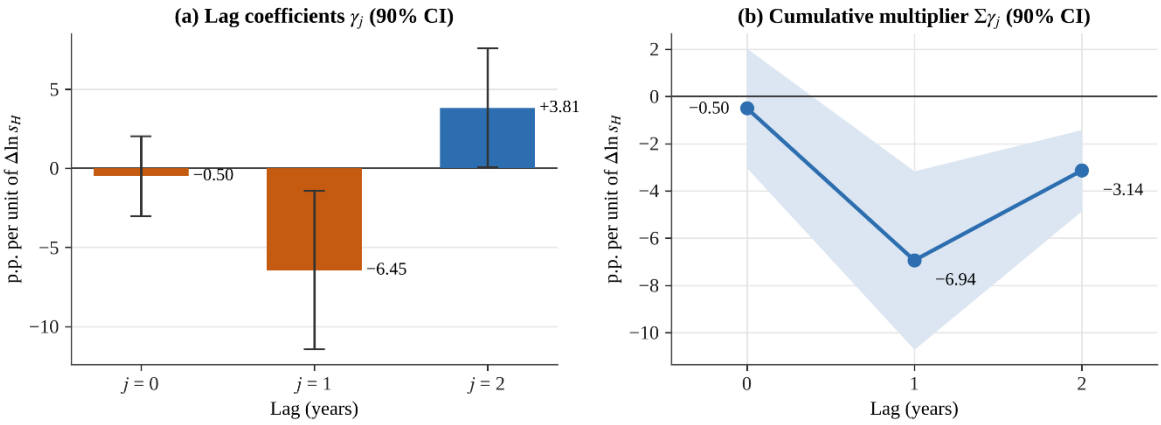


Figure 4. Lagged effect of changes in tertiary enrollment on economic growth: lag coefficients and the cumulative multiplier

Table 9. Estimates of the distributed-lag models (dependent variable: $g_{y,t}$, %)

Regressor	DL(2)	DL(3)	DL(4)
Constant (α)	5.542*** (0.286)	5.384*** (0.260)	5.306*** (0.306)
γ_0: $\Delta \ln s_{H,t}$	−0.496 (1.534)	−0.433 (1.754)	−1.166 (1.463)
γ_1: $\Delta \ln s_{H,t-1}$	−6.446** (3.041)	−4.482 (3.003)	−3.963 (2.443)
γ_2: $\Delta \ln s_{H,t-2}$	3.805* (2.284)	2.969 (3.087)	5.261 (3.399)
γ_3: $\Delta \ln s_{H,t-3}$	—	−0.784 (2.307)	−0.025 (2.232)
γ_4: $\Delta \ln s_{H,t-4}$	—	—	−3.988 (3.097)
$\Delta \ln s_{K,t}$	−1.925 (2.245)	−3.399* (1.941)	−3.310* (1.724)
D20	−4.605*** (0.344)	−4.921*** (0.418)	−5.040*** (0.439)
Cumulative multiplier $\Sigma\gamma_j$	−3.138*** (1.044)	−2.730** (1.121)	−3.880*** (1.324)
R^2 / adjusted R^2	0.750 / 0.646	0.803 / 0.685	0.818 / 0.658
Durbin–Watson	1.081	1.594	1.237
Period; n	2008–2025; 18	2009–2025; 17	2010–2025; 16

The results are robust to the lag length: in the DL(3) and DL(4) models, too, the first lag is negative and the second positive, while the cumulative multiplier ranges from −2.73 to −3.88 and is always significant. The Almon PDL(4,2) model (Table 10, Figure 5) smooths the lag coefficients and confirms the same structure: a negative effect in the

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current year ($w_0 = -2.65$), positive effects at lags 1–3 (+0.64, +1.58 and +0.15) and again a negative value at lag 4. This last estimate is imprecise (standard error 3.45) and mainly reflects the small number of observations at the end of the sample.

Table 10. Lag weights from the Almon PDL(4,2) model (2010–2025; n = 16; R^2 = 0.793)

Lag j	**0**	**1**	**2**	**3**	**4**	Σw_j
w_j (Almon)	−2.65	0.64	1.58	0.15	−3.63	−3.91
Std. error	1.35	1.50	1.81	0.40	3.45	1.42

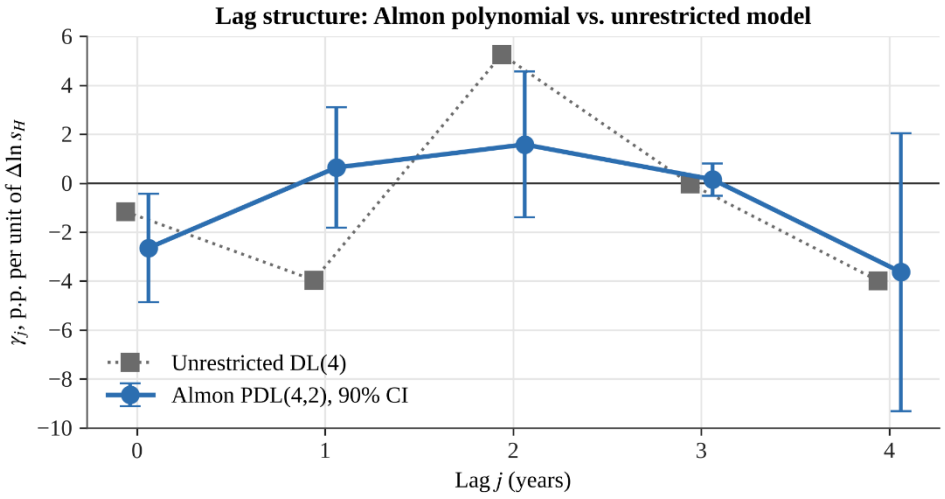


Figure 5. Lag structure: comparison of the Almon polynomial and the unrestricted DL(4) models

4.6. Education expenditure

According to Table 11, the lagged level of education expenditure (as a share of GDP) has a positive effect on current growth (ζ = 3.40; p = 0.078): if the expenditure share increases by 10% (for example, from 5.0% to 5.5% of GDP), the growth of GDP per capita in the following year is approximately 0.34 percentage points higher. Because the number of observations is small (n = 13), this result should be regarded as preliminary.

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Table 11. Effect of education expenditure on economic growth (dependent variable: $g_{y,t}$, %)

Regressor	Coefficient	Std. error (HAC)	p-value
Constant	−0.760	3.181	0.811
$\ln EDU_{t-1}$	3.396*	1.925	0.078
D20	−5.747***	0.475	0.000
R^2; n	0.727; 13 (2012–2024)		

Note: ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

4.7. Summary of the findings

The empirical results reveal the following patterns: (1) physical capital accumulation has a positive effect in the levels regression, but this effect is imprecisely estimated in the short time series; (2) the expansion of tertiary enrollment affects growth not immediately but with a lag and in a J-shaped pattern, negatively at first and positively from the second year onwards; (3) in the data up to 2025, the cumulative effect is still negative, and long-run cointegration is not confirmed; and (4) education expenditure has a positive effect on growth with a one-year lag.

5. Discussion

5.1. Interpretation of the results

The short-run negative effect of enrollment growth can be explained by several mechanisms. The first is a *labor supply effect*: when the number of students rises sharply, some young people of working age temporarily withdraw from production or move to part-time employment. The second is a *resource reallocation effect*: households divert funds from consumption and investment to pay tuition fees, while the resources available per student decline because the government has not raised the share of education expenditure. The third is a *quality effect*: when enrollment grows rapidly, the teaching staff, classroom capacity and laboratory facilities do not expand at the same pace, which affects the productivity of graduates. The positive reversal in the second year can be explained by part-time (correspondence) and evening students, who account for a substantial share of enrollment, applying their knowledge in the workplace, and by the gradual accumulation of the human capital stock. Given that bachelor's programs last four years, graduates of the "big wave" admitted in 2019–2021 began to enter the labor

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market only in 2023–2025; therefore, the full positive effect may not yet be reflected in the data.

5.2. Comparison with previous studies

The insignificant or negative coefficient on $\ln s_H$ in the levels regression is consistent with the findings of Pritchett [20] and Krueger and Lindahl [12], who also found a weak relationship between the quantity of education and growth in within-country time variation. The lagged positive effect and the positive role of expenditure, in turn, are consistent with the conclusions of Gyimah-Brempong et al. [8] and Hanushek and Woessmann [10]: the contribution of human capital to growth depends not on quantity but on quality and on the resources with which education is provided. Unlike the cross-country results of MRW [15], the elasticity with respect to s_H is not positive in the single-country time series; the reason is that enrollment in Uzbekistan first declined for a long period and then rose sharply, so that s_H behaves not as a steady-state share but as a variable in transition. The situation that Holmes [11] and Hanushek [9] warned about, namely that the mass expansion of higher education does not automatically accelerate growth, is also observed in the case of Uzbekistan.

5.3. Policy implications

The results provide a basis for the following policy conclusions: (i) the expansion of enrollment should be accompanied by a commensurate increase in public expenditure on education, since expenditure has a positive effect on growth; (ii) monitoring of education quality (accreditation, international rankings and graduate employment indicators) should be given no lower priority than enrollment indicators; (iii) integrating graduates into the labor market through internships, dual education and cooperation with employers can shorten the lag; and (iv) budget planning should take into account that the returns to education investment materialize with a lag of two to four years.

5.4. Limitations and directions for future research

The results should be interpreted in light of the following limitations. First, the number of observations is small (25 for the ARDL model, 16–18 for the DL models and 13 for the expenditure model), which reduces the power of the tests. Second, the period of rapid enrollment growth, 2017–2025, coincides with currency liberalization, broad economic reforms and the pandemic, so the effects cannot be fully disentangled. Third, the gross enrollment ratio does not reflect the quality of education or the distinction between full-time and part-time (correspondence) study; although the methodological break of 2005 is controlled for with a dummy variable, measurement error may persist. Fourth, the possibility of reverse causality (growth → demand for education) cannot be ruled out.

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Future research could use regional panel data (14 regions), data on the number of graduates and indicators of education quality, and re-estimate the ARDL model as new observations up to 2030 become available.

6. Conclusions

1. Tertiary enrollment in Uzbekistan rose from 8.6% to 59.8% (almost sevenfold) in 2017–2025, whereas public expenditure on education remained at around 4–6% of GDP; as a result, the resources available per student declined in relative terms.
2. The static OLS estimates of the augmented Solow (MRW) model showed signs of spurious regression, and the MRW restriction was rejected (Wald $\chi^2 = 6.76$; $p = 0.009$), which points to the need for dynamic models in the time series of a single country.
3. The bounds test for the ARDL(1,2,2,2) model ($F = 3.43 < 3.47$) did not confirm long-run cointegration among the variables over 1999–2025; the only significant long-run factor is an exogenous trend of 4.8% per year.
4. The distributed-lag models revealed a J-shaped lagged effect of enrollment growth on economic growth: negative at a one-year lag ($\gamma_1 = -6.45$) and positive at a two-year lag ($\gamma_2 = +3.81$). The cumulative multiplier is still negative (−3.14), meaning that the full positive effect of the enrollment expansion has not yet materialized.
5. Education expenditure has a positive effect on growth with a one-year lag ($\zeta = 3.40$; $p = 0.078$).

Policy recommendations. The policy of expanding enrollment should be combined with measures to raise education expenditure to at least 6% of GDP, to strengthen quality monitoring and to integrate graduates effectively into the labor market; a lag of two to four years should be taken into account when assessing the efficiency of education investment; and higher education statistics (graduate employment and enrollment by mode of study) should be published openly.

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Appendix A. Data used in the study

Table A1. Data used in the stud

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Year	y, US$	s_K, %	s_H, %	n, %	EDU, %	g_y, %
1999	1,362.5	17.10	12.98	1.40	—	2.81
2000	1,397.0	19.60	13.01	1.26	—	2.50
2001	1,437.6	21.10	13.30	1.22	—	2.86
2002	1,476.5	21.20	13.74	1.23	—	2.67
2003	1,520.4	20.80	13.84	1.22	—	2.93
2004	1,613.6	24.50	13.85	1.23	—	5.96
2005	1,704.3	28.00	9.84	1.25	—	5.47
2006	1,807.9	29.61	9.96	1.29	—	5.90
2007	1,951.6	28.21	9.83	1.40	—	7.65
2008	2,095.4	32.04	9.90	1.53	—	7.11
2009	2,228.1	29.95	9.82	1.60	—	6.14
2010	2,359.6	25.54	8.86	1.58	—	5.74
2011	2,512.7	24.94	8.58	1.53	5.82	6.28
2012	2,657.1	28.06	7.85	1.51	6.03	5.59
2013	2,823.0	25.45	7.98	1.57	5.74	6.06
2014	2,990.6	25.91	7.99	1.66	5.71	5.77
2015	3,172.9	24.42	8.05	1.73	5.49	5.92
2016	3,327.9	23.80	8.34	1.73	5.38	4.77
2017	3,425.8	26.26	8.65	1.69	4.48	2.90
2018	3,559.9	33.50	9.96	1.73	4.84	3.84
2019	3,733.5	37.55	13.03	1.81	6.30	4.76
2020	3,723.8	35.62	15.97	1.87	4.44	−0.26
2021	3,952.0	35.91	22.09	1.94	4.14	5.95
2022	4,109.1	33.32	31.70	2.01	5.23	3.90
2023	4,281.2	33.03	45.76	2.02	5.47	4.10
2024	4,477.9	39.05	56.52	1.97	—	4.49
2025	4,732.7	33.28	59.81	1.88	—	5.53

Note: y — GDP per capita (constant 2015 US$); s_K — gross capital formation, % of GDP; s_H — gross tertiary enrollment ratio; n — population growth; EDU — education expenditure, % of GDP; $g_y = 100 \cdot \Delta \ln y$. *Source:* World Bank WDI [25], UNESCO UIS [23].

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