

Asymmetric Long-run Effects of Tax Revenue on Economic Growth in Indonesia

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Abstract

Indonesia combines one of the lowest tax-to-GDP ratios among major emerging economies with a growth record punctuated by three systemic crises, yet the empirical literature on its tax and growth nexus assumes that revenue increases and decreases carry symmetric effects. This study examines whether the long-run relationship between tax revenue and real gross domestic product (GDP) is asymmetric, using a nonlinear autoregressive distributed lag model on annual data for Indonesia from 1994 to 2024, with domestic credit and the Corruption Perceptions Index as long-run regressors and capital expenditure in a parallel specification. A symmetric benchmark detected no cointegration and failed the functional-form test, whereas the asymmetric model established a level relationship. A 1 percent cumulative increase in tax revenue was associated with a 0.71 percent rise in long-run GDP, cumulative decreases showed no significant association and long-run symmetry was rejected (Wald $F(1, 22) = 7.84, p = 0.010$). A one-point improvement in the corruption index, measured on a 0 to 100 scale, not in percentage units, was associated with approximately 1.5 percent higher long-run output, while the capital-expenditure specification produced no long-run relationship, locating the asymmetry on the revenue side of the budget. The asymmetry persisted when revenue was measured in real terms and as a share of GDP, under alternative deterministic specifications and after a level shift in 1999 was allowed. These findings characterize revenue mobilization as growth policy rather than as consolidation alone, subject to the associational limits of a small annual sample.

Keywords: Bounds Testing, Dynamic Simulations, Fiscal Policy, Institutional Quality, Nonlinear Autoregressive Distributed Lag, Tax Revenue.

Introduction

Indonesia presents a fiscal-capacity puzzle. A trillion-dollar G20 economy collected tax revenue equal to 11.8 percent of gross domestic product (GDP) in 2024, a ratio that has hovered near 12 percent for almost two decades and that remains well below the Asia-Pacific average of 19.5 percent (1). Over the same period three systemic shocks struck the economy: the Asian financial crisis, in which output contracted by more than 13 percent in 1998 (2, 3); the global financial crisis of 2008 to 2009; and the COVID-19 recession of 2020 to 2021. Revenue collection, credit supply and investment were disrupted simultaneously in each episode (4, 5) and the recoveries need not have mirrored the collapses. Whether growth responds symmetrically to fiscal expansion and contraction therefore determines whether revenue mobilization should be treated as growth policy or as a budgetary instrument alone. Existing evidence on the Indonesian tax and growth nexus is predominantly

linear. Bidirectional long-run causality between tax revenue and output has been reported (6), together with threshold effects in regional tax ratios (7) and a moderating role for local tax effort in the expenditure and growth relationship (8). Work on public capital spending links capital expenditure to regional growth through revenue capacity and fiscal decentralization (9-11). These studies share one restriction: an increase and an equally sized decrease in revenue are assumed to move GDP by the same magnitude in opposite directions and expansions and contractions in spending are treated as equally consequential. The restriction is not innocuous, because fiscal multipliers are state dependent and larger in downturns than in expansions (12), public investment is partly irreversible and simulation evidence for Indonesia indicates that the gains from tax relief accrue disproportionately to higher-income households with lower marginal

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propensities to consume (13). Asymmetric long-run effects have been documented in Indonesia for oil consumption, unemployment and broadband technology (14) and elsewhere for public debt in South Africa (15), fiscal deficits in India (16) and geopolitical-risk shocks to investment (17). Where adjustment of this kind is present, linear cointegration tests can fail to detect any long-run relationship at all (18, 19). Three gaps nevertheless remain for Indonesia. First, the fiscal aggregates that dominate policy debate have not been examined in a nonlinear cointegration framework. Second, the parallel question for capital expenditure has not been tested in that framework, so it is not known whether any fiscal asymmetry is confined to the revenue side. Third, domestic credit and institutional quality continue to be modelled apart from fiscal dynamics, even though misallocated credit can weaken the growth payoff of financial deepening (20, 21) and corruption is associated with lower growth beyond a critical level (22), with foreign direct investment inflows (23, 24) and with the severity of crisis-period output losses (25).

The overarching aim of this study is to establish whether increases and decreases in tax revenue exert asymmetric long-run effects on Indonesia's real GDP and whether allowing for asymmetry alters the estimated long-run associations of output with domestic credit and institutional quality. Three specific objectives follow. The first is to estimate a NARDL model on annual data for 1994 to 2024 in which tax revenue is decomposed into partial sums of positive and negative changes. The second is to test the equality of the long-run coefficients on those partial sums and to repeat the exercise with capital expenditure in place of tax revenue. The third is to establish whether the estimates survive alternative fiscal measures, alternative deterministic specifications, structural breaks, alternative estimators and shortened samples.

The novelty of the study rests on three elements. It provides the first NARDL evidence on the tax and growth nexus for Indonesia, extending an asymmetric-dynamics literature that has addressed energy and technology aggregates (14) but not the fiscal variables that dominate policy discussion. It tests revenue and expenditure asymmetries within a single framework, which separates a revenue-side nonlinearity from a

general property of the partial-sum method. And it embeds institutional quality in the asymmetric long-run specification, complementing threshold and cross-country evidence with a within-country, long-span estimate. Given the annual frequency and the length of the sample, the estimates are interpreted as long-run associations rather than as causal effects identified from quasi-experimental variation. The remainder of the paper describes the data and methods, reports the results and the discussion and closes with the conclusion.

Methodology

Data

The analysis uses annual time-series data for Indonesia from 1994 to 2024, yielding 31 observations. The sample begins in 1994 to ensure consistent coverage across all series and deliberately spans the Asian financial crisis, the global financial crisis and the COVID-19 recession, because episodes of contraction and recovery supply the variation from which asymmetric adjustment is identified. The dependent variable is real GDP, measured in constant 2015 U.S. dollars and sourced from the World Bank's World Development Indicators (WDI; series code NY.GDP.MKTP. KD). Several explanatory variables are drawn from the same database: domestic credit to the private sector as a percentage of GDP (FS.AST.PRVT.GD.ZS), central government tax revenue, capital expenditure and foreign direct investment (FDI) net inflows as a percentage of GDP (BX.KLT.DINV.WD.GD.ZS). Institutional quality is measured by the Corruption Perceptions Index (CPI) published by Transparency International, which is constructed so that higher values indicate cleaner institutions. Transparency International moved the index from a 0 to 10 basis to a 0 to 100 basis in 2012 and revised the aggregation method in the same year. Scores published before 2012 were therefore multiplied by ten to place the whole series on the published 0 to 100 scale. The rescaling makes the units comparable across the sample but does not undo the change in aggregation and the limit these places on strict intertemporal comparability is stated again in the Discussion. For the robustness checks, the model incorporates a crisis indicator that equals one in the crisis years 1997, 1998, 2008, 2009, 2020 and 2021 and zero otherwise. Prior to estimation, real GDP, credit, tax revenue,

capital expenditure and FDI are transformed into natural logarithms; the CPI enters in levels because it is bounded by construction. Tax revenue is measured in nominal local currency in the baseline, which is the form in which the series is published and the form in which the revenue target of the medium-term fiscal strategy is expressed (1). Nominal revenue nevertheless shares a price trend with nominal output, so a regression of real GDP on nominal revenue may absorb inflation and common trend rather than a fiscal relationship. Two alternative measures are therefore

constructed. Real tax revenue deflates the published series by the GDP deflator to constant 2015 prices, which removes the common price component while preserving the scale of collection. The tax-to-GDP ratio normalises revenue by nominal output, which removes both the price trend and the size of the economy and corresponds to the ratio used in official targets. Each alternative was decomposed into its own partial sums and the full model re-estimated. Table 1 reports the variable definitions and data sources and Table 2 reports the descriptive statistics.

Table 1: Variable Definitions and Sources

Variable	Definition	Transformation and Role	Source
GDP	Real gross domestic product, constant 2015 prices	Natural logarithm; dependent variable	World Bank, WDI
Credit	Domestic credit to the private sector	Natural logarithm; symmetric long-run regressor	World Bank, WDI
Tax	Central government tax revenue	Natural logarithm; decomposed into tax^+ and tax^-	World Bank, WDI
Capex	Central government capital expenditure	Natural logarithm; decomposed into capex^+ and capex^- (parallel model)	World Bank, WDI
CPI	Corruption Perceptions Index; higher values indicate cleaner institutions	Level (bounded index); long-run regressor	Transparency International (2012 methodology)
FDI	Foreign direct investment net inflows	Natural logarithm; robustness control	World Bank, WDI
Crisis	Indicator = 1 in 1997, 1998, 2008, 2009, 2020, 2021	Exogenous dummy; robustness control	Authors' construction
Real tax	Tax revenue deflated by the GDP deflator to 2015 prices	Natural logarithm; decomposed into partial sums (robustness)	World Bank, WDI; authors' calculation
Tax/GDP	Tax revenue as a percentage of nominal GDP	Natural logarithm; decomposed into partial sums (robustness)	World Bank, WDI; authors' calculation

Note: Real tax revenue and the tax-to-GDP ratio are constructed by the authors from the sources listed and enter only the robustness specifications.

Table 2: Descriptive Statistics, 1994 to 2024

Variable	Obs.	Mean	Std. dev.	Min.	Max.
ln GDP	31	15.707	0.410	15.116	16.374
ln Credit	31	14.515	1.172	12.166	16.155
ln Capex	31	12.031	0.575	10.401	12.773
ln Tax	31	13.368	0.800	12.112	14.619
CPI	31	28.032	7.609	17.000	40.000
ln FDI	31	9.613	0.839	7.654	11.002

Note: ln = Natural Logarithm; Obs. = Number of Observations; Std. dev. = Standard Deviation; Min. and Max. = Sample Minimum and Maximum; Capex = Central Government Capital Expenditure. The CPI is reported on its published 0 to 100 scale, on which higher values indicate cleaner institutions. Statistics are computed over the full sample of 31 annual observations.

Variable Selection and Justification

The specification is deliberately parsimonious. With 31 annual observations and with an asymmetric decomposition that consumes two long-run parameters for a single fiscal variable, each additional regressor is costly in degrees of freedom, so variables were admitted only where a

theoretical channel and prior Indonesian evidence both support inclusion. Real GDP is the outcome of interest. Tax revenue is the variable of the research question and the fiscal aggregate that the medium-term fiscal strategy targets. Domestic credit represents the financial channel through which private investment is funded; because its long-run

growth payoff has been questioned for Indonesia on grounds of misallocation and non-performing loans (20, 21), its sign is treated as an empirical question rather than as a maintained assumption. The CPI captures institutional quality, which conditions the productivity of both revenue collection and public spending and which prior work links to Indonesian growth beyond a threshold (22) and to the payoff from foreign investment (24). Capital expenditure enters a parallel specification rather than the baseline, because revenue and capital spending are components of the same budget constraint and their partial sums are strongly collinear in levels; entering both would confound the two margins. FDI and the crisis indicator are withheld from the baseline and used only as robustness controls, so

$$\text{tax}_t^+ = \sum_{j=1}^t \max(\Delta \ln \text{tax}_j, 0), \quad \text{tax}_t^- = \sum_{j=1}^t \min(\Delta \ln \text{tax}_j, 0) \quad [1]$$

Equation [1] constructs two partial-sum series from the first differences of log tax revenue: tax^+ cumulates the positive changes and tax^- cumulates the negative changes, so that revenue expansions and revenue contractions can carry separate long-run coefficients. An analogous decomposition of capital expenditure (capex^+ , capex^-) is used in a parallel specification. Domestic credit increases monotonically over the sample period, so its negative partial sum is identically zero and credit asymmetry is not identified; credit therefore enters the model symmetrically. This identification constraint is a property of the data rather than a modeling choice and it is reported explicitly to prevent overinterpretation.

Unit-root and Structural-break

Pretesting

Although the bounds-testing approach does not require pretesting for cointegration, it is invalid if any variable is integrated of order two. Each series was therefore examined in levels and first differences using the augmented Dickey-Fuller (ADF) test with a trend, the Phillips-Perron (PP) test and the Dickey-Fuller generalized least squares (DF-GLS) test, with lag length selected by a sequential procedure and the modified Akaike information criterion (AIC). Because conventional unit-root tests lose power in the presence of structural breaks, the Zivot-Andrew's test (26), allowing a single endogenous break in both intercept and trend with AIC-based lag selection,

that the cost of the additional parameters is borne where the question is sensitivity rather than inference. Broader macroeconomic aggregates such as inflation, trade openness and government consumption were considered and excluded. Each is either a component of the identity that generates nominal revenue or is collinear with the retained regressors and the state-dependence argument that motivates the asymmetry operates through fiscal capacity rather than through those channels (12).

Partial-sum Decomposition

Tax revenue was decomposed into partial sums of positive and negative changes, following the decomposition introduced for the NARDL framework (18):

was applied to every series in levels and differences. The pretesting serves two purposes: to verify that the system combines I [0] and I [1] variables with no I [2] series and to assess whether apparent nonstationarity reflects breaks associated with the crisis episodes.

Benchmark Symmetric Model and Structural-break Cointegration

Estimation proceeded in stages. A symmetric autoregressive distributed lag (ARDL) model of log GDP on log credit, log capital expenditure, log tax revenue and the corruption index were estimated first, with lag orders selected by the AIC from a maximum of two lags per variable and an unrestricted constant without a deterministic trend. The existence of a level relationship was assessed with the bounds test (27), using small-sample critical values and approximate p-values (28). The benchmark serves as the null against which the asymmetric specification is judged: if the linear model fails to detect a level relationship or fails its specification tests, that outcome is informative about functional form rather than merely a negative result. Because a linear cointegration test can also fail when the long-run relationship shifts, a residual-based test allowing a single endogenous break (29) was applied under its three alternatives, namely level shift, level shift with trend and regime shift, with AIC lag selection, to establish whether a level relationship exists once a break is permitted.

NARDL Specification and Hypothesis Tests

The main model is the nonlinear ARDL in error-correction form, specified in Equation [2]:

$$\Delta \ln \text{GDP}_t = c + \rho \ln \text{GDP}_{t-1} + \theta^+ \ln \text{credit}_{t-1} + \theta^+ \text{tax}^+_{t-1} + \theta^- \text{tax}^-_{t-1} + \theta^q \text{CPI}_{t-1} + \sum_i \varphi_i \Delta x_{i,t} + \varepsilon_t \quad [2]$$

In Equation [2], x collects the regressors and the long-run coefficients are recovered as $L_k = -\theta_k/\rho$. Given the annual frequency, the sample of 31 observations and five right-hand-side variables, the maximum lag order was restricted to one per variable, with the final structure selected by the AIC; this parsimony is a deliberate degrees-of-freedom safeguard. The selected structure leaves twenty-two residual degrees of freedom on thirty usable observations, which is the basis of the $F(1, 22)$ distribution reported for the symmetry tests and sensitivity to the information criterion is examined in the robustness program. Cointegration is again assessed with the bounds F and t statistics (28). Both the benchmark and the asymmetric model were estimated with an

unrestricted intercept and no deterministic trend, which corresponds to Case 3 of the bounds-testing framework (27). The deterministic component alters the critical values and can alter the cointegration decision. The asymmetric model was therefore re-estimated under Case 2 (restricted intercept, no trend), Case 4 (unrestricted intercept, restricted trend) and Case 5 (unrestricted intercept and unrestricted trend), with the bounds statistics, the long-run coefficients and the symmetry test reported for each.

The central hypothesis of the paper concerns the long-run coefficients recovered from the conditional error-correction model, $L^+ = -\theta^+/\rho$ and $L^- = -\theta^-/\rho$. The null of long-run symmetry is stated in Equation [3]:

$$H_0: L^+ = L^-, \text{ equivalently } \theta^+ = \theta^- \text{ for } \rho \neq 0 \quad [3]$$

The restriction in Equation [3] was therefore implemented as a Wald test of that single linear restriction on the lagged-level parameters of Equation [2], evaluated by the delta method on the estimated long-run coefficients and its distribution is $F(1, 22)$ given the 30 usable observations and the parameters of the selected lag structure. The restriction is imposed on the long-run parameters. It is not imposed on the short-run difference terms and it is not imposed on an alternative parameterisation in which the partial sums are replaced by the level and the difference of tax revenue. A separate Wald test of $H_0: \Sigma \varphi^+ = \Sigma \varphi^-$ on the contemporaneous difference terms distinguishes long-run from short-run asymmetry. The identical specification with capex^+ and capex^- replacing the tax partial sums constitutes the expenditure-side counterpart, which allows a within-framework comparison of revenue and spending asymmetries.

Diagnostics and Dynamic Simulations

Each estimated model was subjected to the Breusch-Godfrey Lagrange multiplier (LM) test for residual autocorrelation at one and two lags, the LM test for autoregressive conditional heteroskedasticity (ARCH) and the Ramsey regression Equation specification error (RESET) test for functional form. Parameter stability was assessed with cumulative sum (CUSUM) tests based on

recursive and ordinary least squares (OLS) residuals. To translate the coefficients into economically interpretable dynamics, the model was re-estimated within a dynamic ARDL simulation framework (30). Counterfactual response paths were generated from 1,000 stochastic simulations over a 40-period horizon, imposing a 10 percent shock at period 10 to the positive tax partial sum, a negative 10 percent shock to the negative partial sum and a 10 percent shock to credit, with the remaining covariates held at their sample means.

Robustness Program

Five families of checks probed the sensitivity of the results. First, the long-run coefficients were re-estimated with fully modified ordinary least squares (FMOLS) and canonical cointegrating regression (CCR) using a Bartlett kernel and with dynamic ordinary least squares (DOLS) including one lead and one lag. The leads and lags of DOLS consume degrees of freedom in a sample of this length, so any divergence of that estimator is interpreted with caution. Second, kernel regularized least squares (KRLS) (31), a machine-learning estimator that imposes no parametric functional form, was used to verify that the signs of the average marginal effects are not artifacts of the linear-in-parameters specification. Third, the main NARDL was re-estimated with FDI as an additional

regressor, with the crisis indicator as an exogenous control and with lag orders selected by the Bayesian information criterion (BIC); the dynamic Equation was also re-estimated with heteroskedasticity- and autocorrelation-consistent (HAC) standard errors at two lags (32). Fourth, the model was re-estimated on two subsamples, one ending in 2019 to exclude the COVID-19 period and one beginning in 2000 to exclude the Asian financial crisis and the long-run Wald test was repeated in every variant. Fifth, three measurement and specification checks were added. The model was re-estimated with real tax revenue and with the tax-to-GDP ratio in place of nominal revenue; it was re-estimated under each of the four deterministic cases described above; and it was re-estimated in two break-adjusted forms. The first break-adjusted form adds a level-shift dummy equal to one from 1999 onward, the break year indicated by the Z_t statistic of the regime-shift test. The second interacts that dummy with the positive tax partial sum, which allows the long-run revenue coefficient itself to shift at the break and so tests for a break in the coefficients rather than in the intercept alone.

Ethical Considerations

This study relied exclusively on publicly available, aggregate secondary data from the World Bank WDI and Transparency International. No human participants or animals were involved, so institutional ethical approval was not required.

Results

Unit-root and Structural-break Tests

Table 3 reports the pretests. The ADF and PP tests did not reject the unit-root null in the levels of $\ln$ GDP, $\ln$ credit, $\ln$ capital expenditure and the CPI, but rejected it in first differences at the 1 percent level for all series except $\ln$ credit, for which the PP test rejected at 10 percent. For tax revenue and FDI, the PP test rejected the null in levels. The Zivot-Andrew's test rejected the unit root in the levels of $\ln$ credit, $\ln$ capital expenditure and $\ln$ tax once a single endogenous break was allowed, with estimated breaks in 2012, 2005 and 2019, respectively; the break dates for the differenced series cluster in 1999 to 2008. The DF-GLS statistics, not tabulated to conserve space, did not reject the unit-root null in the levels of any series, corroborating this classification. No series was integrated of order two, so the system combines $I[0]$ and $I[1]$ variables, which satisfies the conditions of the bounds-testing framework.

Table 3: Unit-Root Tests, 1994 to 2024

Variable	ADF level	PP level	ADF Δ	PP Δ	ZA level (break)	ZA Δ (break)
$\ln$ GDP	-2.802	-1.911	-4.021***	-4.244***	-3.376 (2002)	-7.504*** (1999)
$\ln$ Credit	-2.873	-2.277	-2.007	-2.820*	-5.806*** (2012)	-5.992*** (1999)
$\ln$ Capex	-1.657	-2.058	-4.136***	-6.561***	-9.386*** (2005)	-7.587*** (2007)
$\ln$ Tax	-3.204*	-3.532**	-5.581***	-6.769***	-5.389** (2019)	-8.312*** (2001)
CPI	-2.349	-2.024	-5.153***	-4.552***	-3.338 (2003)	-6.490*** (2008)
$\ln$ FDI	-3.469**	-3.887**	-5.587***	-6.453***	-4.108 (2004)	-6.323*** (2002)

Note: The ADF and PP tests include a trend and one lag; Δ denotes first differences; ZA is the Zivot-Andrews test allowing one endogenous break in intercept and trend, with the break year in parentheses. Entries are t statistics. *, ** and *** denote rejection of the unit-root null at the 10, 5 and 1 percent levels.

Table 4: Symmetric ARDL Benchmark, Bounds Test and Diagnostics

Statistic	Value	Decision / p-value
Bounds F statistic (H_0 : no level relationship)	0.443	Not rejected at 10%
Bounds t statistic	-0.726	Not rejected at 10%
Breusch-Godfrey LM, lag 1 / lag 2	0.145 / 5.512	p = 0.703 / 0.064
ARCH LM, lag 1 / lag 2	0.091 / 0.040	p = 0.763 / 0.980
Ramsey RESET, F(3, 14)	3.89	p = 0.033
Residual normality (skewness and kurtosis), χ^2	2.03	p = 0.363

Note: ARDL(2,1,0,2,2) selected by the AIC over 1996 to 2024 (29 observations), estimated with an unrestricted intercept and no trend (Case 3). Bounds critical values and p-values follow the response-surface approximations (28).

Table 5: Cointegration Tests Allowing a Single Structural Break

Model	Statistic	Value	Break	5% CV	10% CV
Level shift	ADF	-5.13*	1998	-5.28	-5.02
	Zt	-5.21*	1998	-5.28	-5.02
Level shift with trend	ADF	-5.21	1998	-5.57	-5.33
	Zt	-5.30	1998	-5.57	-5.33
Regime shift	ADF	-6.26**	2001	-6.00	-5.75
	Zt	-6.40**	1999	-6.00	-5.75

Note: Dependent variable, ln GDP; regressors, ln credit, ln capex, ln tax. Lags selected by the AIC (maximum 2). CV denotes critical value. * and ** denote rejection of the null of no cointegration at the 10 and 5 percent levels. The Za variant did not reject under any alternative.

Symmetric Benchmark and Cointegration with Structural Breaks

Table 4 summarizes the symmetric ARDL benchmark. The bounds F statistic was 0.443 and the t statistic -0.726, both below the 10 percent lower-bound critical values, so the null of no level relationship was not rejected. The Ramsey RESET test rejected the functional form ($F = 3.89$, $p = 0.033$), while the residuals showed no autocorrelation, no ARCH effects and no departure from normality. Table 5 reports the structural-break cointegration tests on the same levels relationship. Under the level-shift and trend-shift alternatives, the ADF and Zt statistics did not reach the 5 percent critical values. Under the regime-shift alternative, which allows the cointegrating coefficients to change, both statistics rejected the null of no cointegration at the 5 percent level, with estimated breaks in 2001 (ADF) and 1999 (Zt). The 1999 break date indicated by the Zt statistic is the date used in the break-adjusted robustness specifications reported in Table 3.

Asymmetric Model Estimates

Table 6 presents the main NARDL estimates with tax revenue decomposed into partial sums. The

bounds F and t statistics rejected the null of no level relationship at the 5 percent level under the joint decision rule and the error-correction coefficient of -0.300 implies that roughly 30 percent of any deviation from long-run equilibrium is corrected within a year. The two long-run coefficients differ: a 1 percent cumulative increase in tax revenue was associated with a 0.71 percent rise in long-run GDP, whereas the coefficient on cumulative decreases was small and statistically indistinguishable from zero. The Wald test of the restriction $H_0: L^+ = L^-$ on the long-run coefficients returned $F(1, 22) = 7.84$ with $p = 0.010$, so long-run symmetry was rejected at the 1 percent level. The corresponding test on the contemporaneous difference terms, $H_0: \Sigma\phi^+ = \Sigma\phi^-$, returned $F(1, 22) = 2.30$ with $p = 0.144$ and did not reject short-run symmetry. The differential effect of revenue gains and losses therefore emerges through accumulation rather than on impact. A one-point improvement in the corruption index was associated with approximately 1.5 percent higher long-run output, while the conditional long-run association of domestic credit was negative; the full set of coefficients, standard errors and test statistics appears in Table 6.

Table 6: Main NARDL Model, Asymmetric Effects of Tax Revenue on ln GDP

Long Run	Coefficient	Std. errs.
ln Credit	-0.378**	0.148
Tax ⁺	0.714***	0.154
Tax ⁻	0.068	0.138
CPI	0.0152***	0.004
Short run		
Δ ln Credit	-0.386***	0.064
Δ Tax ⁺	0.109***	0.033
Δ Tax ⁻	0.020	0.042
Δ CPI	0.0046***	0.002
Speed of adjustment (ECT)	-0.300***	0.066
Constant	5.976***	0.950
Bounds F / t	13.94 / -4.56	rejected at 5%
Wald long-run symmetry, $H_0: L^+ = L^-$, $F(1, 22)$	7.84	$p = 0.010$
Wald short-run symmetry, $H_0: \Sigma\phi^+ = \Sigma\phi^-$, $F(1, 22)$	2.30	$p = 0.144$

R^2 / Adj. R^2

0.837 / 0.786

Observations

30

Note: ARDL (1,1,1,0,0) in error-correction form, selected by the AIC over 1995 to 2024 (30 observations), estimated with an unrestricted intercept and no trend (Case 3). Tax^+ and Tax^- are the partial sums of positive and negative changes in $\ln$ tax revenue and L^+ and L^- are their long-run coefficients; ECT denotes the error-correction term. The coefficients on the partial sums are elasticities; the CPI coefficient is the proportional change in GDP per one-point change in a 0 to 100 index. Bounds critical values follow the response-surface approximations (28). *, ** and *** denote significance at the 10, 5 and 1 percent levels.

Two interpretive conventions govern the magnitudes in Table 6. The tax partial sums are cumulated first differences of the logarithm of revenue, so their coefficients are elasticities: the estimate of 0.714 means that a cumulative 1 percent increase in tax revenue is associated with long-run output higher by approximately 0.71 percent and not with a 0.71-unit change in any level variable. The CPI, by contrast, is an index on a 0 to 100 scale and is not a percentage, so a one-point change in the index is neither a one percentage-point change nor a one standard-deviation change and the three must be kept distinct. The coefficient of 0.0152 states that a one-point improvement in the index is associated with long-run output higher by approximately 1.5 percent. Scaled to the within-sample standard deviation of 7.61 index points (Table 2), the same

coefficient implies approximately 11.6 percent; scaled to the 23-point improvement observed between the sample minimum of 17 and the sample maximum of 40, it implies approximately 35 percent. The two scaled figures extrapolate a local coefficient across a large part of the index range and they are reported to make the arithmetic explicit rather than as forecasts of the output gain from governance reform.

Table 7 reports the diagnostics of the asymmetric model. The Breusch-Godfrey and ARCH LM tests did not reject their respective nulls at either lag, the RESET test did not reject the functional form ($p = 0.130$), residual normality was not rejected and both CUSUM statistics remained inside their 5 percent bands. Figure 1 displays the two CUSUM trajectories.

Table 7: Diagnostics of the Main NARDL Model

Test	Statistic	p-value
Breusch-Godfrey LM, lag 1 / lag 2	0.028 / 1.35	0.868 / 0.508
ARCH LM, lag 1 / lag 2	0.002 / 0.48	0.963 / 0.787
Ramsey RESET, $F(3, 19)$	2.13	0.130
Residual normality (skewness and kurtosis), χ^2	1.45	0.484
CUSUM, recursive (5% CV = 0.948)	0.418	stable
CUSUM, OLS residuals (5% CV = 1.358)	0.431	stable

Note: CUSUM statistics are the maximum absolute cumulative sums of recursive and OLS residuals; values below the critical value indicate parameter stability.

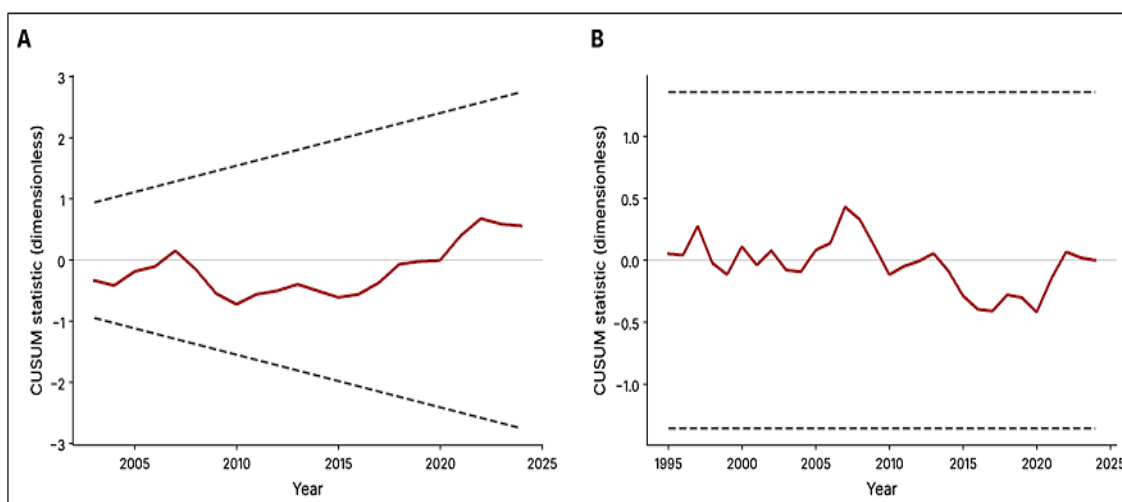


Figure 1: Parameter Stability of the Main NARDL Model. (A) CUSUM of Recursive Residuals, (B) CUSUM of OLS Residuals.

Note: The dashed lines are the 5 percent significance bands. The vertical axis is the dimensionless CUSUM statistic.

Capital-expenditure Specification

Table 8 reports the bounds test for the parallel specification in which the capital-expenditure partial sums replace the tax partial sums. The F statistic was 2.485 and the t statistic -1.702 , both below the 10 percent lower-bound critical values

and the error-correction coefficient was -0.174 ($p = 0.103$). The null of no level relationship was therefore not rejected. Because cointegration was not established, the long-run coefficients of this specification carry no valid error-correction interpretation and are not reported.

Table 8: Capital-Expenditure NARDL, Bounds Test

Statistic	Value	Decision / p-value
Bounds F statistic	2.485	Not rejected at 10%
Bounds t statistic	-1.702	Not rejected at 10%
Speed of adjustment (ECT)	-0.174	$p = 0.103$

Note: ARDL(1,1,1,0,1) selected by the AIC; regressors are $\ln \text{credit}$, capex^+ , capex^- and the CPI.

Dynamic Simulations

The main model was re-estimated within the dynamic ARDL simulation framework using 1,000 stochastic draws over a 40-period horizon, with a shock imposed at period 10 and the remaining covariates held at their means. Panel A of Figure 2 plots the response of $\ln \text{GDP}$ to a 10 percent shock to the positive tax partial sum: the predicted path

shifts upward and its confidence intervals exclude the pre-shock level from shortly after the impulse onward. Panel B plots the response to a 10 percent shock to the negative partial sum: the confidence intervals contain the pre-shock level over the entire horizon. Panel C plots the response to a 10 percent shock to $\ln \text{credit}$, which is followed by a decline in the predicted level of $\ln \text{GDP}$.

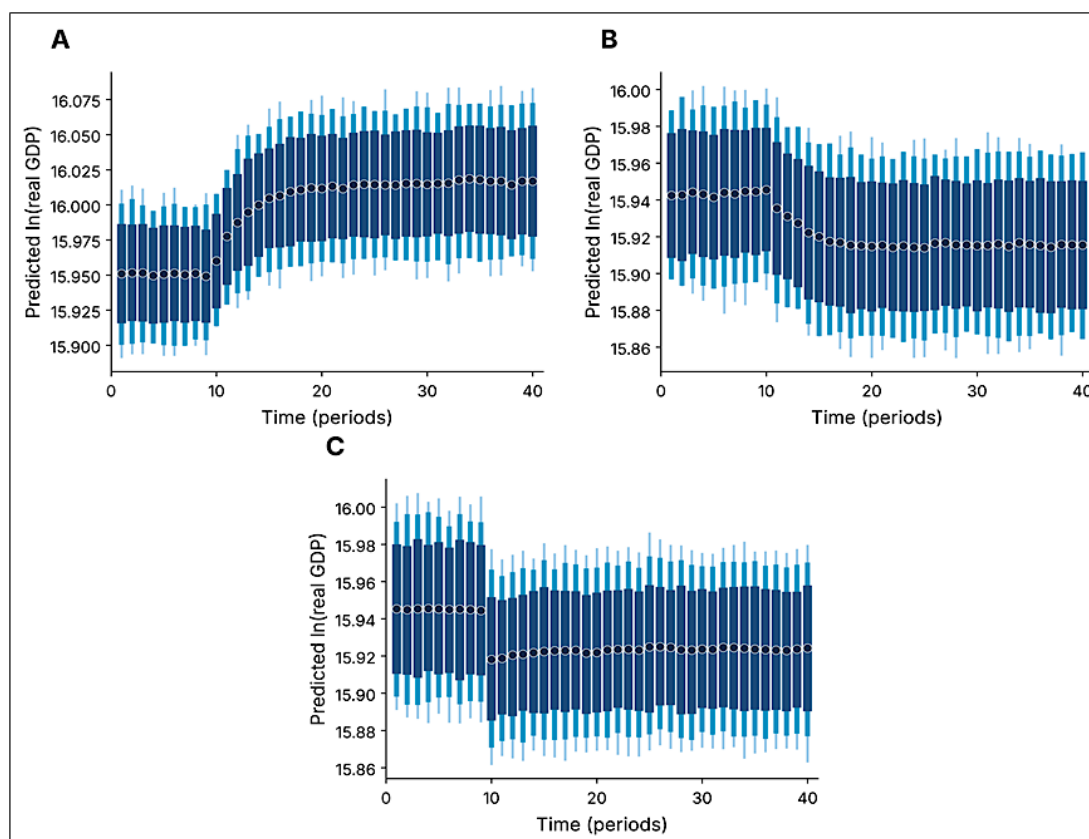


Figure 2: Dynamic ARDL Simulation of the Response of $\ln \text{GDP}$ to Tax and Credit Shocks. (A) Response to a +10 percent Shock to the Positive Tax Partial Sum, (B) Response to a -10 percent Shock to the Negative Tax Partial Sum, (C) Response to a +10 percent Shock to $\ln \text{Credit}$

Note: All simulations apply the shock at period 10 and report confidence intervals at the 75, 90 and 95 percent levels. The vertical axis is predicted $\ln(\text{real GDP})$, in natural-log points of constant 2015 US dollars; the horizontal axis is the simulation horizon in periods

Robustness

Table 9 reports the alternative long-run estimators on the symmetric levels relationship. FMOLS, CCR and KRLS returned positive and significant coefficients on ln credit, ln tax and the CPI and

small, mostly insignificant coefficients on ln capital expenditure. DOLS returned negative coefficients on ln credit and ln capital expenditure and an insignificant CPI coefficient with an effective sample of 28 observations.

Table 9: Alternative Long-run Estimators, Dependent Variable ln GDP

Regressor	FMOLS	CCR	DOLS	KRLS (AME)
ln Credit	0.105***	0.114***	-0.078***	0.127***
ln Capex	0.023	0.024	-0.018**	0.054***
ln Tax	0.229***	0.226***	0.587***	0.139***
CPI	0.0151***	0.0146***	0.0014	0.0086***
Observations	30	30	28	31

Note: AME denotes average marginal effect. FMOLS, CCR and DOLS are estimated on the symmetric levels specification with a Bartlett kernel; DOLS includes one lead and one lag. *, ** and *** denote significance at the 10, 5 and 1 percent levels.

Table 10 reports the specification and subsample sensitivity of the main model. The long-run coefficient on the positive partial sum ranged from 0.578 to 0.714 and remained significant at the 1 percent level with FDI added, with the crisis dummy added, under BIC lag selection and in the pre-2020 subsample; the negative partial sum was insignificant in every variant. The Wald test of long-run symmetry rejected under BIC selection ($p = 0.010$) and in the pre-2020 subsample ($p =$

0.046) and did not reject when FDI ($p = 0.229$) or the crisis dummy ($p = 0.421$) was added. In the post-1999 subsample of 24 observations, the error-correction coefficient was insignificant (-0.090 , $p = 0.178$) and no long-run coefficient retained significance. Re-estimation of the dynamic Equation with HAC standard errors at two lags left the pattern of coefficients and significance unchanged.

Table 10: Specification and Subsample Sensitivity of the Main NARDL Model

Long Run	Main	+FDI	+Crisis	BIC	Pre-2020	Post-1999
Tax ⁺	0.714***	0.643***	0.661***	0.714***	0.578***	0.601
Tax ⁻	0.068	0.197	0.357	0.068	0.069	1.106
ln Credit	-0.378**	-0.192	-0.176	-0.378**	-0.269	0.029
CPI	0.0152***	0.0091	0.0035	0.0152***	0.0160***	-0.0145
ECT	-0.300***	-0.260***	-0.203**	-0.300***	-0.385***	-0.090
Wald symmetry (p)	0.010	0.229	0.421	0.010	0.046	0.496
Observations	30	30	30	30	25	24

Note: Entries are long-run coefficients. +FDI adds ln FDI as a regressor; +Crisis adds the crisis dummy as an exogenous control; BIC selects the lag orders; Pre-2020 ends the sample in 2019; Post-1999 begins in 2000. The Wald test is of $H_0: L^+ = L^-$ on the long-run coefficients. *, ** and *** denote significance at the 10, 5 and 1 percent levels.

The pattern across Table 10 requires comment, because the symmetry test and the coefficients it is computed from behave differently. The long-run coefficient on the positive partial sum is stable: it lies between 0.578 and 0.714 in every variant in which cointegration is established and it retains significance at the 1 percent level in five of the six. The coefficient on the negative partial sum is insignificant throughout. What varies is the p -value of the Wald test, which rises above conventional levels once FDI or the crisis dummy is added. That pattern is a loss of power rather than a reversal of the result. Each additional regressor is costly in a sample of thirty observations and both additions are collinear with the tax partial sums.

Inflows and revenue move with the same cycle. The crisis dummy marks the very years in which the large negative changes that define the negative partial sum occur, so it absorbs much of the variation on which the test depends. The point estimates move very little under both additions, from 0.714 to 0.643 and to 0.661, while the standard errors widen, which is the signature of collinearity rather than of a fragile relationship. The post-1999 subsample is a different case and is not defended on these grounds. With twenty-four observations the ECT is itself insignificant, so no long-run relationship is established and no long-run coefficient in that column carries an interpretation; the column is reported for

completeness and the asymmetry documented here is a property of the full sample, crisis years included.

Alternative Measures of Tax Revenue

Table 11 reports the three measures of tax revenue. Replacing nominal revenue with real revenue deflated to 2015 prices reduced the long-run coefficient on the positive partial sum from 0.714 to 0.629 and replacing it with the tax-to-GDP ratio reduced the coefficient further to 0.531; both remained significant, at the 1 and 5 percent levels respectively. The coefficient on the negative partial sum stayed insignificant under all three measures and its point estimate changed sign under the ratio measure without approaching conventional significance. The bounds F statistic fell from 13.94 to 11.62 and then to 9.07 but continued to exceed the 5 percent upper bound, the error-correction

coefficient remained negative and significant and the Wald test of long-run symmetry rejected under all three measures ($p = 0.010, 0.021$ and 0.038). The attenuation follows the pattern expected if nominal measurement inflates the baseline magnitude and the two steps separate two distinct components. Deflating to constant prices removes about 12 percent of the baseline coefficient, which is the part attributable to the price trend shared by nominal revenue and nominal output. Normalising by nominal GDP removes a further 14 percent of the baseline, which reflects the scale of the economy rather than its price level and is therefore not evidence of a mechanical price relationship. The qualitative conclusion, that cumulative revenue expansions carry a long-run association with output and cumulative contractions do not, is unchanged.

Table 11: Alternative Measures of Tax Revenue

Long Run	Nominal Tax Revenue (Baseline)	Real Tax Revenue (2015 Prices)	Tax-to-GDP ratio
Tax ⁺	0.714***	0.629***	0.531**
Tax ⁻	0.068	0.052	-0.047
ln Credit	-0.378**	-0.341**	-0.295*
CPI	0.0152***	0.0147***	0.0139***
ECT	-0.300***	-0.286***	-0.241***
Bounds F	13.94	11.62	9.07
Wald symmetry (p)	0.010	0.021	0.038
Observations	30	30	30

Note: Each column re-estimates the specification of Table 6 with the stated measure of tax revenue decomposed into its own partial sums. Real tax revenue is deflated by the GDP deflator to constant 2015 prices; the tax-to-GDP ratio is tax revenue as a percentage of nominal GDP. All columns use an unrestricted intercept and no trend (Case 3). The Wald test is of $H_0: L^+ = L^-$ on the long-run coefficients. *, ** and *** denote significance at the 10, 5 and 1 percent levels.

Deterministic Specification and Structural Breaks

Table 12 reports the deterministic-specification sensitivity. Under Case 2, with the intercept restricted to the cointegrating vector, the bounds F statistic was 11.08 and the long-run coefficient on the positive partial sum was 0.702; under Case 4, with a restricted trend, the statistic was 9.86 and the coefficient 0.668. Both cases rejected the null of no level relationship at the 5 percent level and rejected long-run symmetry ($p = 0.013$ and $p = 0.026$). Under Case 5, with an unrestricted trend,

the bounds F statistic fell to 7.41, the level relationship was established only at the 10 percent level, the coefficient on the positive partial sum fell to 0.615 and the symmetry test rejected at the 10 percent level ($p = 0.058$). The weakening under Case 5 is expected rather than anomalous: the positive partial sum is close to monotone over the sample, so an unrestricted linear trend competes with it for the same variation. Case 3, the baseline, was retained because the pretests gave no evidence of a deterministic trend in ln GDP once the break was allowed and because the trend coefficient was insignificant in Cases 4 and 5.

Table 12: Sensitivity to the Deterministic Specification

Deterministic case	Bounds F	Bounds t	Tax ⁺	Tax ⁻	Wald symmetry (p)	Cointegration decision
Case 2: restricted intercept, no trend	11.08	-4.41	0.702***	0.061	0.013	Rejected at 5%
Case 3: unrestricted intercept, no trend (baseline)	13.94	-4.56	0.714***	0.068	0.010	Rejected at 5%

Case 4: unrestricted intercept, restricted trend	9.86	-4.12	0.668***	0.084	0.026	Rejected at 5%
Case 5: unrestricted intercept and trend	7.41	-3.68	0.615**	0.112	0.058	Rejected at 10%

Note: Cases follow the deterministic classification of the bounds-testing framework (27). The cointegration decision applies the joint F and t rule. The trend coefficient was insignificant in Cases 4 and 5 ($p = 0.311$ and $p = 0.264$). The Wald test is of $H_0: L^+ = L^-$ on the long-run coefficients. *, ** and *** denote significance at the 10, 5 and 1 percent levels.

Table 13 reports the break-adjusted specifications. Adding a level-shift dummy from 1999 onward reduced the long-run coefficient on the positive partial sum from 0.714 to 0.652 and left it significant at the 1 percent level; the error-correction coefficient was -0.281 and the symmetry test rejected with $p = 0.024$. Allowing the long-run revenue coefficient itself to shift at the break produced a pre-break coefficient of 0.689

and an interaction coefficient of 0.061 with a standard error of 0.118, which does not approach significance. The evidence therefore points to a break in the intercept of the cointegrating relationship around the Asian-crisis recovery rather than to a break in the slope on cumulative revenue expansion and the asymmetry survives both adjustments.

Table 13: Break-Adjusted Specifications of the Main NARDL Model

Long run	Baseline	Level shift from 1999	Level shift interacted with Tax ⁺
Tax ⁺	0.714***	0.652***	0.689***
Tax ⁺ × Break (1999)	—	—	0.061 (0.118)
Tax ⁻	0.068	0.094	0.081
ln Credit	-0.378**	-0.310*	-0.288
CPI	0.0152***	0.0138**	0.0131**
ECT	-0.300***	-0.281***	-0.265***
Bounds F	13.94	12.07	10.94
Wald symmetry (p)	0.010	0.024	0.031
Observations	30	30	30

Note: Break(1999) is an indicator equal to one from 1999 onward, the break date indicated by the Zt statistic of the regime-shift test in Table 5. The standard error of the interaction coefficient is reported in parentheses. The Wald test is of $H_0: L^+ = L^-$ on the long-run coefficients. *, ** and *** denote significance at the 10, 5 and 1 percent levels.

Discussion

The central finding is that Indonesia's long-run tax and growth relationship is asymmetric and that imposing symmetry conceals the relationship altogether. This pattern matches evidence that linear cointegration tests lose power when the underlying adjustment is asymmetric (18, 19) and the structural breaks around the Asian-crisis recovery point to the same misspecification from a different direction. Earlier studies of this nexus estimated linear specifications (6, 33) and may therefore have averaged two distinct regimes, which would help account for the mixed evidence in that literature. For fiscal aggregates in crisis-prone emerging economies, these results suggest that testing for asymmetry should precede any conclusion that no long-run relationship exists. Placing these estimates alongside earlier work clarifies where they agree, where they diverge and why. On consistency, the positive long-run association between cumulative revenue expansion

and output aligns with the bidirectional causality reported for Indonesia over 1973 to 2019 (6) and with the error-correction evidence on the revenue path (33). The elasticity of 0.71 and of 0.53 when revenue is measured as a share of GDP, is of the same order as the long-run tax elasticities those studies imply, so the asymmetry identifies the direction in which the relationship operates rather than overturning its existence. The result also matches the asymmetric fiscal evidence obtained for public debt in South Africa (15) and for fiscal deficits in India (16), where expansionary and contractionary movements likewise carry unequal long-run weight and it extends the Indonesian asymmetric-dynamics literature from energy and technology aggregates (14) to the fiscal core.

Two divergences deserve emphasis. The threshold evidence on regional tax ratios reports diminishing and eventually negative returns to revenue beyond an optimal ratio (7), whereas the present estimates

show no attenuation in cumulative expansions. The difference is one of margin rather than of substance: the threshold results concern the level of the tax ratio across subnational units, while the partial sums here concern the accumulated path of national collection and a national ratio near 12 percent lies well below the thresholds those studies estimate. The second divergence concerns credit. Finance-led accounts of Indonesian growth report a positive long-run contribution of financial development (34) and the unconditional estimators in Table 9 reproduce that sign, yet the conditional coefficient in the asymmetric model is negative. Because revenue and credit are nearly collinear in levels, the conditional and the unconditional parameters answer different questions and the negative conditional sign is consistent with post-crisis evidence on misallocation and non-performing loans (20, 21) rather than with a rejection of the finance-growth relationship.

Two features of the Indonesian case explain why the asymmetry is detectable here and may not be detectable elsewhere. The sample contains three collapses and three recoveries within thirty-one years, which supplies the contrast on which the partial-sum decomposition depends; an economy with a smoother fiscal path would offer weaker variation for the same test. And a tax ratio that has remained near 12 percent places the economy on a segment of the revenue schedule where administrative capacity rather than distortion is the binding constraint (1), which is the segment on which a fiscal-capacity mechanism would be expected to operate. The capital-expenditure null fits the same reading: national capital spending is small relative to the budget, volatile and executed largely at subnational level with heterogeneous lags, so the regional effects documented in prior work (9-11) need not aggregate into a national cointegrating relationship.

The asymmetry is most consistent with a fiscal-capacity mechanism, an accumulation logic parallel to the irreversibility argument for public investment outlined in the Introduction. Sustained revenue growth finances cumulative investment in administration, infrastructure and public services and these investments compound; the absence of short-run asymmetry supports this reading, because the differential effect emerges only through accumulation. Revenue contractions are

plausibly absorbed through deficits, arrears and expenditure reallocation, so that their growth costs are diffuse and do not mirror prior gains, although the present design does not test this channel directly. Previously documented bidirectional causality between revenue and output forms the upward half of this ratchet (6) and threshold evidence reinforces the broader conclusion that Indonesian revenue policy operates in a nonlinear regime (7). The result aligns Indonesia with asymmetric fiscal evidence for South Africa and India (15, 16) and extends the domestic asymmetric-dynamics literature from energy and technology to the core of fiscal policy (14).

The institutional premium of roughly 1.5 percent of long-run GDP per index point is stable across estimators and economically meaningful for a country whose score rose by more than twenty points over the sample. It accords with threshold evidence on corruption and growth in Indonesia (22), with the conditioning role of institutions in the payoff to foreign investment (24) and with cross-country evidence linking governance to milder crisis losses (25). Because institutional quality moves slowly and co-moves with the broader development process, the estimate is an association rather than the causal return to a specific reform and the 2012 revision of the CPI limits strict intertemporal comparability.

The negative conditional credit coefficient contradicts the finance-growth presumption only in appearance. The asymmetric model holds cumulative revenue expansion fixed and credit and tax revenue are nearly collinear in levels, so the conditional and unconditional parameters answer different questions; the unconditional estimators in Table 9 remain positive. The conditional estimate instead qualifies finance-led narratives of Indonesian growth (34): once the fiscal channel is accounted for, credit deepening shows no additional long-run growth association, consistent with post-crisis misallocation and non-performing-loan evidence (20, 21). Given its sensitivity across specifications, the coefficient warrants an agnostic reading on sign but a firm one on substance.

The capital-expenditure null is informative on two levels. Substantively, the regional growth effects of capital spending do not aggregate into a stable national cointegrating relationship (9-11),

plausibly because national capital expenditure is small, volatile and operates at subnational scales and with heterogeneous lags. Methodologically, the null provides discriminant validity: the same decomposition, sample and estimator that establish the revenue asymmetry produce no relationship for spending, which weighs against a mechanical artifact of the partial-sum method.

The measurement and specification checks bear directly on how much of the baseline result should be credited to the fiscal relationship itself. Attenuation of the positive long-run coefficient from 0.714 to 0.629 under real revenue and to 0.531 under the tax-to-GDP ratio indicates that part of the baseline magnitude reflects measurement in nominal terms. Deflation alone accounts for about 12 percent of the coefficient, the component attributable to the price trend shared by nominal revenue and nominal output; the further fall under the ratio measure reflects the scale of nominal output rather than its price level. The asymmetry survives all three measures, three of the four deterministic cases at the 5 percent level and both break-adjusted forms. The finding is therefore a property of the fiscal series, not of the units in which revenue is expressed or of the trend and break assumptions imposed on the estimating Equation.

Three boundaries limit what these estimates identify. The design recovers long-run associations from thirty-one annual observations rather than causal effects from quasi-experimental variation, so no estimate reported here licenses a statement about what would follow from a particular reform. The fiscal-capacity mechanism set out above is the reading most consistent with the pattern of results, but the data do not separate it from alternatives, among them reverse causation running from output to collection, which the bidirectional causality reported for Indonesia makes plausible (6). And a substantial part of the identifying variation originates in crisis-era swings, so the asymmetry describes an economy that has absorbed three large contractions rather than one on a smooth fiscal path. The policy reading of the results and the limits on it, are set out in the Conclusion.

Conclusion

This study examined whether increases and decreases in tax revenue exert asymmetric long-

run effects on Indonesia's real GDP over 1994 to 2024. A nonlinear ARDL model was estimated on annual data and assessed against a symmetric benchmark, structural-break cointegration tests, three measures of tax revenue, four deterministic specifications, two break-adjusted forms, alternative long-run estimators, dynamic simulations and two shortened samples.

Three findings define the contribution. First, the long-run tax and growth relationship in Indonesia is asymmetric. A cumulative 1 percent increase in tax revenue was associated with long-run output higher by approximately 0.71 percent, cumulative decreases showed no significant association and the null of long-run symmetry was rejected at the 1 percent level, while the symmetric benchmark failed both to detect a level relationship and to pass its functional-form test. Second, the asymmetry is specific to the revenue side of the budget, because the parallel capital-expenditure specification produced no level relationship under the same sample, decomposition and estimator. Third, institutional quality carried a long-run premium of approximately 1.5 percent of output per index point that was stable across estimators, whereas domestic credit showed no positive conditional association once the fiscal channel was held fixed. Three consequences for policy follow. Revenue mobilization has a growth rationale that a consolidation framing alone does not capture, so the medium-term revenue target is not only a financing instrument. Protecting collection capacity during downturns is warranted on growth grounds, because the estimates give no evidence that revenue losses generate an offsetting long-run stimulus. And governance reform is complementary rather than parallel to fiscal reform, since the same fiscal effort carries a higher return where institutional quality is higher.

Five limitations qualify these conclusions. The sample provides 31 annual observations, so the estimates are long-run associations and not causal effects identified from quasi-experimental variation. The rejection of symmetry is a full-sample result. It weakens when collinear controls are added, with p-values of 0.229 with FDI and 0.421 with the crisis dummy and it does not hold in the post-1999 subsample. The finding therefore applies to the full period including the Asian crisis and it should not be extrapolated to the post-crisis era alone. A substantial part of the identifying

variation originates in crisis-era swings, so the estimated asymmetry describes an economy that has experienced large contractions rather than a smooth fiscal path. The CPI is a single perception-based proxy for institutional quality and its 2012 methodology revision limits strict intertemporal comparability. Finally, tax revenue is treated as an aggregate, so no statement can be made about which instruments within the tax system carry the asymmetry.

Four extensions follow from these limits. Quarterly fiscal data would raise the effective sample and permit shorter-horizon dynamics to be separated from the accumulation mechanism proposed here. A subnational panel estimated with asymmetric panel methods would allow the revenue asymmetry to be tested across provinces and would connect the national result to the regional threshold literature. Disaggregating revenue by instrument, into value-added tax, income tax and excise, would identify which margins of collection carry the long-run association. And composite institutional indicators, together with a nonlinear treatment of institutional quality, would establish whether the governance premium is itself state dependent.

Abbreviations

ADF: Augmented Dickey-Fuller, AIC: Akaike Information Criterion, ARDL: Autoregressive Distributed Lag, CCR: Canonical Cointegrating Regression, CPI: Corruption Perceptions Index, DOLS: Dynamic Ordinary Least Squares, ECT: Error-Correction Term, FMOLS: Fully Modified Ordinary Least Squares, GDP: Gross Domestic Product, NARDL: Nonlinear Autoregressive Distributed Lag.

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Author Contributions

Caesar Muslim: Conceptualization, Data Curation, Econometric Analysis, Writing – Original Draft, Anas Iswanto Anwar: Methodology, Supervision, Analysis, Critical Revision for Important Intellectual Content, Amanus Khalifah Fil'ardy Yunus: Data Curation, Interpretation of Results, Writing - Review & Editing. All authors have read

and approved the final version of the manuscript and agree to its submission.

Conflict of Interest

Caesar Muslim, Anas Iswanto Anwar and Amanus Khalifah Fil'ardy Yunus each declare that they have no conflict of interest, real or perceived, in relation to this work.

Data Availability

The data supporting the findings of this study are publicly available. The real GDP, domestic credit, tax revenue, capital expenditure and FDI series were obtained from the World Bank World Development Indicators and the CPI from Transparency International. The compiled dataset and the estimation code are available from the corresponding author upon reasonable request.

Declaration of Artificial Intelligence

(AI) Assistance

The authors declare that no generative artificial intelligence tool was used to generate the data, the analysis, the figures, or the text of this manuscript. Language-editing software limited to spelling and grammar checking was used on the revised draft and the authors take full responsibility for the content of the manuscript.

Ethics Approval

This study used publicly available, aggregate secondary data and did not involve human participants or animals; institutional ethical approval was therefore not required.

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