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Does Institutional Quality Condition the Output Returns to Human Capital? Time-Series Evidence from Kenya, 1963–2023

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Abstract

Institutional theory predicts that human capital contributes more to output where institutions are strong, but evidence for this conditional relationship comes mainly from multi-country panels, and no study has tested it for Kenya over a long period. This study examines whether institutional quality moderates the long-run effect of human capital on real output in Kenya from 1963 to 2023. It uses Penn World Table 11.0 data on output, capital, employment and schooling, life expectancy and a composite index of rule of law, corruption control and impartial public administration from V-Dem v16. The analysis applies the ARDL bounds testing approach with finite-sample critical values and reports the conditional marginal effect of human capital across the observed range of institutional quality. A stable long-run relationship exists only when health is included alongside schooling ($F = 4.44$, finite-sample $p = 0.021$), with about 76% of deviations corrected within a year. Schooling and health are both positively related to long-run output; a 1% rise in life expectancy is associated with a 1.1% rise in output. Institutional quality has no significant direct effect. Stronger institutions do not raise the return to schooling: the conditional elasticity is positive at every observed level of institutional quality, and the interaction is negative, significant at the 5% level in three of ten specifications. The results align Kenya with single-country evidence from Nigeria and Brazil rather than with panel studies. Health investment emerges as a direct contributor to long-run output, and governance reforms are better judged by their effects on service delivery and human capital accumulation than by an assumed amplification of returns to schooling.

Keywords: human capital; institutional quality; health; ARDL bounds test; moderation; Kenya

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1. Introduction

Kenya Vision 2030 commits the country to sustained real gross domestic product (GDP) growth of 10% a year as the route to becoming a newly industrialising, middle-income economy (Republic of Kenya, 2007). Actual growth has fallen well short of that target. Real GDP grew by 4.7% in 2024 and by 4.6% in 2025 (Kenya National Bureau of Statistics [KNBS], 2026), less than half the rate the plan requires. The shortfall is long-standing: growth averaged 7.2% in the 1970s but 2.2% in the 1990s, recovered to 7.1% in 2007 after the 2002 change of government, and fell to 0.23% in 2008 amid post-election violence, drought and the global financial crisis (Kimenyi et al., 2016). Because Vision 2030 places education, training and health within its social pillar, explaining the shortfall depends in large part on whether investment in human capital is producing the growth returns that the plan assumes. The current Fourth Medium Term Plan (2023 to 2027) pursues the Vision through the Bottom-Up Economic Transformation Agenda, with healthcare among its five priority sectors (Republic of Kenya, 2023).

Kenya is a useful setting for examining the returns to human capital because both its human capital and its institutions have changed substantially since independence. Access to schooling has expanded in successive waves, most visibly after the abolition of primary school fees in 2003, which raised enrolment among poorer pupils without substantially lowering the test scores of pupils already enrolled (Lucas & Mbiti, 2012). Institutional quality has followed a less steady path. Bigsten and Moene (1996) trace Kenya's poor growth performance to a ruling coalition that retained power through an oversized public sector, administrative corruption and the overlap of political office with private business interests. The 2010 Constitution introduced new limits on executive power and devolved government to 47 counties (Republic of Kenya, 2010), and devolution has begun to weaken the patronage-based allocation of development resources (Kanyinga, 2016). Kenya therefore combines large gains in schooling with marked changes in governance.

Growth theory gives human capital a prominent place. In the augmented Solow model, human capital enters the aggregate production function as a separate accumulable input, and adding it accounts for a large share of the variation in income across countries (Mankiw et al., 1992). In endogenous growth models, Lucas (1988) treats the accumulation of human capital as the source of sustained growth, and Romer (1990) ties the rate of growth to

the stock of human capital employed in producing new ideas. Benhabib and Spiegel (1994) argue that human capital raises growth mainly by enabling countries to adopt and create technology, rather than as a conventional factor of production. On each of these accounts, a country that raises the education and health of its workforce should, other things equal, grow faster.

The macroeconomic evidence is less conclusive than the theory. Pritchett (2001) shows that the rapid expansion of schooling in developing countries after 1960 was not matched by comparable growth in output per worker, and Hanushek and Woessmann (2008) find that cognitive skills, rather than years of schooling, predict long-run growth. In Sub-Saharan Africa, Abu Alfoul et al. (2024) report that the long-run effect of education on growth across 18 countries is statistically insignificant, which they attribute to educational exclusion and low educational quality. This weak aggregate link may reflect the conditions under which human capital is used rather than a lack of productivity.

Institutional theory describes those conditions. North (1990) defines institutions as the formal and informal rules that structure economic interaction and shape incentives. Where these rules reward rent-seeking more than production, talented people move into activities that redistribute income rather than create it (Murphy et al., 1991), and corruption lowers investment and growth (Mauro, 1995). Weak institutions can also reduce the return to human capital directly. In Uganda, Reinikka and Svensson (2004) show that most of a central government grant intended for schools was captured by local officials before it reached them. Cross-country studies identify institutional quality as a fundamental determinant of long-run income (Acemoglu et al., 2005; Rodrik et al., 2004), although Glaeser et al. (2004) argue that human capital is the more basic cause of growth and that institutions improve as populations become more educated. These arguments imply a conditional relationship in which the growth return to human capital depends on institutional quality and is higher where government is effective and corruption is contained.

Tests of this conditional relationship have produced conflicting results, and most come from cross-country panels. Dias and Tebaldi (2012) find that structural institutions shape human capital accumulation and long-run economic performance. Adams-Kane and Lim (2016) show that government effectiveness raises income partly through its effect on human capital, and Abu Alfoul et al. (2024) report that stronger institutions increase the growth effect of education in Sub-Saharan Africa. Single-country time-series studies point the other way. Raifu et al. (2021) find that institutional quality does not strengthen the relationship between human capital and growth in Nigeria, and Doré and Teixeira (2023) find that, over two centuries of Brazilian data, schooling has a lasting positive effect on growth while institutional quality does not. Evidence for Kenya treats the two influences separately. Otieno (2016) finds that education expenditure per worker raised output in the long run between 1967 and 2010, Chepkemboi and Korir (2025) find that government effectiveness and control of corruption had positive and significant effects on growth between 1997 and 2022, and Bigsten and Moene (1996) model the growth costs of rent-seeking, but none of these studies asks whether institutional quality conditions the return to human capital. Panel estimates impose one moderating effect on countries with very different institutional histories, whereas a country-specific time series traces how the return to human capital changed as one country's institutions changed, which is the estimate its policymakers need. To the authors' knowledge, no study has tested whether institutional quality moderates the relationship between human capital and economic growth in Kenya within a time-series framework.

Three measurement problems compound this gap. First, composite indices such as the Human Development Index include income per capita, so using them as the human capital variable in a regression of output builds a mechanical correlation into the estimates. Second, the Worldwide Governance Indicators begin only in 1996 (Kaufmann et al., 2011), which restricts a Kenyan time series to fewer than 30 annual observations and excludes the first three decades after independence. Third, moderation cannot be inferred from the sign or significance of an interaction coefficient alone. The quantity of interest is the marginal effect of human capital on output at observed levels of institutional quality, together with its standard error (Brambor et al., 2006).

This study addresses the gap and these three problems. It examines whether institutional quality moderates the long-run effect of human capital on real output in Kenya, using annual data for 1963 to 2023 and the autoregressive distributed lag (ARDL) bounds testing approach of Pesaran et al. (2001). Output, physical capital, employment and human capital are taken from Penn World Table 11.0, whose human capital index is built from years of schooling and assumed rates of return to education and contains no income component (Feenstra et al., 2015). Health, the second component of human capital, is measured by life expectancy at birth (Bloom et al., 2004). Institutional quality is measured with a composite of the Varieties of Democracy (V-Dem) indices of rule of law, political corruption and impartial public administration, which are available for Kenya across the whole period (Coppedge et al., 2026). Both variables are mean-centred before the interaction term is formed so that their separate coefficients remain interpretable. The study asks whether human capital and institutional quality

each affect real output in the long run, and whether the marginal effect of human capital rises with institutional quality.

The study contributes to the literature in three ways. It provides a country-specific time-series test of the interaction between institutions and human capital for Kenya, covering six decades since independence, whereas earlier evidence has come mainly from multi-country panels. It reports the conditional marginal effect of human capital across the observed range of institutional quality instead of relying on the interaction coefficient alone, showing how far governance reform changes the return to education. It also combines a human capital measure with no income component with governance data that predate 1996, and adds health to schooling, removing one source of spurious correlation and lengthening the sample available for Kenya. The results show that schooling and health both have positive long-run relationships with output, and that a stable long-run relationship exists only when health is included. Institutional quality has no significant direct long-run effect, and stronger institutions do not raise the return to schooling: the conditional elasticity of output with respect to schooling is positive at every observed level of institutional quality, while the interaction is negative and significant at the 5% level only in some specifications. The remainder of the paper is organised as follows. Section 2 reviews the literature and states the hypotheses, Sections 3 and 4 present the theory, data and method, Section 5 reports the results, Section 6 discusses them, and Section 7 concludes.

2. Literature review and hypotheses

2.1 Human capital and economic growth

The theoretical case for human capital as a source of growth takes two forms. In the first, human capital is a factor of production that is accumulated like physical capital. Mankiw et al. (1992) add it to the Solow model and show that the augmented model explains a large share of cross-country differences in income per worker, with diminishing returns to each reproducible factor still applying. In the second, human capital drives growth itself. Lucas (1988) models the accumulation of skills as an activity with external effects that sustain growth without diminishing returns, and Romer (1990) makes the rate of technological progress depend on the stock of human capital engaged in research. Benhabib and Spiegel (1994) combine these ideas with a catch-up mechanism in which the stock of human capital governs how quickly a country adopts technologies developed elsewhere. The distinction matters for estimation: the factor view predicts a link between the level of human capital and the level of output, whereas the endogenous growth view predicts a link between the level of human capital and the rate of growth.

Early cross-country evidence supported a positive association. Barro (1991) finds that initial school enrolment rates predict subsequent growth in income per capita across 98 countries. Later work weakened this conclusion. Benhabib and Spiegel (1994) find that growth in human capital enters growth regressions insignificantly and often with a negative sign, Pritchett (2001) shows that the large expansion of schooling in developing countries after 1960 was not matched by growth in output per worker, and Bils and Klenow (2000) argue that much of the correlation between schooling and growth runs from expected growth to schooling rather than the reverse. Two explanations have been offered for these weak results. Krueger and Lindahl (2001) and de la Fuente and Doménech (2006) attribute them to measurement error in cross-country schooling data, which biases estimated coefficients towards zero, and show that the effect of schooling becomes larger and more stable when data quality improves. Hanushek and Woessmann (2008) argue that years of schooling are a poor proxy for skills and find that cognitive skills, measured by international test scores, have a strong and robust association with long-run growth.

Health is the second component of human capital. Bloom et al. (2004) estimate an aggregate production function and find that life expectancy has a positive and significant effect on output, independent of schooling and experience. Microeconomic evidence on returns to education is more consistent than the macroeconomic evidence. In a review of estimates from 139 countries, Psacharopoulos and Patrinos (2018) report that the private return to an additional year of schooling averages about 9% and is highest in low-income countries and in Sub-Saharan Africa. The gap between consistently positive private returns and weak aggregate effects is one of the main puzzles in this literature, and it motivates the search for conditions under which human capital is, or is not, converted into output.

African evidence reflects the same tension. Gyimah-Brempong et al. (2006) find that all levels of education, including higher education, have positive and significant effects on the growth of income per capita in African countries. Abu Alfoul et al. (2024), using a panel of 18 Sub-Saharan countries over 2000 to 2020, find that the long-run effect of education on growth is insignificant on its own, which they attribute to high rates of educational exclusion and low educational quality. Lee and Kim (2009) find that secondary education and institutions matter for growth in lower-income countries, while technology and higher education matter more in

upper middle-income and high-income countries, which suggests that the return to a given type of human capital depends on the stage of development and the surrounding institutions.

2.2 Institutions and economic growth

New institutional economics treats institutions as the rules that structure economic, political and social interaction (North, 1990). Secure property rights, enforceable contracts and constraints on those who hold power reduce transaction costs and the risk of expropriation, and so raise the expected return to investment of all kinds. Hall and Jones (1999) show that differences in social infrastructure, meaning the institutions and government policies that determine the economic environment, account for much of the variation in output per worker across countries. Knack and Keefer (1995) find that measures of property rights security and the quality of public administration from commercial risk ratings are strongly related to growth and investment.

The main challenge in this literature is reverse causality, since richer countries can afford better institutions. Acemoglu et al. (2001) address it by using settler mortality as an instrument for current institutions and find large effects of institutions on income per capita. Rodrik et al. (2004) compare institutions, geography and trade integration in a common framework and conclude that institutional quality dominates once it is instrumented. Acemoglu et al. (2019) find that democratisation raises GDP per capita by about 20% over the following 25 years, operating partly through greater investment in schooling and health. Glaeser et al. (2004) dispute the direction of this relationship. They argue that the commonly used institutional indicators measure policy outcomes rather than durable constraints on government, and that human capital is the more basic cause of both growth and institutional improvement. This disagreement bears directly on the present study, because it implies that human capital and institutions may be substitutes, complements or stages in a single process.

Corruption is the institutional failure most often studied in Sub-Saharan Africa. Mauro (1995) finds that corruption lowers investment and therefore growth. Mo (2001) estimates that a 1% increase in the level of corruption reduces the growth rate by about 0.72%, with political instability as the most important transmission channel and lower human capital as a further one. Ndulu and O'Connell (1999) relate Africa's growth slowdown after the mid-1970s to authoritarian systems in which governing elites favoured rent extraction over broad-based development, and note faster growth after political liberalisation in the 1990s in countries that avoided conflict.

2.3 The conditional relationship between institutions and human capital

The two literatures meet in the proposition that institutions determine how productively human capital is used. Three mechanisms support this proposition. The first is allocation. Murphy et al. (1991) show that when rent-seeking pays better than production, able people move into law, the bureaucracy and other redistributive activities, and the growth return to their skills falls. The second is leakage. Reinikka and Svensson (2004) find that schools in Uganda received on average only 13% of a central government capitation grant during 1991 to 1995, with the remainder captured by local officials, so that public spending on education produced far less human capital than budgeted. The third is complementarity with investment. If corruption lowers physical investment (Mauro, 1995), educated workers are combined with less capital and their marginal product falls. Each mechanism implies that the marginal effect of human capital on output rises with institutional quality.

Empirical tests of this proposition have used three designs. The first estimates direct and indirect channels. Dias and Tebaldi (2012) model institutions as a determinant of human capital accumulation and find, for a panel covering 1965 to 2005, that structural institutions affect long-run economic performance whereas political institutions show no significant relationship with productivity or growth. Adams-Kane and Lim (2016) use a two-stage strategy in which government effectiveness affects educational outcomes, which in turn affect income, and conclude that institutional quality mediates the effect of human capital. The second design adds an interaction term to a growth regression. Ali et al. (2018), using a panel of 132 countries over 1996 to 2011, find that human capital raises per capita GDP growth only in the presence of high-quality legal institutions and better economic opportunities. Abu Alfoul et al. (2024) find that institutional quality, particularly control of corruption and political stability, strengthens the effect of education on growth in Sub-Saharan Africa. The third design uses long single-country time series. Raifu et al. (2021) find that institutional quality does not enhance the relationship between human capital and growth in Nigeria over 1984 to 2016, and Doré and Teixeira (2023) find that, over two centuries of Brazilian data, years of schooling have a positive and lasting effect on growth while institutional quality is not a significant long-run driver.

Two patterns stand out. Panel studies tend to find that institutions amplify the effect of human capital, while the few single-country studies tend not to. Panel estimates impose a common moderating effect on countries with different institutional histories, so this divergence may reflect genuine heterogeneity across countries rather than a contradiction. In addition, most of these studies infer moderation from the sign and significance of the

interaction coefficient alone. Brambor et al. (2006) show that this practice can mislead, because the effect of one variable in an interaction model is conditional on the value of the other and its significance can differ across the range of the moderator. Evidence that reports conditional marginal effects across the observed range of institutional quality is therefore scarce.

Table 1: Empirical studies of the interaction between institutions and human capital

Study	Sample and period	Design	Finding on moderation
Dias and Tebaldi (2012)	Cross-country panel, 1965 to 2005	Theoretical model and panel estimation	Structural institutions shape long-run performance; political institutions do not
Adams-Kane and Lim (2016)	Cross-country	Two-stage: institutions to education to income	Government effectiveness works partly through human capital
Ali et al. (2018)	132 countries, 1996 to 2011	Panel fixed effects with interactions	Human capital raises growth only with high-quality legal institutions
Abu Alfoul et al. (2024)	18 Sub-Saharan countries, 2000 to 2020	Panel ARDL with interactions	Institutions strengthen the education effect
Raifu et al. (2021)	Nigeria, 1984 to 2016	Time series	Institutions do not enhance the human capital effect
Doré and Teixeira (2023)	Brazil, 1822 to 2019	ARDL	Schooling matters; institutions not a significant long-run driver
This study	Kenya, 1963 to 2023	ARDL with centred interaction and conditional marginal effects	No positive moderation; interaction negative once health is included, significant at 5% in some specifications

2.4 Evidence from Kenya

Kenyan evidence on human capital and growth comes mainly from time-series studies of education. Otieno (2016) estimates a Cobb-Douglas production function for 1967 to 2010 and finds that a 1% increase in education expenditure per worker raises output by about 0.5% in the long run, with causality running from education spending to growth. Microeconomic evidence shows that education policy has expanded access. Lucas and Mbiti (2012) find that the abolition of primary school fees in 2003 increased the number of pupils completing primary school, particularly among poorer households, without substantially reducing the test scores of pupils already enrolled.

Kenyan evidence on institutions takes two forms. Political economy studies link Kenya's growth record to the character of its governing coalitions. Bigsten and Moene (1996) argue that growth slowed as the ruling coalition used an expanding public sector, administrative corruption and the fusion of political office with business interests to retain power. Kimenyi et al. (2016) associate the recovery after the 2002 change of government with improvements in governance, and attribute the collapse of growth in 2008 partly to post-election violence. Kanyinga (2016) finds that devolution under the 2010 Constitution has begun to weaken the patronage-based allocation of development resources. Econometric studies are more recent. Chepkemboi and Korir (2025) find that government effectiveness and control of corruption had positive and significant effects on growth over 1997 to 2022, while voice and accountability had a negative effect and political stability none.

These studies treat human capital and institutions as separate determinants of growth. None asks whether the return to human capital in Kenya depends on institutional quality, and those that use governance indicators are limited to the period after 1996.

2.5 Hypotheses

The literature reviewed above leads to three hypotheses. Following the factor and endogenous growth views (Lucas, 1988; Mankiw et al., 1992) and Kenyan time-series evidence (Otieno, 2016), the first concerns the direct effect of human capital. Following the institutional literature (Acemoglu et al., 2001; North, 1990) and the Kenyan evidence on governance (Chepkemboi & Korir, 2025), the second concerns the direct effect of institutional quality. Following the allocation, leakage and complementarity mechanisms (Mauro, 1995; Murphy et al., 1991; Reinikka & Svensson, 2004) and the panel evidence on interaction effects (Abu Alfoul et al., 2024; Ali et al., 2018), the third concerns their interaction.

H1: Human capital has a positive long-run effect on real output in Kenya.

H2: Institutional quality has a positive long-run effect on real output in Kenya.

H3: The long-run marginal effect of human capital on real output increases with institutional quality.

H3 is tested against the alternative suggested by Glaeser et al. (2004), Raifu et al. (2021) and Doré and Teixeira (2023), under which the return to human capital does not vary with institutional quality, or under which human capital and institutions act as substitutes so that the interaction is negative.

3. Theoretical framework

3.1 Production technology

The framework extends the augmented Solow model of Mankiw et al. (1992) in two directions suggested by the institutional literature reviewed in Section 2. Aggregate output in year t is produced with physical capital, labour and human capital:

$$Y_t - A_t K_t^{\alpha} L_t^{\gamma} H_t^{\theta_t} \qquad (1)$$

where Y is real output, K the physical capital stock, L employment, H human capital per worker and A total factor productivity (TFP). The elasticities α and γ are constant. The elasticity of output with respect to human capital, θ, carries a time subscript because it is allowed to depend on the institutional environment.

3.2 Two roles for institutions

Institutions enter the model in two ways. The first follows North (1990) and Hall and Jones (1999), for whom institutions determine the efficiency with which all inputs are combined. Institutional quality, G, therefore shifts TFP:

$$A_t = A_0 \exp(\lambda G_t + \varepsilon_t) \qquad (2)$$

where λ measures the direct effect of institutions on output and ε captures other influences on productivity. This is the additive channel assumed in most of the growth literature, including the original version of this study.

The second role is specific to human capital. The allocation, leakage and complementarity mechanisms described in Section 2.3 (Mauro, 1995; Murphy et al., 1991; Reinikka & Svensson, 2004) all imply that a given stock of skills contributes more to output when rent-seeking is less rewarding, when public education spending reaches schools, and when skilled workers have more capital to work with. The output elasticity of human capital is therefore modelled as a linear function of institutional quality:

$$\theta_t = \theta_0 + \theta_1 G_t \qquad (3)$$

The parameter θ_0 is the elasticity that would prevail if G were zero, and θ_1 measures how much that elasticity changes as institutions improve. The institutional view predicts $\theta_1 > 0$. The view of Glaeser et al. (2004), in which human capital is the more basic cause of growth and institutions largely follow from it, is consistent with $\theta_1 = 0$, while a substitution relationship, in which human capital compensates for weak institutions, implies $\theta_1 < 0$.

The specification nests the growth accounting framework used to construct the Penn World Table, in which output is Cobb-Douglas in capital and human-capital-augmented labour with no role for institutions in the return to skills (Feenstra et al., 2015). That framework corresponds to $\theta_1 = 0$ and $\gamma = \theta_0$, so it can be tested as a restriction on the estimated model rather than imposed.

3.3 The estimating equation

Substituting (2) and (3) into (1) and taking natural logarithms gives:

$$\ln Y_t = \ln A_0 + \alpha \ln K_t + \gamma \ln L_t + \theta_0 \ln H_t + \lambda G_t + \theta_1 (G_t \times \ln H_t) + \varepsilon_t \qquad (4)$$

The interaction term arises directly from equation (3); it is a consequence of the theory, not an addition to it. Differentiating (4) gives the elasticity of output with respect to human capital:

$$\frac{\partial \ln Y_t}{\partial \ln H_t} = \theta_0 + \theta_1 G_t \qquad (5)$$

Equation (5) shows why the coefficient on human capital in (4) cannot be read on its own. It is the effect of human capital when G equals zero, which may lie outside the observed range of institutional quality, and the

multicollinearity between ln H and its product with G inflates its standard error (Brambor et al., 2006). Both problems are addressed by centring the two variables on their sample means before forming the product. Let lower-case letters denote logarithms, and let h and g denote the centred values of ln H and G. Equation (4) becomes:

$$y_t = \beta_0 + \beta_1 k_t + \beta_2 l_t + \beta_3 h_t + \beta_4 g_t + \beta_5 (h_t \times g_t) + \varepsilon_t \qquad (6)$$

Centring changes the meaning of the main effects but not the interaction. The coefficient $\beta_3 = \theta_0 + \theta_1 \bar{G}$ is the elasticity of output with respect to human capital at the sample mean of institutional quality, $\beta_4 = \lambda + \theta_1 \cdot \text{mean}(\ln H)$ is the effect of institutional quality at the sample mean of human capital, and $\beta_5 = \theta_1$. The quantity of interest for policy is the conditional elasticity at any observed level of institutional quality, together with its standard error:

$$ME(g) = \beta_3 + \beta_5 g, \qquad \text{Var}[ME(g)] = \text{Var}(\hat{\beta}_3) + g^2 \text{Var}(\hat{\beta}_5) + 2g\text{Cov}(\hat{\beta}_3, \hat{\beta}_5) \qquad (7)$$

Equation (7) is evaluated across the full range of g observed in Kenya between 1963 and 2023, so that the analysis identifies the levels of institutional quality, if any, at which human capital has a statistically significant effect on output.

3.4 From levels to growth

Equation (6) describes a long-run relationship between the level of output and the levels of its determinants. It does not imply that output is always at that level. If the variables are integrated of order one and cointegrated, deviations from (6) are corrected over time, and the growth of output in any year depends on the growth of the inputs and on the size of the previous year's deviation from the long-run path. The ARDL model set out in Section 4 estimates both parts: the long-run coefficients in (6), which answer the hypotheses, and the short-run dynamics with the speed of adjustment, which connect the level relationship to annual growth. This structure reconciles the factor view of human capital, which predicts level effects, with the policy interest in growth: a permanent improvement in institutional quality raises the long-run output path through β_4 and through the larger return to the existing stock of human capital, and output grows faster during the adjustment to that higher path.

3.5 Testable predictions

Table 2: Hypotheses, parameters and expected signs

Hypothesis	Parameter in equation (6)	Expected sign	Theoretical basis
H1: human capital raises long-run output	β_3 (elasticity at mean institutional quality)	Positive	Lucas (1988); Mankiw et al. (1992)
H2: institutional quality raises long-run output	β_4 (effect at mean human capital)	Positive	North (1990); Hall and Jones (1999); Acemoglu et al. (2001)
H3: the return to human capital rises with institutional quality	β_5, and ME(g) increasing and significant over part of the observed range of g	Positive	Murphy et al. (1991); Mauro (1995); Reinikka and Svensson (2004)
Alternative to H3	β_5	Zero or negative	Glaeser et al. (2004)
Controls	β_1 (capital), β_2 (labour)	Positive	Solow model; Mankiw et al. (1992)

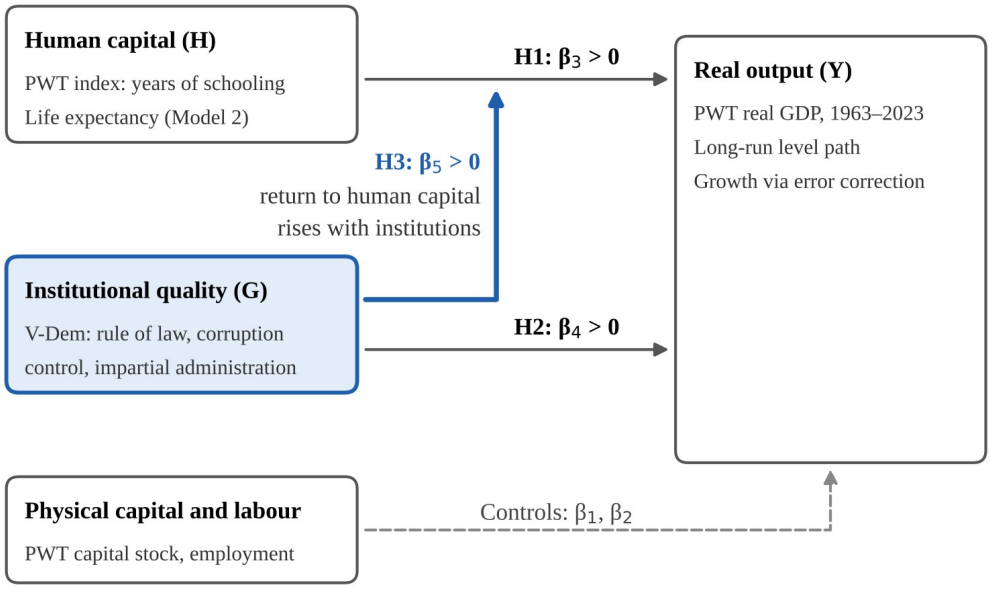


Figure 1: Conceptual framework
Source: Authors' construction based on Section 3.

Human capital and institutional quality each affect long-run output directly (H1 and H2). Institutional quality also acts on the link between human capital and output, raising the return to a given stock of skills (H3). Physical capital and labour enter as controls.

4. Data and econometric method

4.1 Data and variables

The study uses annual data for Kenya from 1963, the year of independence, to 2023, the last year available in Penn World Table (PWT) version 11.0. This gives 61 observations for every variable, with no interpolated or projected values. Output, physical capital, employment and human capital are taken from PWT 11.0 (Feenstra et al., 2015), and institutional quality from version 16 of the Varieties of Democracy (V-Dem) dataset (Coppedge et al., 2026). Table 3 defines the variables.

Table 3: Variables, measures and sources

Variable	Symbol	Measure	Source	Transformation
Real output	Y	Real GDP at constant national prices (PWT code rgdpna)	PWT 11.0	Natural log, y
Physical capital	K	Capital stock at constant national prices (rnna)	PWT 11.0	Natural log, k
Labour	L	Number of persons engaged (emp)	PWT 11.0	Natural log, l
Human capital	H	Human capital index based on years of schooling and returns to education (hc)	PWT 11.0	Natural log, centred, h
Institutional quality	G	First principal component of three V-Dem indices (below)	V-Dem v16	Standardised, centred, g
Rule of law		Rule of law index (v2x_rule), 0 to 1	V-Dem v16	Input to G
Political corruption		Political corruption index (v2x_corr), 0 to 1, reversed so that higher values mean less corruption	V-Dem v16	Input to G
Public administration		Rigorous and impartial public administration (v2clrspct)	V-Dem v16	Input to G
Life expectancy	LE	Life expectancy at birth, years (e_pelifeex)	V-Dem v16	Natural log, centred, Model 2

Three choices in Table 3 respond directly to weaknesses in earlier work. First, the PWT capital stock replaces gross capital formation, because the production function in equation (1) relates output to the stock of capital rather than to the annual flow of investment. Second, the PWT human capital index contains no income component, so the mechanical correlation that arises when an income-based index such as the Human Development Index is regressed against output is avoided. Third, the V-Dem indices are available for Kenya across the whole sample, whereas the Worldwide Governance Indicators begin only in 1996 (Kaufmann et al., 2011).

Institutional quality is constructed in three steps. Each of the three V-Dem indices is standardised to zero mean and unit variance over 1963 to 2023, with the corruption index multiplied by minus one so that all three increase with institutional quality. Principal component analysis is then applied to the three standardised series, and the first principal component is retained as G. The loadings, the share of variance explained by the first component and the correlations among the three inputs are reported with the results. Finally, G is rescaled to unit standard deviation, so that a one-unit change corresponds to a one-standard-deviation change in institutional quality over Kenya's post-independence history. The three indices measure the dimensions emphasised in Section 2.3: constraints on arbitrary action (rule of law), the diversion of public resources (corruption) and the competence and impartiality of the bureaucracy that delivers education spending (public administration).

Descriptive statistics and the correlations among the variables are reported in Section 5.1 (Table 4).

4.2 Unit roots and structural breaks

The ARDL bounds test is valid when the regressors are integrated of order zero, order one or a mixture of the two, but not of order two (Pesaran et al., 2001). The order of integration of each series is therefore established before estimation. Three tests are applied to the level and first difference of every variable: the augmented Dickey-Fuller test (Dickey & Fuller, 1979) and the Phillips-Perron test (Phillips & Perron, 1988), both with a null hypothesis of a unit root, and the KPSS test (Kwiatkowski et al., 1992), whose null hypothesis is stationarity. Using tests with opposite null hypotheses guards against the low power of unit root tests in a sample of 61 observations.

Kenya's post-independence history includes several events that may have shifted the level or trend of output and institutions, among them the economic decline of the 1980s and 1990s, the return to multiparty politics in the early 1990s, the change of government in 2002 and the post-election crisis of 2008 (Kimenyi et al., 2016), as well as the COVID-19 shock of 2020. Conventional unit root tests can mistake a structural break for a unit root. The Zivot and Andrews (1992) test, which allows one endogenously determined break in intercept and trend, is therefore also reported. If it identifies a significant break, the break date is used to define a dummy variable in the ARDL model; break dates are taken from the test, not imposed in advance.

The interaction term h × g is the product of two centred series and has its own time-series properties, which are tested in the same way. Its order of integration does not need to match that of its components, since the bounds test does not require a common order of integration among the regressors.

4.3 The ARDL bounds testing model

The long-run relationship in equation (6) is estimated with the autoregressive distributed lag (ARDL) approach of Pesaran and Shin (1999) and Pesaran et al. (2001). The approach has three advantages in this setting. It does not require all regressors to share the same order of integration; it performs well in small samples compared with system-based cointegration methods; and, with sufficient lags of the dependent variable and the regressors, it yields consistent long-run estimates even when some regressors are endogenous (Pesaran & Shin, 1999). The conditional error correction form of the model is:

$$\Delta y_t = c_0 + \phi\, y_{t-1} + \sum_{j=1}^{5} \pi_j\, x_{j,t-1} + \sum_{i=1}^{p-1} \gamma_i\, \Delta y_{t-i} + \sum_{j=1}^{5} \sum_{i=0}^{q_j-1} \delta_{ji}\, \Delta x_{j,t-i} + \boldsymbol{\psi}'\mathbf{D}_t + u_t \qquad (8)$$

where x_1 to x_5 are k, l, h, g and h × g, D is the vector of break dummies identified in Section 4.2, and u is a white-noise error. The long-run coefficients of equation (6) are recovered as $\beta_j = -\pi_j / \varphi$. The lag orders p and q_j are selected by the Akaike information criterion over all combinations of one or two lags, a limit set by the sample size, with every model compared on a common estimation sample. With five regressors and two lags, the most general model estimates about 20 parameters from 59 usable observations.

Cointegration is tested in two stages. The F-test of the joint null hypothesis $\varphi = \pi_1 = \ldots = \pi_5 = 0$ is compared with lower and upper critical bounds; a statistic above the upper bound indicates a long-run relationship regardless of whether the regressors are I(0) or I(1). The t-test of $\varphi = 0$ is then used to rule out the degenerate case in which the lagged regressors are significant but the lagged dependent variable is not. Because the asymptotic critical values of Pesaran et al. (2001) are too permissive in samples of this size (Narayan, 2005), finite-sample critical values and p-values for both the F-test and the t-test are simulated for the actual sample size and lag structure, with 10,000 replications for the F-test, following the simulation logic behind the small-sample tables of Narayan (2005) and the response surfaces of Kripfganz and Schneider (2020). Asymptotic p-values are reported for comparison. The model includes an unrestricted intercept and no trend (Pesaran et al., 2001, Case III), with a trend added as a robustness check.

Two versions of the model are estimated. Model 1 is equation (6), in which human capital is measured by schooling alone. Model 2 adds health, the second component of human capital discussed in Section 2.1 (Bloom et al., 2004), by including the centred log of life expectancy and its interaction with g. If cointegration is established, each model is re-estimated in error correction form:

$$\Delta y_t = c_0 + \sum_{i=1}^{p-1} \gamma_i\, \Delta y_{t-i} + \sum_{j=1}^{5} \sum_{i=0}^{q_j-1} \delta_{ji}\, \Delta x_{j,t-i} + \boldsymbol{\psi}'\mathbf{D}_t + \phi\, ECT_{t-1} + u_t \qquad (9)$$

where ECT is the deviation of output from the long-run relationship in equation (6). A negative and significant φ confirms adjustment towards the long-run path, and its absolute value is the share of a deviation corrected within one year.

4.4 Conditional marginal effects

Hypothesis H3 is evaluated from the long-run coefficients by computing the conditional elasticity of output with respect to human capital in equation (7) at every level of institutional quality observed between 1963 and 2023. Because the long-run coefficients are ratios of the short-run parameters in equation (8), their variances and covariance are obtained by the delta method. The conditional elasticity and its 95% confidence interval are plotted against g, with a rug plot of the observed values of g, following Brambor et al. (2006). The range of g over which the elasticity is significantly different from zero is reported, together with the corresponding years. Moderation is supported if β_5 is positive and the conditional elasticity is significant over at least part of the observed range of institutional quality.

4.5 Diagnostic tests

The estimated model is checked for serial correlation with the Breusch-Godfrey LM test at one and two lags, for heteroskedasticity with the Breusch-Pagan-Godfrey test, for non-normality of residuals with the Jarque-Bera test and for functional form with the Ramsey RESET test. Parameter stability is assessed with the CUSUM and

CUSUM of squares tests based on recursive residuals (Brown et al., 1975), and both plots are reported. Variance inflation factors for the long-run regressors are reported before and after centring.

4.6 Robustness

Five sets of checks test whether the conclusions depend on the estimator or on measurement choices. First, the long-run coefficients are re-estimated with fully modified OLS (Phillips & Hansen, 1990), dynamic OLS (Stock & Watson, 1993) and canonical cointegrating regression (Park, 1992), which correct for endogeneity and serial correlation in different ways. Second, G is replaced by the simple average of the three standardised V-Dem indices and by each index separately, to show whether any one dimension of institutions drives the result. Third, a linear trend is added to the model (Pesaran et al., 2001, Case V). Fourth, heteroskedasticity-robust (HC1) standard errors are computed for the long-run coefficients. Fifth, the direction of causality is examined with the Toda and Yamamoto (1995) procedure, which is valid regardless of the order of integration of the series, to address the possibility raised by Glaeser et al. (2004) and Bils and Klenow (2000) that output growth drives both human capital and institutions.

4.7 Software and data availability

Estimation is carried out in Python 3.13 with statsmodels 0.15 (ARDL, error correction, bounds tests and diagnostics) and arch 8.0 (unit root tests and the FMOLS, DOLS and CCR estimators). The Penn World Table file was verified against the SHA-1 checksum published on DataverseNL. The dataset and code will be deposited in a public repository on acceptance. All data sources are publicly available.

5. Results

5.1 Descriptive statistics and the institutional index

Table 4 summarises the data. Real GDP rose from US$27.7 billion in 1963 to US$315.2 billion in 2023 at constant national prices, an increase of 2.43 log points, while the PWT human capital index rose from 1.22 to 2.61. Output, capital, employment and schooling all reach their minimum in 1963 and their maximum in 2023, and their logarithms are almost perfectly correlated (pairwise correlations between 0.96 and 1.00). Institutional quality and life expectancy follow different paths. Institutional quality is only moderately correlated with output (0.70), and life expectancy even less so (0.47), because neither rose steadily.

Table 4: Descriptive statistics, Kenya, 1963 to 2023 (n = 61)

Variable	Mean	Std. dev.	Minimum	Maximum	Year of minimum	Year of maximum
Real GDP (US$ million, constant national prices)	115,523	77,594	27,696	315,169	1963	2023
Capital stock (US$ million, constant national prices)	243,859	174,429	56,394	721,478	1963	2023
Persons engaged (million)	11.26	5.49	3.86	19.95	1963	2023
Human capital index (PWT)	1.79	0.40	1.22	2.61	1963	2023
Institutional quality index, G	0.00	1.00	−1.17	1.97	1984 to 1989	2021
Life expectancy at birth (years)	60.30	3.56	54.83	66.93	2000	2023

Sources: Penn World Table 11.0; V-Dem v16. G is standardised to mean zero and unit standard deviation over 1963 to 2023.

The three V-Dem inputs to the institutional index are strongly correlated: 0.95 between rule of law and impartial public administration, 0.79 between rule of law and (reversed) corruption, and 0.71 between corruption and public administration. The first principal component explains 87.9% of their joint variance, with nearly equal loadings of 0.60 on rule of law, 0.55 on corruption control and 0.59 on public administration, so the index is not dominated by any one dimension. Figure 2 plots it. Institutional quality declined from independence to its lowest level in 1984 to 1989, when it was 1.17 standard deviations below its 1963 to 2023 mean, recovered slowly after the return to multiparty politics, rose sharply after the 2002 change of government and again after the 2010 Constitution, and peaked in 2021. Life expectancy shows a different pattern: it rose to 63.4 years in 1987, fell to 54.8 years in 2000 during the HIV/AIDS epidemic, and recovered to 66.9 years by 2023.

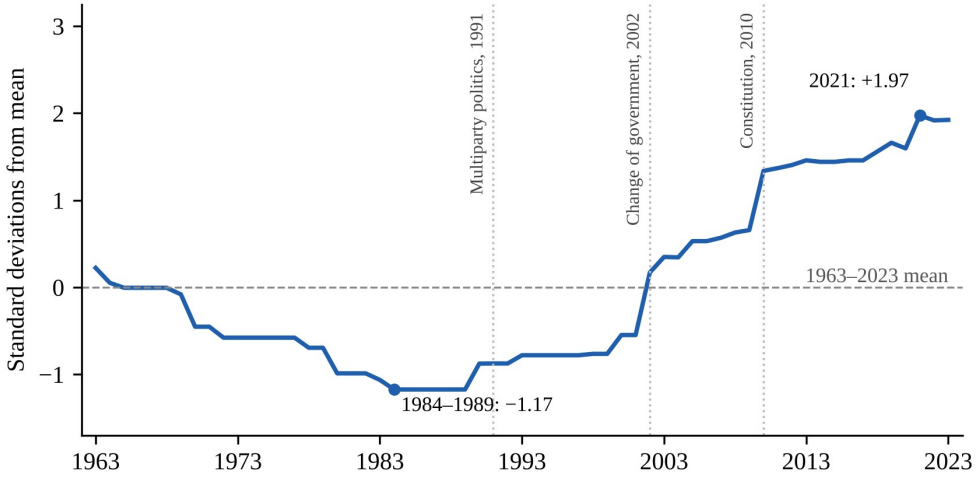


Figure 2: Institutional quality in Kenya, 1963 to 2023

5.2 Unit roots and structural breaks

Table 5 reports the unit root tests. For output, schooling, institutional quality and the health interaction, the ADF and PP tests fail to reject a unit root in levels and reject it decisively in first differences, which indicates integration of order one. The evidence for capital and labour is weaker. The ADF test does not reject a unit root in their first differences ($p = 0.104$ and 0.253), but the PP test does ($p = 0.038$ and 0.019). Both PWT series are smooth by construction, which lowers the power of the ADF test. For life expectancy, the ADF test rejects a unit root in levels against a trend-stationary alternative ($p = 0.013$), the PP test does not, and neither rejects in first differences, reflecting the fall and recovery of the HIV/AIDS period. Second differences of every series are clearly stationary (all ADF p-values below 0.03), so no variable is integrated of order two, which is the condition the bounds test requires.

Table 5: Unit root and structural break tests

Variable	ADF level	ADF first difference	PP level	PP first difference	KPSS level	KPSS first difference	Zivot-Andrews (break year)	Order
y	−1.71 (0.747)	−7.07 (0.000)	−2.05 (0.576)	−7.05 (0.000)	0.133 (0.071)	0.112 (0.529)	−3.38 (1992)	I(1)
k	−3.05 (0.118)	−2.55 (0.104)	−1.44 (0.850)	−2.96 (0.038)	0.162 (0.036)	0.202 (0.265)	−4.05 (2012)	I(1)
l	−1.90 (0.653)	−2.08 (0.253)	0.13 (0.995)	−3.22 (0.019)	0.243 (0.006)	0.539 (0.032)	−4.32 (1994)	I(1)
h	−2.76 (0.212)	−6.17 (0.000)	−2.17 (0.506)	−4.04 (0.001)	0.164 (0.034)	0.183 (0.303)	−3.44 (2004)	I(1)
g	−1.78 (0.717)	−4.21 (0.001)	−1.80 (0.706)	−7.65 (0.000)	0.322 (0.001)	0.749 (0.010)	−3.56 (1980)	I(1)
h × g	1.05 (1.000)	−2.20 (0.205)	0.33 (0.996)	−7.06 (0.000)	0.268 (0.003)	0.521 (0.035)	−2.93 (2003)	I(1)
le	−3.87 (0.013)	−2.06 (0.259)	−1.63 (0.779)	−1.95 (0.307)	0.142 (0.058)	0.185 (0.298)	−3.59 (1989)	I(0) or I(1)
le × g	−1.20 (0.911)	−3.99 (0.001)	−0.70 (0.973)	−4.24 (0.001)	0.191 (0.018)	0.283 (0.152)	−3.56 (2002)	I(1)

Notes: Test statistics with p-values in parentheses. Levels include an intercept and trend; first differences include an intercept. ADF lags selected by AIC (maximum 4). KPSS has stationarity as its null hypothesis. The Zivot-Andrews test allows one break in intercept and trend; its 5% critical value is −5.08. le is the centred log of life expectancy.

The Zivot-Andrews statistics range from −2.93 to −4.32, all well above the 5% critical value of −5.08. The unit root findings are therefore not an artefact of an unmodelled structural break. The most likely break years the test identifies are historically plausible (1980 for institutions, 1992 for output, 2003 and 2004 for schooling and its interaction, 1989 for life expectancy), but none is statistically significant, so no break dummies are added to the ARDL models.

5.3 Cointegration

Table 6 reports the bounds tests. For Model 1, in which human capital is measured by schooling alone, the F-statistic of 3.20 lies between the finite-sample lower and upper bounds at the 5% level (2.81 and 4.18), and the t-statistic on the lagged level of output (−3.65) lies between the simulated t-bounds (−2.86 and −4.13). Both tests are therefore inconclusive: they reject the absence of a long-run relationship if all regressors are I(0), but not if they are I(1), which is the case the unit root tests support. Model 1 does not provide reliable evidence of cointegration.

Adding health changes this result. In Model 2 the F-statistic of 4.44 exceeds the finite-sample upper bound at the 5% level (3.85), with a finite-sample p-value of 0.021 against I(1) regressors, and the t-statistic of −5.51 exceeds the simulated upper t-bound at the 1% level (−5.13). Output, physical capital, labour, schooling, health, institutional quality and the two interaction terms share a stable long-run relationship once both components of human capital are included. Model 2 is therefore the preferred specification, and Model 1 is reported for comparison. Because Model 2 was estimated after Model 1's bounds test proved inconclusive, its results are interpreted alongside the robustness checks in Section 5.6 rather than on their own.

Table 6: ARDL bounds tests for cointegration

	Model 1: schooling	**Model 2: schooling and health**
Selected ARDL order	(2, 1, 1, 2, 1, 1)	(2, 2, 1, 2, 1, 1, 2, 1)
Number of regressors, k	5	7
Observations	59	59
F-statistic	3.20	4.44
Finite-sample 5% bounds, I(0) / I(1)	2.81 / 4.18	2.57 / 3.85
Finite-sample 1% bounds, I(0) / I(1)	3.94 / 5.40	3.52 / 4.89
Finite-sample p-value, I(0) / I(1)	0.027 / 0.157	0.002 / 0.021
Asymptotic p-value, I(0) / I(1)	0.009 / 0.106	0.000 / 0.004
t-statistic on y(t−1)	−3.65	−5.51
Simulated 5% t-bounds, I(0) / I(1)	−2.86 / −4.13	−2.78 / −4.45
Conclusion	Inconclusive	Cointegrated

Notes: Case III (unrestricted intercept, no trend). ARDL order lists the lag of output followed by k, l, h, g, h × g and, in Model 2, le and le × g. Finite-sample F critical values and p-values are simulated for the actual sample and lag structure (10,000 replications); t-bounds are simulated with 3,000 (Model 1) and 1,500 (Model 2) replications.

5.4 Long-run estimates

Table 7 reports the long-run coefficients, which correspond to equation (6) with health added in Model 2. Because human capital and institutional quality are centred, the coefficient on h is the long-run elasticity of output with respect to schooling at Kenya's average level of institutional quality, and the coefficient on g is the effect of institutional quality at average schooling.

Table 7: Long-run coefficients (dependent variable: log real GDP)

Variable	Model 1: coefficient (SE)	Model 1: p-value	Model 2: coefficient (SE)	Model 2: p-value
Physical capital, k	0.101 (0.227)	0.657	0.178 (0.146)	0.231
Labour, l	−0.536 (0.252)	0.039	−0.656 (0.271)	0.020
Schooling, h (β_3)	4.003 (0.841)	< 0.001	4.093 (0.685)	< 0.001
Institutional quality, g (β_4)	−0.041 (0.038)	0.282	0.003 (0.024)	0.910
h × g (β_5)	0.280 (0.318)	0.385	−0.365 (0.190)	0.062
Health, le			1.147 (0.340)	0.002
le × g			0.122 (0.212)	0.567
Constant	11.504 (2.663)	< 0.001	10.834 (1.612)	< 0.001

Notes: Standard errors by the delta method. With heteroskedasticity-robust (HC1) standard errors, the t-statistics in Model 2 are 5.96 for h, 4.43 for le and −1.89 for h × g.

Three findings stand out. First, schooling has a large, positive and highly significant long-run association with output in both models (4.00 and 4.09), which is consistent with H1. The size of the coefficient should not be read as a structural elasticity. The PWT human capital index and real GDP both rise almost monotonically over the sample (correlation 0.99), so with no trend term the schooling coefficient also absorbs the trend in total factor productivity, and the variance inflation factor for h is about 208 even after centring. Second, health has an independent positive effect: a 1% rise in life expectancy is associated with a 1.15% rise in long-run output ($p = 0.002$), and including it is what produces cointegration. Third, institutional quality has no significant direct effect on output in either model ($p = 0.282$ and 0.910), so H2 is not supported.

The interaction between schooling and institutional quality is not significant in Model 1 (0.280, $p = 0.385$). In Model 2 it is negative and significant only at the 10% level (−0.365, $p = 0.062$). Neither model supports H3, which predicts a positive coefficient. The negative coefficient in Model 2 points, if anything, towards the alternative in which the return to schooling is lower, not higher, when institutions are stronger. The interaction between health and institutions is positive but not significant (0.122, $p = 0.567$). The negative long-run coefficient on labour is implausible as an output elasticity and reflects the near-collinearity of employment with schooling (correlation 0.99) rather than a negative contribution of workers to output.

5.5 Short-run dynamics and diagnostics

Table 8 reports the error correction form of Model 2. The error correction term is −0.764 ($p < 0.001$): about 76% of any deviation of output from its long-run path is corrected within one year, so the long-run relationship reasserts itself quickly. In the short run, output growth rises with the current growth of schooling (3.66, $p < 0.001$) and of life expectancy (2.99, $p = 0.024$), and falls with lagged schooling growth (−4.98, $p = 0.011$), a pattern consistent with adjustment costs as cohorts with more schooling enter the labour force. The short-run interaction between schooling growth and institutional quality is negative (−0.291, $p = 0.035$), matching the sign of the long-run interaction. The error correction term in Model 1 is −0.437 ($p = 0.001$).

Table 8: Error correction model for Model 2 (dependent variable: Δy)

Regressor	Coefficient (SE)	p-value
Error correction term, y(t−1)	−0.764 (0.139)	< 0.001
Δy(t−1)	0.286 (0.128)	0.031
Δk	0.430 (0.427)	0.320
Δk(t−1)	0.487 (0.413)	0.246
Δl	−0.210 (0.443)	0.638
Δh	3.660 (0.952)	< 0.001
Δh(t−1)	−4.979 (1.864)	0.011
Δg	0.031 (0.026)	0.238
Δ(h × g)	−0.291 (0.133)	0.035
Δle	2.992 (1.278)	0.024
Δle(t−1)	−1.333 (1.162)	0.258
Δ(le × g)	0.529 (0.301)	0.087
Constant	8.278 (1.970)	< 0.001
R^2 / adjusted R^2	0.641 / 0.467	

Table 9 reports the diagnostic tests. Neither model shows serial correlation at one or two lags, which is the condition on which ARDL inference depends. Residuals are consistent with normality, and the RESET test finds no evidence of misspecified functional form. The Breusch-Pagan-Godfrey test indicates mild heteroskedasticity in both models (p = 0.055 and 0.037 for the LM version); the heteroskedasticity-robust t-statistics reported under Table 7 leave the conclusions unchanged. The CUSUM statistics stay inside their 5% bands throughout, and the maximum CUSUM of squares deviations (0.132 and 0.131) are well below their simulated 5% critical values (0.259 and 0.280), so the parameters are stable over the sample.

Table 9: Diagnostic tests

Test	Model 1: statistic (p)	Model 2: statistic (p)
Breusch-Godfrey LM, 1 lag	2.669 (0.102)	2.050 (0.152)
Breusch-Godfrey LM, 2 lags	3.071 (0.215)	2.622 (0.270)
Breusch-Pagan-Godfrey LM	22.040 (0.055)	31.320 (0.037)
Jarque-Bera	0.885 (0.642)	5.522 (0.063)
Ramsey RESET, F	0.038 (0.846)	0.456 (0.503)
CUSUM	Within 5% bands	Within 5% bands
CUSUM of squares, maximum deviation (5% critical value)	0.132 (0.259)	0.131 (0.280)

Centring reduces collinearity where it was introduced by the interaction: the variance inflation factor falls from 1,325 to 18 for the interaction term and from 195 to 6 for institutional quality. The high values for capital, labour and schooling (149 to 208) remain, because those series trend together and centring cannot separate them.

5.6 Conditional marginal effects

Figure 3 plots the long-run elasticity of output with respect to schooling across the full range of institutional quality observed in Kenya, following equation (7). The elasticity is 4.52 (95% confidence interval 3.05 to 5.99) at the weakest institutions, recorded in 1984 to 1989, and 3.37 (1.81 to 4.93) at the strongest, recorded in 2021. It is positive and significant at every level of institutional quality observed between 1963 and 2023. The decline across the range is the interaction coefficient of Table 7, which is significant only at the 10% level, and the two confidence intervals overlap almost entirely. The marginal-effects evidence therefore does not show that stronger institutions raise the return to schooling, as H3 predicts. If anything, the return is slightly lower when institutions are stronger, but this difference is not estimated precisely.

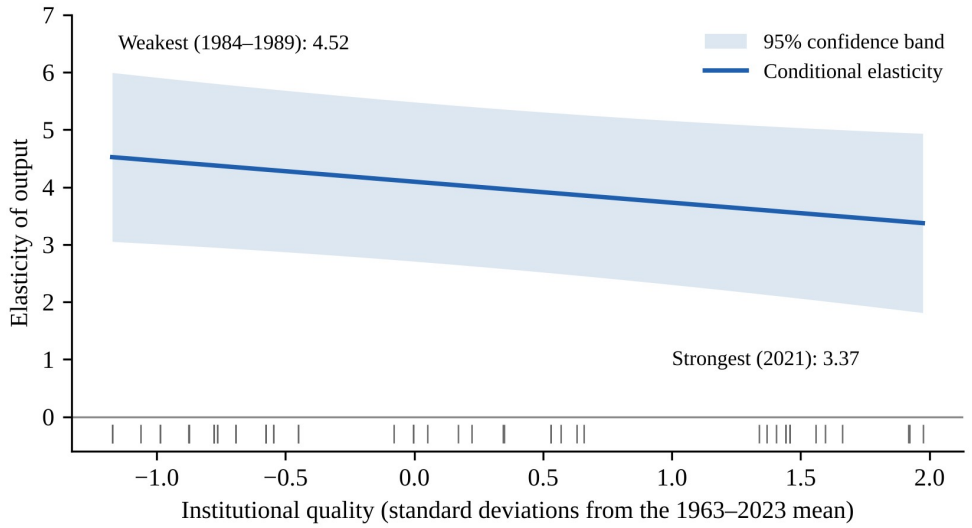


Figure 3: Conditional elasticity of output with respect to schooling

The elasticity of output with respect to life expectancy moves in the opposite direction. It rises from 1.00 (0.54 to 1.47) at the weakest institutions to 1.39 (−0.04 to 2.82) at the strongest and is significant in 58 of the 61 years, losing significance only at the highest levels of institutional quality, where few observations lie and the confidence band widens.

5.7 Robustness and causality

Table 10 reports the main long-run coefficients of Model 2 under alternative estimators, deterministic terms and measures of institutional quality. The bounds test confirms cointegration in every ARDL variant (asymptotic p-values against I(1) regressors between 0.002 and 0.006).

Table 10: Robustness of the Model 2 long-run coefficients: coefficient (t-statistic)

Specification	Schooling, h	h × g	Health, le	Institutional quality, g
Model 2, ARDL (Table 7)	4.093 (5.97)	−0.365 (−1.92)	1.147 (3.37)	0.003 (0.11)
HC1 robust standard errors	4.093 (5.96)	−0.365 (−1.89)	1.147 (4.43)	0.003 (0.12)
Linear trend added (Case V)	5.559 (4.24)	−0.393 (−1.98)	1.175 (3.34)	0.014 (0.55)
Fully modified OLS	2.113 (12.73)	−0.190 (−3.20)	1.748 (13.63)	0.012 (1.62)
Canonical cointegrating regression	2.098 (10.61)	−0.199 (−3.35)	1.737 (11.73)	0.014 (1.98)
Dynamic OLS (one lead, one lag)	3.442 (8.50)	−0.404 (−1.56)	1.095 (2.72)	0.027 (0.84)
G = simple average of three indices	4.073 (5.94)	−0.368 (−1.97)	1.149 (3.35)	0.002 (0.10)
G = rule of law only	4.027 (5.72)	−0.323 (−1.42)	1.148 (3.46)	0.009 (0.24)
G = corruption control only	2.931 (3.64)	−0.365 (−2.76)	0.903 (1.45)	−0.007 (−0.41)
G = impartial public administration only	4.194 (5.31)	−0.079 (−0.35)	1.039 (3.46)	−0.025 (−0.80)
Model 1, schooling only (Table 7)	4.003 (4.76)	0.280 (0.88)		−0.041 (−1.09)

Three results are robust. First, schooling has a positive and significant long-run association with output in every specification, although its size varies from 2.10 to 5.56 and depends on whether a trend is included, which

confirms that the coefficient also captures trend productivity. Second, health has a positive long-run coefficient in all ten Model 2 variants, significant in nine, the exception being the specification that measures institutions by corruption control alone. Third, institutional quality has no significant direct effect at the 5% level in any specification except canonical cointegrating regression, where the coefficient is small (0.014) and only just significant (t = 1.98).

The interaction between schooling and institutional quality is never positive in any specification that includes health. It is negative in all ten Model 2 variants, significant at the 5% level in three (fully modified OLS, canonical cointegrating regression and the corruption-only index), significant at the 10% level in four more, and insignificant in the remaining three. The corruption dimension of institutions carries the negative interaction, while the public administration dimension shows almost none (−0.079, t = −0.35). Under fully modified OLS and canonical cointegrating regression the interaction between health and institutional quality is positive and significant (0.605, t = 6.70, and 0.599, t = 5.20), but it is insignificant under ARDL and dynamic OLS, so this result is fragile.

The Toda and Yamamoto (1995) tests, based on a vector autoregression in output, schooling and institutional quality with three lags selected by the Akaike criterion and one additional lag, find no Granger causality in any direction (all p-values between 0.43 and 0.94). This rules out strong short-horizon feedback from output to schooling or institutions, the concern raised by Bils and Klenow (2000) and Glaeser et al. (2004), but with 61 annual observations the test has low power, and the absence of predictive causality is not evidence against a long-run relationship.

6. Discussion

The results support H1 but not H2 or H3. Human capital, measured by both schooling and health, has a positive long-run relationship with Kenya's output. Institutional quality has no direct long-run effect. The interaction between schooling and institutional quality is never positive once health is included, and in several specifications it is significantly negative. This section relates each finding to the literature reviewed in Section 2 and considers what explains it.

6.1 Human capital and output

The positive long-run association between schooling and output is consistent with the factor view of Mankiw et al. (1992), with the endogenous growth view of Lucas (1988), and with earlier Kenyan time-series evidence that education per worker raised output between 1967 and 2010 (Otieno, 2016). The size of the coefficient, however, cannot be read as the return to an additional year of schooling. It ranges from 2.10 to 5.56 across estimators, rises when a trend is added, and the human capital index is almost perfectly correlated with output and capital. The coefficient therefore combines the contribution of schooling with part of the trend in productivity. This is the time-series counterpart of the measurement problem that Krueger and Lindahl (2001) identify in cross-country data, and the safe conclusion concerns the sign and significance of the relationship rather than its magnitude.

The health result is new for Kenya and changes how the schooling evidence should be read. Model 1, with schooling alone, fails to establish cointegration, whereas Model 2, with life expectancy added, does so at the 5% level, and life expectancy has a long-run elasticity of about 1.1 that survives nine of ten robustness checks. This matches the production function evidence of Bloom et al. (2004), who find a positive effect of life expectancy on output that is independent of schooling. In Kenya the health channel has a clear historical counterpart. Life expectancy fell from 63.4 years in 1987 to 54.8 years in 2000 during the HIV/AIDS epidemic, and the 1990s were the decade of slowest growth since independence, at 2.2% a year (Kimenyi et al., 2016). A model that omits health attributes part of that slowdown to the wrong variables, which suggests that Kenyan growth studies relying on schooling or education spending alone may be misspecified.

6.2 Institutional quality and output

The absence of a direct long-run effect of institutional quality contrasts with the cross-country literature, in which institutions are a fundamental determinant of income (Acemoglu et al., 2001; Hall & Jones, 1999; Rodrik et al., 2004), and with recent Kenyan time-series evidence, which finds positive effects of government effectiveness and control of corruption over 1997 to 2022 (Chepkemboi & Korir, 2025). Three differences in design help to reconcile these findings. First, the sample periods differ. The governance indicators used in earlier Kenyan work begin in 1996, so they cover only the period after 2002 in which institutions and growth improved together. Over 1963 to 2023 the relationship is less regular: institutional quality declined steadily through the 1960s and 1970s while growth averaged 5.7% and 7.2% (Kimenyi et al., 2016), and only later did the two move together. Second, cross-country studies identify the effect of institutions from persistent differences in levels

between countries, built up over generations, whereas a single-country time series identifies it from changes within one country over a few decades. Third, the model controls for schooling and health. If better institutions raise output mainly by improving the accumulation of human capital, as Dias and Tebaldi (2012) and Adams-Kane and Lim (2016) argue, then much of their effect is already captured by h and le, and the remaining direct effect would be small. The results are therefore consistent with institutions mattering for Kenya's output, but not as a separate input over and above human capital.

6.3 Why institutions do not raise the return to schooling

The finding that stronger institutions do not raise the return to schooling places Kenya alongside the single-country studies of Nigeria (Raifu et al., 2021) and Brazil (Doré & Teixeira, 2023) and against the panel evidence for Sub-Saharan Africa (Abu Alfoul et al., 2024) and for a large cross-section of countries (Ali et al., 2018). This is the pattern anticipated in Section 2.3: a moderating effect estimated as an average across many countries need not hold within any one of them. The Kenyan evidence indicates that panel-based conclusions should not be transferred to country-level policy without country-specific evidence.

Three explanations for the negative sign of the interaction are compatible with the data, and the present design cannot distinguish between them. The first is substitution. When institutions are weak, education may be particularly valuable because educated workers and managers are better able to deal with arbitrary rules and corrupt officials, so that the return to schooling is higher than it would be under impartial institutions. That the negative interaction is driven by the corruption index rather than the public administration index is consistent with this reading, which is close to the argument of Glaeser et al. (2004) that human capital is the more basic factor. The second explanation is timing. The largest improvement in institutional quality, after 2002, coincided with the fastest expansion of schooling, following the abolition of primary school fees in 2003 (Lucas & Mbiti, 2012). If the marginal return to schooling falls as the stock of schooling grows, the interaction term will pick up that decline, because in a single time series a rising institutional index cannot be separated from the passage of time. The third explanation is measurement. The PWT human capital index is built from years of schooling and assumed rates of return to education (Feenstra et al., 2015), so it does not record changes in the quality of schooling, which Hanushek and Woessmann (2008) show matters more for growth than years of attendance.

These results bear on the mechanisms set out in Section 2.3. Leakage of education spending (Reinikka & Svensson, 2004) and the diversion of talent into rent-seeking (Murphy et al., 1991) imply that weak institutions reduce the amount of human capital produced from a given budget, or the use made of it, but they do not require that the return to each unit of measured schooling vary with institutions. If these mechanisms work mainly through the accumulation of human capital, then institutions act as a mediator rather than a moderator, and a moderation test of the kind used here, which conditions on the stock of human capital, would not detect them. The tentative positive interaction between health and institutions under fully modified OLS and canonical cointegrating regression fits this reading, since institutional quality is more plausibly linked to the delivery of health services than to the productivity of a given year of schooling, but that result is not robust across estimators.

The analysis also has a methodological implication. The earlier version of this study inferred a positive moderating effect from an interaction coefficient that was not statistically significant, supported by a joint test of three coefficients. Reporting the conditional marginal effect across the observed range of institutional quality, as Brambor et al. (2006) recommend, shows directly that the return to schooling does not rise with institutional quality in Kenya, and prevents that kind of over-reading.

7. Conclusion

7.1 Conclusions

This study asked whether institutional quality conditions the long-run return to human capital in Kenya. Using annual data for 1963 to 2023 from Penn World Table 11.0 and V-Dem v16, an ARDL bounds framework and conditional marginal effects, it reaches three conclusions. First, human capital is positively related to Kenya's long-run output, and both of its components matter. Schooling is positive and significant in every specification, and health has an independent effect: a 1% rise in life expectancy is associated with a rise of about 1.1% in long-run output. A stable long-run relationship appears only when health is included. Second, institutional quality has no significant direct long-run effect on output once schooling and health are controlled for. Third, stronger institutions do not raise the return to schooling. The conditional elasticity of output with respect to schooling is positive at every level of institutional quality observed since independence, falling slightly from 4.52 at the weakest institutions to 3.37 at the strongest, and the interaction between schooling and institutions is negative in every specification that includes health, though significant at the 5% level in only three of ten.

These results place Kenya with the single-country evidence from Nigeria and Brazil rather than with the panel evidence for Sub-Saharan Africa. A moderating effect of institutions estimated on average across many countries does not hold within Kenya's own history.

7.2 Policy implications

Four implications follow, each tied to a specific result.

1. Health spending is growth policy as well as social policy. The long-run elasticity of output with respect to life expectancy, and the coincidence of the HIV/AIDS decline in life expectancy with the slowest growth decade since independence, indicate that sustained investment in population health contributes to output directly. This supports the inclusion of healthcare among the five priority sectors of the Fourth Medium Term Plan (Republic of Kenya, 2023) and argues for evaluating health programmes partly by their contribution to long-run output.
2. Investment in schooling remains justified, but its measured contribution rests on years of schooling, not on what pupils learn. Because the human capital index used here does not reflect quality, and because cognitive skills matter more for growth than attendance (Hanushek & Woessmann, 2008), education policy should track learning outcomes alongside enrolment and completion as access continues to expand (Lucas & Mbiti, 2012).
3. Governance reforms should not be justified on the expectation that they will raise the return to existing schooling, since the Kenyan evidence does not support that effect. Their case rests on other grounds that this study does not reject: reducing leakage of public spending before it reaches schools and clinics (Reinikka & Svensson, 2004) and improving the accumulation of human capital (Adams-Kane & Lim, 2016; Dias & Tebaldi, 2012). Reforms should therefore be evaluated by their effect on service delivery and on the quantity and quality of education and health produced, not by an assumed multiplier on the return to schooling.
4. Panel-based findings should be applied to Kenyan policy with caution. Where cross-country averages and country-specific evidence disagree, as they do here, policy design should rest on the country-specific evidence or on studies that allow the effect to differ across countries.

7.3 Limitations and future research

The study has five main limitations. First, it relies on a single country and 61 annual observations, which limits statistical power, especially for the interaction terms and the causality tests. Second, output, capital, employment and schooling trend together, with variance inflation factors between 149 and 208 even after centring, so the individual long-run coefficients are imprecise and the magnitude of the schooling coefficient partly reflects trend productivity; the sign and significance of the coefficients are more reliable than their size. Third, the preferred model, which adds health, was estimated after the schooling-only model's bounds test proved inconclusive. Although health was motivated by the literature in Section 2.1 and the result holds across ten robustness checks, this sequence carries some risk of specification search. Fourth, each measure has known weaknesses: the PWT human capital index is based on years of schooling with assumed returns, the life expectancy series is smoothed, the V-Dem indices are expert-coded, and the evidence on whether capital and employment are integrated of order one rather than two is mixed. Fifth, the ARDL approach reduces but does not eliminate concerns about endogeneity between output, human capital and institutions.

Three extensions follow. A mediation design, such as a system that models institutional quality as a determinant of schooling and health accumulation, would test the channel that this study suggests but cannot estimate. County-level data available since devolution in 2013 would allow the interaction to be estimated from variation across Kenya's 47 counties rather than over time alone. Quality-adjusted measures of human capital, such as national examination results or harmonised learning outcomes, would show whether the return to what pupils learn, rather than to years enrolled, depends on institutions.

Declarations

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Data availability. All data are publicly available. Penn World Table 11.0 is available from the Groningen Growth and Development Centre and DataverseNL (https://doi.org/10.34894/FABVLR); the file used was verified against the published SHA-1 checksum. V-Dem v16 is available from the V-Dem Institute (https://doi.org/10.23696/vdemds26). The analysis dataset and the code that reproduces every table and figure are available from the corresponding author on request and will be deposited in a public repository on acceptance.

Ethics statement. Not applicable. The study uses only publicly available country-level secondary data and involves no human participants.

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