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Do Foreign Direct Investment and Foreign Remittances Drive Economic Growth in Bangladesh? New Evidence from an Augmented ARDL Approach, 1990–2025

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Abstract

Foreign direct investment and remittances are widely credited with sustaining Bangladesh's growth, yet the time-series evidence is mixed and often rests on bivariate models, short samples and incomplete cointegration testing. This study re-examines their growth effects using World Development Indicators data for 1990–2025, measured in constant 2015 US dollars, within an augmented production function that controls for gross fixed capital formation and labor. Long- and short-run effects are estimated with the autoregressive distributed lag bounds-testing approach, using both the F and t statistics with asymptotic and simulated small-sample critical values. A model containing only output, foreign direct investment and remittances shows no valid long-run relationship. In the augmented model, a stable long-run relationship exists ($F = 8.25$; $t = -4.22$), with about 17% of any disequilibrium corrected each year, but it is anchored by domestic fixed investment (elasticity 0.97). The long-run elasticities of foreign direct investment (-0.003, $p = 0.78$) and remittances (-0.039, $p = 0.25$) are statistically insignificant, as are their short-run effects. Three alternative cointegrating estimators, a pandemic dummy, alternative control variables and Toda–Yamamoto causality tests reveal no positive effect of either inflow. Recursive estimates show that the positive remittance elasticity visible in samples ending before 2016 has since disappeared. Policy should therefore aim at converting external inflows into productive domestic investment rather than at maximizing inflow volumes.

Keywords: Foreign direct investment, Remittances, Economic growth, ARDL bounds test, Capital formation, Bangladesh

JEL classification: C22, F21, F24, O47, O53

Introduction

Bangladesh has been among the fastest-growing economies in the developing world, with real GDP expanding by an average of about 5.5% a year between 1991 and 2025 (World Bank 2026). Its growth has rested on export-oriented garment manufacturing (Islam 2021), rising domestic investment and large inflows of foreign exchange from migrant workers. Foreign remittances reached US$33.9 billion in 2025, equivalent to 7.4% of GDP and roughly nineteen times the US$1.8 billion of net foreign direct investment (FDI) received in the same year (World Bank 2026). FDI, by contrast, has remained small by regional standards, averaging about 0.5% of GDP over 2016–2025. Remittances support household consumption and cover close to half of the country's import payments; inflows passed US$30 billion in fiscal year 2024–25, helped by a shift from informal *hundi* networks to banking channels after exchange-rate adjustments, higher incentives and enforcement against informal operators (Khondker and Majid 2025). Yet real GDP growth slowed to 4.2% in 2024 and 3.5% in 2025 (fiscal-year basis) just as remittances surged. Whether these external flows raise the productive capacity of the economy, rather than only its spending, therefore remains an open and policy-relevant question.

Theory gives grounds for both optimism and doubt. FDI may raise growth by adding to the capital stock and by transferring technology and managerial practice, but these gains depend on the absorptive capacity and financial depth of the host economy (Borensztein et al. 1998; Alfaro et al. 2004). Remittances can relax household credit constraints, deepen the financial system and finance investment (Giuliano and Ruiz-Arranz 2009; Azizi 2020; Dash 2020), but because they are largely compensatory transfers they may also be consumed, reduce labor supply or appreciate the real exchange rate (Chami et al. 2005). The cross-country evidence reflects this ambiguity. Bird and

Choi (2020) find that FDI raises growth whereas remittances lower it, and a meta-analysis of 538 estimates from 95 studies finds that about 40% of remittance–growth estimates are positive, 40% are insignificant and 20% are negative, with an average effect that is positive but economically small once publication bias is corrected (Cazachevici et al. 2020).

Evidence for Bangladesh is equally divided. Siddique et al. (2012) and Wadood and Hossain (2015) find that remittances drive growth, and Golder et al. (2023) report asymmetric but growth-enhancing remittance effects; Sutradhar (2020) and Orchi and Ahmed (2025) find negative effects; Chowdhury et al. (2023) find no short- or long-run link; Milon et al. (2024) report positive long-run but negative short-run effects of foreign capital inflows; and Sarker (2024) finds FDI-led growth that weakens during crises. Three features of this literature help explain the disagreement. First, many studies regress output on FDI and remittances alone and leave out domestic investment and labor, although Cazachevici et al. (2020) show that omitting other sources of finance biases remittance estimates and that time-series studies report systematically larger effects. Second, most samples end before 2020 and therefore miss the COVID-19 shock, the balance-of-payments strain of 2022–2024 and the remittance surge of 2024–2025. Third, the bounds-testing procedure is often applied only in part: the t-test on the error-correction coefficient, which guards against spurious evidence of a levels relationship (Pesaran et al. 2001), and small-sample critical values (Narayan 2005) are seldom reported.

This study addresses these gaps in three ways. First, it uses actual World Development Indicators (WDI) data through 2025, released in October 2026 and measured in constant prices, so that no observation is extrapolated or forecast. Second, it estimates the effects of FDI and foreign remittances (FR) within an augmented production function that controls for gross fixed capital formation and labor, and compares the result with the parsimonious specification common in

earlier work, so that the consequences of omitting these controls can be seen directly. Third, it applies the full bounds-testing procedure, reporting both the F and t statistics against asymptotic and simulated finite-sample critical values, and complements it with diagnostic and stability tests, three alternative cointegrating estimators, recursive estimates and Toda–Yamamoto causality tests.

The study asks three questions. Do FDI and FR affect Bangladesh's real output in the long run once domestic investment and labor are accounted for? Do they affect output in the short run? And has their relationship with output been stable over time? The evidence answers all three in the negative. A stable long-run relationship exists, but it ties output to domestic fixed investment. In the main specification, neither FDI nor FR has a statistically significant effect in the long or the short run, and no alternative estimator or specification finds a positive effect of either inflow; the modestly positive remittance elasticity found in samples ending before 2016 disappears as more recent years are added. The rest of the paper is organized as follows. The next section reviews the literature and sets out the hypotheses. The Methods section describes the data and the estimation strategy, the Results section reports the findings, the Discussion section interprets them and sets out the study's limitations, and the final section concludes.

Literature review

Theoretical underpinnings

The analysis is grounded in the neoclassical growth model (Solow 1956) and its augmented empirical form (Mankiw et al. 1992), in which output depends on physical capital, labor and the efficiency with which they are combined. External inflows can affect output through two channels. In the *capital-accumulation channel*, FDI and remittances finance additions to the capital stock, either directly (foreign-owned plants, remittance-financed housing and small enterprises) or indirectly by easing financing constraints. In the *productivity channel*, they shift total factor

productivity, for example through technology spillovers from foreign firms or through the financial deepening that remittance flows can bring about. Writing the production function as $Y = A(FDI, FR)K^{\alpha}L^{\beta}$ makes the distinction explicit: conditional on capital (K) and labor (L), the coefficients on FDI and FR capture the productivity channel, while their contribution through capital accumulation works through K. For this reason, the empirical model controls for capital formation and labor and, separately, tests whether FDI and FR predict capital formation itself.

For FDI, endogenous-growth arguments emphasize spillovers of technology and know-how, but these appear only above thresholds of human capital and financial development (Borensztein et al. 1998; Alfaro et al. 2004) and depend on the linkages that foreign firms form with domestic suppliers and customers (Newman et al. 2020). The theoretical predictions for remittances are more divided. Altruistically motivated, compensatory transfers rise when recipient incomes fall and are mostly consumed, which can create moral hazard in labor supply and weaken their link to growth (Chami et al. 2005). They also smooth household consumption against shocks (Combes and Ebeke 2011) and reduce poverty (Adams and Page 2005). Where financial systems are shallow, remittances can substitute for missing credit and finance investment (Giuliano and Ruiz-Arranz 2009), while large inflows may appreciate the real exchange rate and erode the competitiveness of tradable sectors. A contrasting structuralist view, rooted in dependency theory, holds that reliance on external capital can entrench dependence on foreign markets and firms (Dos Santos 1970). Taken together, these perspectives make the sign and size of the effects an empirical question.

Foreign direct investment and economic growth

Early cross-country work found that FDI raises growth only where human capital (Borensztein et al. 1998) or local financial markets (Alfaro et al. 2004) are sufficiently developed. Panel causality

evidence suggests that the link runs more strongly from growth to FDI than the reverse (Choe 2003), studies that deal carefully with endogeneity often find no robust independent FDI effect (Carkovic and Levine 2005), and country-by-country time-series evidence for developing economies frequently fails to detect FDI-led growth (Herzer et al. 2008). Other work is more favorable. Iamsiraroj and Ulubaşoğlu (2015) conclude that the relationship is positive on average, Bird and Choi (2020) find a significant positive FDI effect in a broad panel, and Dinh et al. (2019) find that FDI supports growth in lower-middle-income countries in the long run but not in the short run. Single-country ARDL studies report similar heterogeneity (Mohamed et al. 2021; Rehman et al. 2021), and Das and Sethi (2020) find that FDI matters for growth in India but not in Sri Lanka. For Bangladesh, Sarker (2024) finds positive short- and long-run FDI effects over 1974–2020 that weaken during crisis episodes, while Milon et al. (2024) report that foreign capital inflows promote growth in the long run but not in the short run. A recent legal-economic assessment argues that Bangladesh's FDI regime has been oriented more towards investor protection than development outcomes, which limits the domestic linkages through which FDI could raise productivity (Nasrullah and Rahim 2025).

Remittances and economic growth

Cross-country studies of remittances and growth range from negative effects (Chami et al. 2005; Bird and Choi 2020) to positive effects that depend on financial development (Giuliano and Ruiz-Arranz 2009), with many finding no robust relationship (Barajas et al. 2009). Clemens and McKenzie (2018) give two reasons why remittance effects are hard to detect: recorded remittance growth often reflects changes in measurement rather than new flows, and the growth effect of plausible remittance increases is small relative to the noise in growth data. The meta-analysis of Cazachevici et al. (2020) finds a positive but economically small average effect after correcting

for publication bias, larger effects in Asia than in Africa, and systematically larger estimates in time-series studies and in studies that omit other external flows or ignore endogeneity.

In South Asia, Cooray (2012) and Islam (2022) report positive remittance effects in panel settings, Das and Sethi (2020) find that remittances support growth in both India and Sri Lanka, and Dash (2020) finds that remittances raise domestic investment across six South Asian economies. Similar ARDL designs elsewhere produce mixed results; for Kenya, Mutai et al. (2025) find that remittances raise growth in the short run and that remittances, FDI and imports are all significant determinants of growth in the long run.

Institutions and the policy environment

The growth payoff from both flows depends on the institutional and financial setting. Countries with sound governance, regulation and investment climates are better placed to benefit from FDI (Alfaro et al. 2004), and in Bangladesh regulatory and bureaucratic frictions have limited the development impact of foreign investment (Nasrullah and Rahim 2025). For remittances, the financial system matters in two ways. Remittances deepen financial intermediation: in a panel of 124 developing countries, a 10% rise in the remittance-to-GDP ratio raises credit to the private sector by about 1.7% (Azizi 2020). Household behavior matters as well. Survey evidence from Bangladesh shows that remittances encourage households to save through bank accounts, deposit pension schemes, insurance and cooperative societies (Salahuddin et al. 2022), yet most remittance income still finances household consumption rather than productive investment (Khondker and Majid 2025). Whether these inflows raise aggregate output therefore depends on how far they are converted into productive capital, which is why domestic investment is central to the empirical model below.

Empirical evidence for Bangladesh and the research gap

Table 1 summarizes the studies most relevant to Bangladesh. Their conclusions differ in sign as well as in magnitude. Studies using data that end in the 2000s or early 2010s tend to find that remittances drive growth (Siddique et al. 2012; Wadood and Hossain 2015), whereas more recent studies report mixed, negative or no effects (Sutradhar 2020; Chowdhury et al. 2023; Orchi and Ahmed 2025). Specifications also vary widely: some include domestic investment, trade and aid, while others regress output on remittances or FDI alone. None of the studies listed uses data beyond 2021, and few report the full set of bounds-test statistics. The present study fills these gaps by combining actual data through 2025, an augmented production-function specification and the complete bounds-testing procedure.

Table 1. Selected empirical studies on FDI, remittances and growth relevant to Bangladesh

Study	Coverage	Method	Main finding
Siddique et al. (2012)	Bangladesh, India, Sri Lanka; c. 25 years	VAR Granger causality	Remittance growth causes economic growth in Bangladesh (one-way)
Wadood and Hossain (2015)	Bangladesh, 1972–2013	Johansen cointegration, VECM	Long-run relationship; remittances cause growth
Das and Sethi (2020)	India and Sri Lanka, 1980–2016	VECM, Granger causality	FDI and remittances matter in India; aid and remittances in Sri Lanka
Sutradhar (2020)	Bangladesh, India, Pakistan, Sri Lanka	Country-level and joint regressions	Remittances reduce growth in Bangladesh, Pakistan and Sri Lanka; positive in India
Bird and Choi (2020)	Cross-country panel	Panel growth regressions	FDI raises growth; remittances lower it; aid mixed
Cazachevici et al. (2020)	95 studies, 538 estimates	Meta-analysis	Small positive average remittance effect after correcting for publication bias
Dash (2020)	Six South Asian countries, 1991–2017	Panel cointegration and causality	Remittances raise domestic investment (one-way causality)
Islam (2022)	South Asia, 1986–2019	Panel GLS, FMOLS; Dumitrescu–Hurlin	Remittances raise growth; causality runs from remittances to growth
Chowdhury et al. (2023)	Bangladesh, Sri Lanka, Vietnam, 1990–2019	Panel and country-level causality	Adverse panel effect; no short- or long-run link for Bangladesh
Golder et al. (2023)	Bangladesh, 1988–2020	Nonlinear ARDL	Asymmetric effects; both rises and falls in remittances and financial development associated with higher growth
Milon et al. (2024)	Bangladesh, 1976–2019	ARDL bounds test	Foreign capital inflows raise growth in the long run; short-run effects mostly negative
Sarker (2024)	Bangladesh, 1974–2020	Augmented ARDL; FMOLS, DOLS	FDI raises growth in short and long run; crises disrupt FDI-led growth
Orchi and Ahmed (2025)	Bangladesh, 1986–2021	ARDL and NARDL	Aid and exports raise growth; remittances weigh on it

Study	Coverage	Method	Main finding
This study	Bangladesh, 1990–2025	ARDL bounds (F and t, finite-sample), FMOLS, CCR, DOLS, Toda–Yamamoto	See Results

Source: Compiled by the authors.

Research hypotheses

Following the prevailing view in the policy debate and the theoretical channels described above, the study tests the following hypotheses:

- **H1:** Foreign direct investment has a positive and significant long-run effect on Bangladesh's real GDP.
- **H2:** Foreign remittances have a positive and significant long-run effect on Bangladesh's real GDP.
- **H3:** FDI and foreign remittances have positive and significant short-run effects on Bangladesh's real GDP.

Conceptual framework

Figure 1 presents the conceptual framework. FDI and FR are the variables of interest, and domestic investment and labor enter as controls derived from the production function. The dotted arrows indicate the capital-accumulation channel through which FDI and FR may affect output indirectly, which is examined with Toda–Yamamoto causality tests.

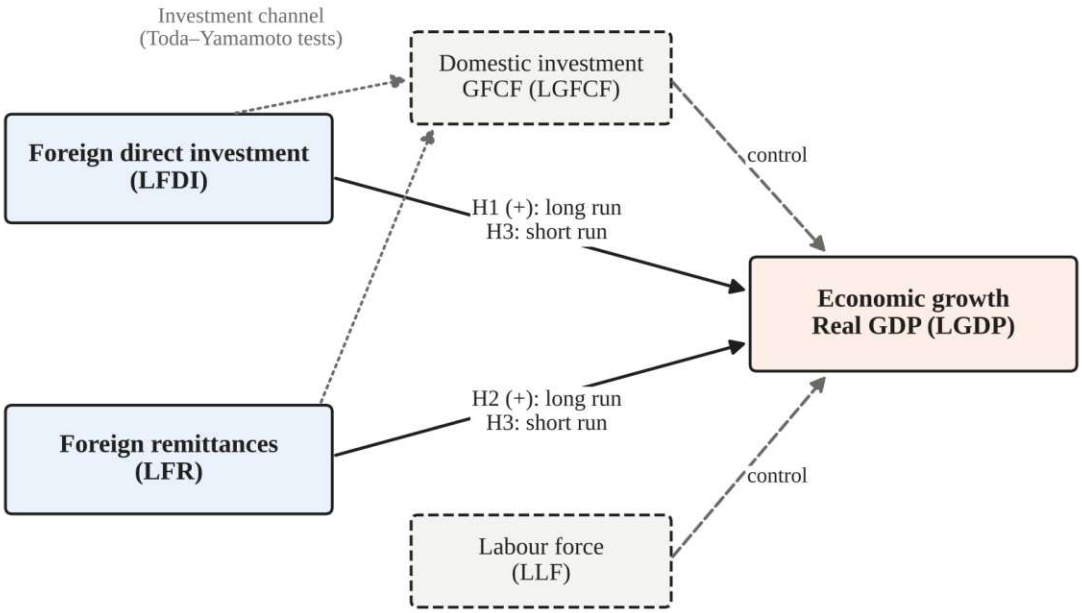


Figure 1. Conceptual framework. Solid arrows: hypothesized effects (H1–H3); dashed arrows: control variables; dotted arrows: indirect investment channel. Source: Developed by the authors.

Methods

Data and variables

The study uses annual data for Bangladesh from 1990 to 2025 (36 observations) drawn from the World Bank's World Development Indicators (World Bank 2026). The data were retrieved on 24 October 2026; the database had last been updated on 8 October 2026. Every observation is an actual published value, and none is forecast or interpolated. The sample starts in 1990 because the labor-force series begins in that year, which also coincides with the start of Bangladesh's trade and financial liberalization.

Table 2 defines the variables. Real GDP and gross fixed capital formation (GFCF) are taken in constant 2015 US dollars. FR is measured by personal remittances received, which comprise personal transfers and compensation of employees. Net FDI inflows and FR, which the WDI reports in current US dollars, are converted to constant 2015 US dollars with Bangladesh's implicit US-dollar GDP deflator (GDP in current US dollars divided by GDP in constant 2015 US dollars),

so that all monetary variables share the same real units. All variables enter in natural logarithms, so the coefficients are elasticities. Net FDI inflows were positive in every year of the sample, so the logarithm is defined throughout. GFCF serves as the proxy for capital, as is common in ARDL growth studies; the robustness checks in the Results section explain why a constructed capital stock is not a usable alternative here. Bangladesh's national accounts are compiled on a July–June fiscal-year basis, whereas the balance-of-payments flows are calendar-year totals. The resulting misalignment is at most six months and is partly absorbed by the dynamic specification.

Table 2. Variables, definitions and data sources

Variable	Definition	WDI code	Unit and transformation	Expected sign
LGDP	Gross domestic product (dependent variable)	NY.GDP.MKTP.KD	Constant 2015 US$; natural log	–
LFDI	Foreign direct investment, net inflows (BoP)	BX.KLT.DINV.CD.WD	Deflated to constant 2015 US$; natural log	+
LFR	Foreign remittances (personal remittances, received)	BX.TRF.PWKR.CD.DT	Deflated to constant 2015 US$; natural log	+
LGFCF	Gross fixed capital formation (proxy for capital)	NE.GDI.FTOT.KD	Constant 2015 US$; natural log	+
LLF	Labor force, total	SL.TLF.TOTL.IN	Persons; natural log	+
LTRADE	Trade, sum of exports and imports (robustness only)	NE.TRD.GNFS.ZS	% of GDP; natural log	±

Note: The deflator used for LFDI and LFR is NY.GDP.MKTP.CD divided by NY.GDP.MKTP.KD (2015 = 1). BoP = balance of payments. Source: World Bank (2026).

Model specification

Taking logarithms of the augmented production function set out in the literature review gives the long-run relationship to be estimated:

$$\ln Y_t = \beta_0 + \beta_1 \ln FDI_t + \beta_2 \ln FR_t + \beta_3 \ln K_t + \beta_4 \ln L_t + \varepsilon_t \quad (1)$$

where Y is real GDP, FDI and FR are real FDI and remittance inflows, K is proxied by real GFCF and L is the labor force. Two versions are estimated. **Model 1** is the parsimonious specification common in earlier studies; it sets $\beta_3 = \beta_4 = 0$ and relates output to FDI and FR only. **Model 2**, the

main specification, includes capital formation and labor. Comparing the two shows how much of any apparent FDI or remittance effect is attributable to omitted domestic inputs.

Unit root tests

Although the ARDL approach accommodates regressors that are I(0), I(1) or mutually cointegrated, it is invalid if any series is I(2). The order of integration of each variable is therefore established with the augmented Dickey–Fuller (ADF; Dickey and Fuller 1979) and Phillips–Perron (PP; Phillips and Perron 1988) tests, whose null hypothesis is a unit root, and the KPSS test (Kwiatkowski et al. 1992), whose null hypothesis is stationarity. Levels are tested with an intercept and trend, and first differences with an intercept. ADF lag lengths are chosen by the Akaike information criterion (AIC) with a maximum of four, and the KPSS bandwidth is selected automatically (Hobijn et al. 2004). Because structural breaks can bias conventional tests towards non-rejection, the Zivot and Andrews (1992) test, which allows one endogenously determined break, is applied as well.

ARDL bounds-testing approach

Cointegration is tested with the ARDL bounds approach of Pesaran and Shin (1999) and Pesaran et al. (2001). The approach suits the present setting because it is valid for any mixture of I(0) and I(1) regressors, performs well in small samples, separates short- and long-run effects within a single equation and, with an adequate lag structure, corrects for the endogeneity of the regressors (Pesaran and Shin 1999). System GMM, by contrast, is a panel estimator and is not applicable to a single-country time series, while FMOLS and DOLS estimate only the long-run relationship; both are therefore used here as robustness checks rather than as the main method. The ARDL(p, q_1, …, q_k) model in levels is

$$y_t = c + \Sigma_{i=1}^{p} \varphi_i \, y_{t-i} + \Sigma_{j=1}^{k} \Sigma_{l=0}^{q} \beta_{jl} \, x_{j,t-l} + u_t \qquad (2)$$

where y is ln Y, the x_j are the k regressors and $q = q_j$ is the lag order of regressor j. Equation (2) can be re-parameterized exactly in error-correction form:

$$\Delta y_t = c + \rho y_{t-1} + \Sigma_{j=1}^{k} \pi_j x_{j,t} + \Sigma_{i=1}^{p-1} \varphi'_i \Delta y_{t-i} + \Sigma_{j=1}^{k} \Sigma_{l=0}^{q-1} \psi_{jl} \Delta x_{j,t-l} + u_t \quad (3)$$

This is the parameterization used by the Stata *ardl* command with the *ec* option (Kripfganz and Schneider 2023). Regressors with $q_j = 0$ enter only through their levels, so π_j is also their impact (short-run) elasticity. The coefficient ρ is the speed of adjustment towards equilibrium, the long-run elasticities are $\theta_j = -\pi_j/\rho$ with delta-method standard errors, and the half-life of a deviation from equilibrium is $\ln(0.5)/\ln(1 + \rho)$ years.

The bounds test involves two statistics. The F-statistic tests H_0: $\rho = \pi_1 = \ldots = \pi_k = 0$, and the t-statistic tests H_0: $\rho = 0$. A levels relationship is accepted only if the F-statistic exceeds the upper, I(1), critical bound and the t-statistic lies below its I(1) bound; values beyond the I(0) bound indicate no levels relationship, and values in between are inconclusive. Reporting the t-test is essential, because a large F-statistic combined with $\rho \approx 0$ is the degenerate case in which the regressors are significant in levels but there is no error correction, and hence no cointegrating relationship (Pesaran et al. 2001). The statistics are compared with the asymptotic critical values of Pesaran et al. (2001) for Case III (unrestricted intercept, no trend). Because these asymptotic values can be unreliable with 35 observations, finite-sample critical values and p-values are also computed by Monte Carlo simulation, following the design of Narayan (2005) and in the spirit of Kripfganz and Schneider (2020). The exact regression in Equation (3), with the same sample size and lag structure, is estimated on 50,000 artificial samples in which y is a random walk unrelated to the regressors, and the regressors are either all I(0) (lower bound) or all I(1) (upper bound).

Lag orders are chosen by a grid search over $p \in \{1, 2\}$ and $q_j \in \{0, 1, 2\}$ for every regressor, with all regressors retained. Every candidate model is estimated on a common sample (1992–2025) so

that the information criteria are comparable. The AIC is the primary criterion and the Schwarz Bayesian criterion (SBC) serves as a check.

Diagnostic and stability tests

The selected models are checked for serial correlation (Breusch–Godfrey LM tests with one and two lags; Breusch 1978; Godfrey 1978), heteroskedasticity (Breusch and Pagan 1979; White 1980), autoregressive conditional heteroskedasticity (ARCH LM test; Engle 1982), non-normality (Jarque and Bera 1980) and functional-form misspecification (Ramsey 1969). Parameter stability is assessed with the CUSUM and CUSUM-of-squares tests of Brown et al. (1975), based on recursive residuals and 5% significance bounds; the CUSUM-of-squares bound is obtained by simulation for the exact number of recursive residuals.

Robustness and causality tests

Four sets of robustness checks are carried out. First, the long-run coefficients of Model 2 are re-estimated with three single-equation cointegration estimators that correct for endogeneity and serial correlation: fully modified OLS (FMOLS; Phillips and Hansen 1990), canonical cointegrating regression (CCR; Park 1992) and dynamic OLS (DOLS; Stock and Watson 1993) with one lead and one lag, each using a Bartlett kernel for the long-run variance. Second, the ARDL model is re-estimated with a COVID-19 impulse dummy for 2020, with trade openness added, without the labor-force variable, and on the pre-pandemic sample 1990–2019. Third, the stability of the FDI and FR elasticities over time is examined with recursive (expanding-window) estimates of the baseline model for samples ending in each year from 2010 to 2025.

Fourth, Granger causality among output, capital formation, FDI and FR is tested with the Toda and Yamamoto (1995) procedure, which remains valid whatever the variables' orders of integration or cointegration. A vector autoregression in levels is estimated with $k + d_{\max}$ lags, where

289 k is the lag order chosen by the SBC and $d_{max} = 1$ is the maximum order of integration, and Wald
290 tests are applied to the first k lags only. These tests also show whether FDI and FR predict capital
291 formation, that is, whether they operate through the capital-accumulation channel.

292 ***Software, replication files and use of AI tools***

293 All estimation was carried out in Python 3.13 using the statsmodels (Seabold and Perktold 2010;
294 version 0.15) and arch (Sheppard 2025; version 8.0) libraries. The ARDL model is estimated in
295 the same parameterization as the Stata *ardl* command, and a Stata do-file that reproduces the main
296 results is included with the replication code. The dataset is provided in Additional file 1 and the
297 replication code in Additional file 2. Figure 2 summarizes the analytical procedure.

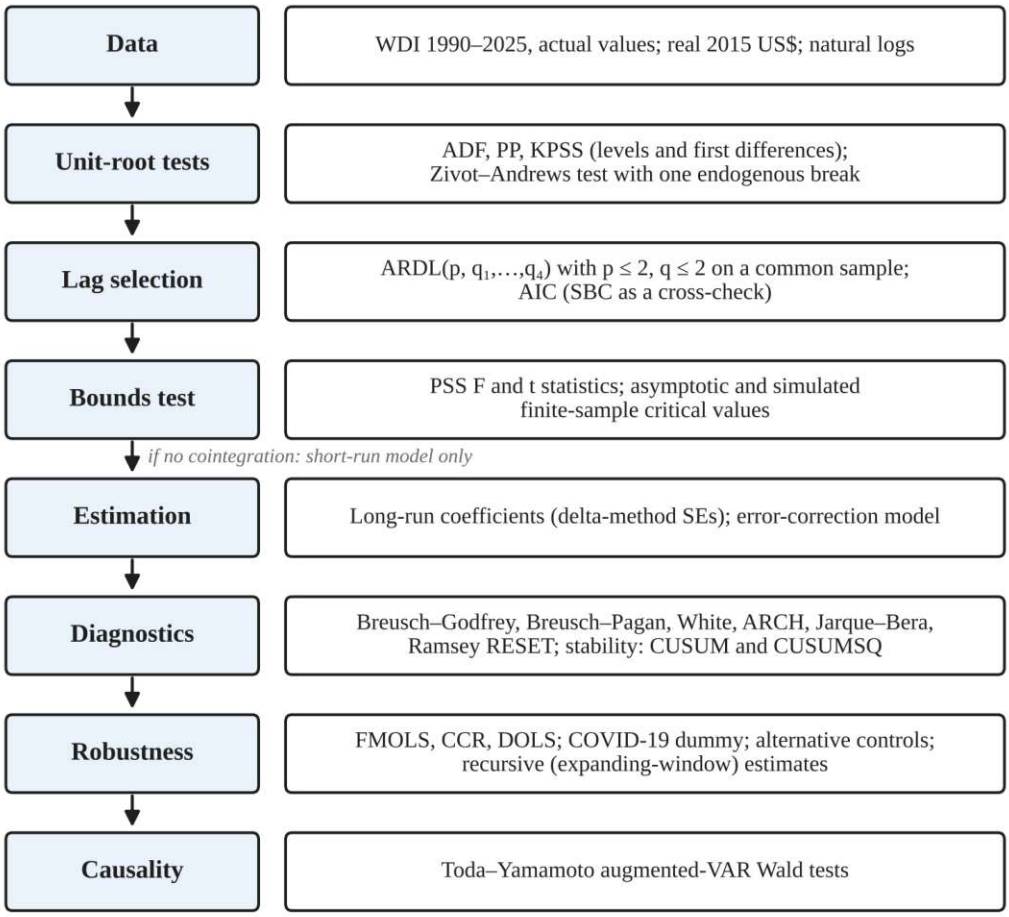


Figure 2. Analytical procedure. Source: Developed by the authors.

In line with the journal's policy on large language models, the use of an artificial intelligence (AI) assistant is documented here. Claude (Anthropic), an AI assistant, was used in preparing this version of the study: to retrieve the World Development Indicators series through the World Bank's public data interface; to write and run the Python code for the estimations, tables and figures; to revise and redraft the manuscript text, building on the authors' earlier version; and to check the reference list against Crossref records. The authors reviewed the code, the results and the text and take full responsibility for the content of the article. The AI assistant is not an author.

Results

Descriptive statistics and trends

Table 3 reports descriptive statistics, and Figure 3 plots the main series. Real GDP rose more than six-fold over the sample, from US$52.9 billion in 1990 to US$348.7 billion in 2025 (2015 prices). Real FR grew even faster, from US$1.3 billion to US$25.9 billion, but its path relative to GDP was far from steady: the remittance-to-GDP ratio climbed to 10.6% in 2012, fell to 4.6% by 2017 and recovered to 7.4% in 2025. FDI is small and volatile, peaking at 1.7% of GDP in 2013 and falling to 0.28% in 2024. Gross fixed capital formation rose from 16.5% of GDP in 1990 to a peak of 32.2% in 2019, before easing to 28.5% in 2025. Among the log-transformed series, only LFDI departs from normality (Jarque–Bera p = 0.013), reflecting the very small inflows of the early 1990s.

Table 3. Descriptive statistics, 1990–2025

Variable	Mean	Std. dev.	Minimum (year)	Maximum (year)
Panel A: Variables in natural units				
Real GDP (US$ billion, 2015 prices)	153.43	90.57	52.85 (1990)	348.69 (2025)
Real foreign remittances (US$ billion, 2015 prices)	9.44	6.57	1.30 (1990)	25.89 (2025)
Real FDI net inflows (US$ million, 2015 prices)	981.3	848.6	2.5 (1991)	2996.5 (2013)
Real gross fixed capital formation (US$ billion, 2015 prices)	41.76	31.86	7.83 (1990)	105.57 (2024)
Labour force (million persons)	56.13	11.61	36.31 (1990)	74.79 (2023)
Foreign remittances (% of GDP)	5.70	2.46	2.46 (1990)	10.59 (2012)
FDI net inflows (% of GDP)	0.580	0.483	0.004 (1991)	1.735 (2013)
Gross fixed capital formation (% of GDP)	25.66	4.74	16.46 (1990)	32.21 (2019)

Variable	Mean	Std. dev.	Minimum	Maximum	Skewness	Kurtosis	JB p-value
Panel B: Variables in natural logarithms (as used in estimation)							
LGDP	25.589	0.590	24.691	26.577	0.148	1.753	0.292
LFDI	19.718	2.080	14.714	21.821	−1.198	3.129	0.013
LFR	22.632	0.919	20.988	23.977	−0.429	1.670	0.153
LGFCF	24.139	0.839	22.781	25.383	−0.056	1.737	0.300
LLF	17.822	0.214	17.408	18.130	−0.289	2.087	0.416

Note: n = 36 for all variables. JB = Jarque–Bera test of normality. Source: Authors' calculations from World Bank (2026).

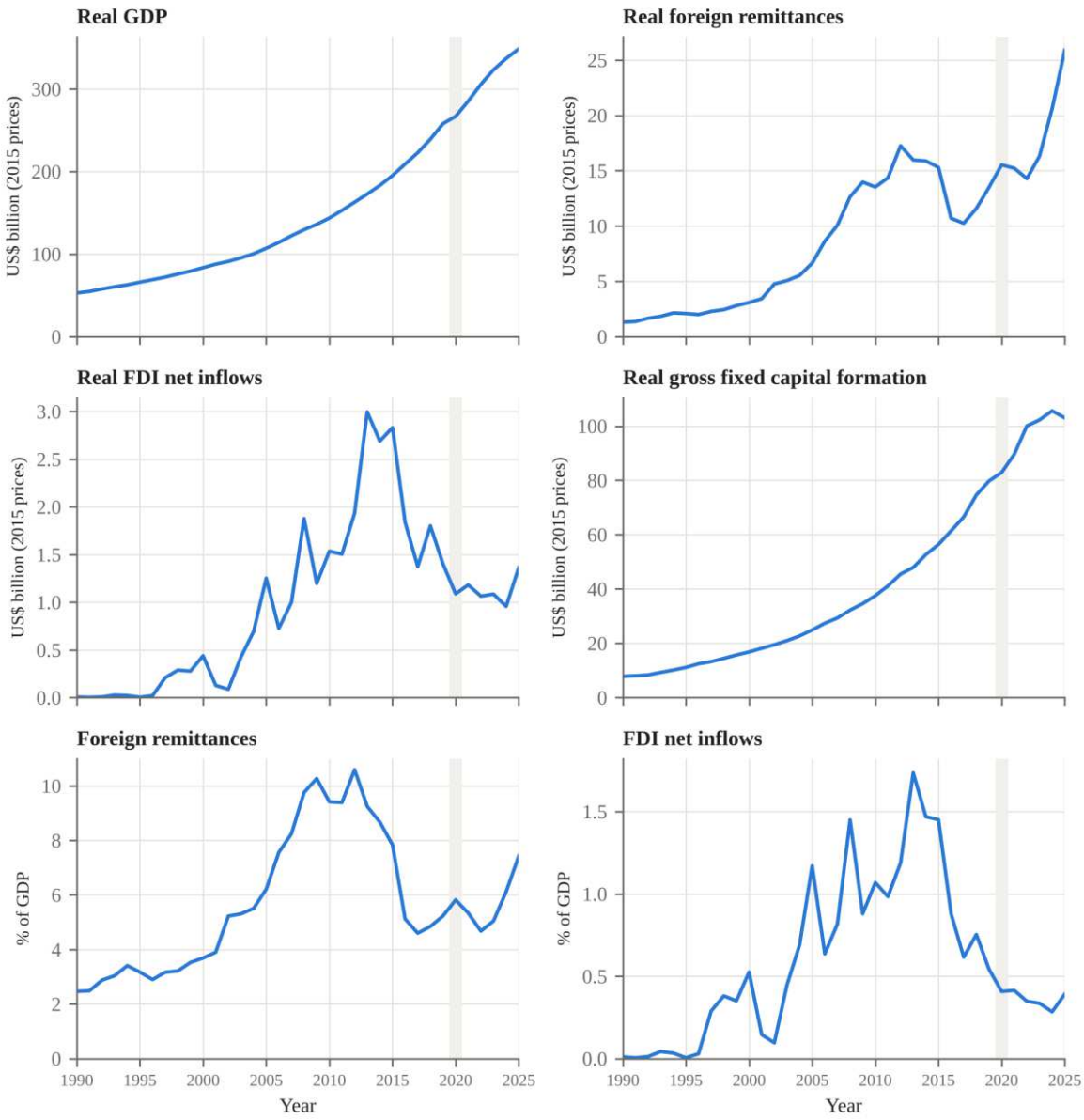


Figure 3. Real GDP, external inflows and fixed capital formation in Bangladesh, 1990–2025. Source: Authors' calculations from World Bank (2026).

Unit root and stationarity test results

Table 4 shows that all variables are integrated of order one. The ADF and PP tests fail to reject a unit root in levels for every series ($p \geq 0.35$) but reject it clearly in first differences ($p \leq 0.01$). The KPSS test confirms the picture: stationarity of levels is rejected at the 5% level for every variable, while stationarity of first differences is not rejected for LFDI, LFR, LGFCF and LTRADE. For

ΔLGDP and ΔLLF, the KPSS statistic exceeds the 5% but not the 1% critical value, a mild discrepancy that the ADF and PP tests do not share. The Zivot–Andrews tests in Table 5 show that allowing for a structural break does not change these conclusions: no series is break-stationary in levels, and all first differences reject the unit-root null at the 5% (LGDP, LFDI, LLF, LTRADE) or 10% (LFR, LGFCF) level. The estimated break in ΔLGDP and ΔLGFCF falls in 2020, consistent with the COVID-19 shock, which motivates the dummy-variable robustness check below. Since no series is I(2), the ARDL bounds approach is applicable.

Table 4. Unit root and stationarity tests

	ADF		PP		KPSS		
Variable	**Level**	**First difference**	**Level**	**First difference**	**Level**	**First difference**	**I(d)**
LGDP	−1.933 [0.638]	−3.677*** [0.004]	−2.440 [0.358]	−3.858*** [0.002]	0.204**	0.538**	I(1)
LFDI	−1.341 [0.877]	−6.484*** [<0.001]	−1.565 [0.806]	−5.605*** [<0.001]	0.238***	0.322	I(1)
LFR	−1.835 [0.688]	−3.813*** [0.003]	−1.496 [0.830]	−3.744*** [0.004]	0.164**	0.152	I(1)
LGFCF	0.589 [0.997]	−3.461*** [0.009]	0.523 [0.997]	−3.731*** [0.004]	0.152**	0.279	I(1)
LLF	−2.113 [0.539]	−3.568*** [0.006]	−1.697 [0.752]	−3.450*** [0.009]	0.191**	0.484**	I(1)
LTRADE	−1.476 [0.837]	−4.789*** [<0.001]	−1.239 [0.902]	−5.359*** [<0.001]	0.221***	0.332	I(1)

Note: Levels include an intercept and trend; first differences include an intercept. ADF and PP: H0 = unit root, MacKinnon p-values in brackets; ADF lags chosen by AIC (maximum 4). KPSS: H0 = stationarity; critical values are 0.119, 0.148 and 0.217 (levels) and 0.347, 0.461 and 0.743 (first differences) at 10%, 5% and 1%. ***, ** and * denote rejection of H0 at 1%, 5% and 10%.

Table 5. Zivot–Andrews unit root test with one structural break

Variable	**Level: statistic**	**Level: break year**	**First difference: statistic**	**First difference: break year**
LGDP	−3.366	2002	−5.273**	2020
LFDI	−3.593	1996	−7.297***	2001
LFR	−3.161	2005	−4.588*	2010
LGFCF	−3.662	2021	−4.597*	2020
LLF	−3.020	1999	−5.427***	2017
LTRADE	−4.363	2016	−5.977***	2014

Note: H0 = unit root. Levels allow a break in intercept and trend (5% critical value −5.07); first differences allow a break in the intercept (5% critical value −4.81). The break year is the first year of the new regime. Lags chosen by AIC (maximum 4); 15% trimming. ***, ** and * denote rejection at 1%, 5% and 10%.

Bounds tests for cointegration

Table 6 reports the bounds tests. For Model 1, the parsimonious specification, both criteria select ARDL(1, 0, 1). Its F-statistic of 4.16 lies between the asymptotic 5% bounds and below the finite-sample 5% lower bound (4.19), and its t-statistic of −0.31 lies well above even the 10% I(0) bound. The error-correction coefficient is essentially zero ($\rho = -0.0024$), which would imply that a deviation from the supposed equilibrium takes almost three centuries to halve. There is therefore no evidence of a long-run relationship among GDP, FDI and FR when domestic inputs are omitted, and the long-run coefficients of Model 1 are not identified; their standard errors are several times their point estimates (Table 7). Model 1 also fails the Breusch–Pagan, ARCH and CUSUM-of-squares tests (Table 9). Results based on this specification, including the large positive remittance elasticities reported in some earlier studies, should therefore be read with caution.

For Model 2, both the AIC and the SBC select ARDL(1, 0, 0, 0, 0). The F-statistic of 8.25 exceeds the 1% upper bound in both the asymptotic (5.06) and the finite-sample (6.28) tables, and the t-statistic of −4.22 lies below the 5% upper bound in both (−3.99 and −4.07). The simulated p-values against the I(1) bound are 0.002 for the F-test and 0.039 for the t-test. Real output, FDI, FR, capital formation and labor are therefore cointegrated at the 5% level according to both statistics.

Table 6. ARDL bounds tests for cointegration (Case III)

	Model 1: LGDP = f(LFDI, LFR); k = 2	**Model 2: LGDP = f(LFDI, LFR, LGFCF, LLF); k = 4**
Selected order (AIC)	ARDL(1, 0, 1)	ARDL(1, 0, 0, 0, 0)
Selected order (SBC)	ARDL(1, 0, 1)	ARDL(1, 0, 0, 0, 0)
Observations (sample)	35 (1991–2025)	35 (1991–2025)
F-statistic	4.164	8.246
t-statistic (on ρ)	−0.312	−4.216
Error-correction coefficient ρ	−0.0024	−0.1654
Asymptotic critical values, Pesaran et al. (2001), Case III: I(0), I(1)		
F, 10%	3.17, 4.14	2.45, 3.52
F, 5%	3.79, 4.85	2.86, 4.01
F, 1%	5.15, 6.36	3.74, 5.06

	Model 1: LGDP = f(LFDI, LFR); k = 2	Model 2: LGDP = f(LFDI, LFR, LGFCF, LLF); k = 4
t, 10%	−2.57, −3.21	−2.57, −3.66
t, 5%	−2.86, −3.53	−2.86, −3.99
t, 1%	−3.43, −4.10	−3.43, −4.60
Finite-sample critical values (50000 simulations, n = 35): I(0), I(1)		
F, 10%	3.37, 4.44	2.70, 3.89
F, 5%	4.19, 5.36	3.28, 4.61
F, 1%	6.10, 7.50	4.63, 6.28
t, 10%	−2.58, −3.22	−2.58, −3.67
t, 5%	−2.93, −3.61	−2.93, −4.07
t, 1%	−3.65, −4.36	−3.64, −4.88
Simulated p-value, F: I(0), I(1)	0.051, 0.122	0.000, 0.002
Simulated p-value, t: I(0), I(1)	0.899, 0.950	0.002, 0.039
Conclusion	No cointegration (t-test fails; $\rho \approx 0$)	Cointegration (F and t, 5%)

Note: Critical values and p-values are given as pairs for the lower, I(0), and upper, I(1), bounds. A levels relationship requires F above the I(1) bound and t below the I(1) bound. Finite-sample values are simulated for the exact regression and sample size used (see Methods). Estimation sample 1991–2025 after differencing.

Long-run and short-run estimates

Table 7 reports the long-run elasticities from Model 2. The long-run relationship is anchored by domestic investment: a 1% permanent increase in real gross fixed capital formation is associated with a 0.97% increase in real GDP ($p < 0.001$). The long-run elasticities of FDI (−0.003, $p = 0.781$) and FR (−0.039, $p = 0.252$) are small, negative and statistically insignificant. Their 95% confidence intervals, [−0.027, 0.021] for FDI and [−0.106, 0.029] for FR, are narrow enough to be informative: they rule out any long-run remittance elasticity larger than about 0.03 and any FDI elasticity larger than about 0.02. Hypotheses H1 and H2 are therefore not supported.

Two features of the control variables deserve comment. The capital elasticity of 0.97 is larger than conventional estimates of capital's share of income. This is to be expected, because GFCF is an investment flow rather than a capital stock and because investment also responds to output (an accelerator effect); the coefficient should be read as a reduced-form long-run co-movement rather than a structural capital share. The labor-force elasticity is negative and marginally significant (−0.77, $p = 0.055$). This counter-intuitive sign reflects the near-perfect collinearity between LLF

and LGFCF (correlation 0.991) and the modeled nature of the labor-force series. LLF is therefore retained as a theory-based control rather than interpreted structurally, and the robustness checks show that dropping it leaves the FDI and FR results unchanged.

Table 7. Long-run elasticities

	Model 2 (main specification)					Model 1
Variable	**Coefficient**	**Std. error**	**t-statistic**	**p-value**	**95% CI**	**Coefficient (SE)**
LFDI	−0.0033	0.0117	−0.280	0.781	[−0.027, 0.021]	1.374 (4.642)
LFR	−0.0385	0.0329	−1.170	0.252	[−0.106, 0.029]	0.269 (2.129)
LGFCF	0.9698***	0.0955	10.156	<0.001	[0.775, 1.165]	–
LLF	−0.7667*	0.3840	−1.997	0.055	[−1.552, 0.019]	–
Constant	17.0481***	4.6671	3.653	0.001	[7.503, 26.593]	15.891 (29.906)

Note: Long-run coefficients $\theta j = -\pi j/\rho$ with delta-method standard errors; p-values and confidence intervals use the t-distribution with 29 degrees of freedom. Model 1 coefficients are shown for comparison only; because Model 1 shows no cointegration (Table 6), they are not identified. ***, ** and * denote significance at 1%, 5% and 10%.

Table 8 reports the error-correction form of Model 2. The error-correction coefficient is −0.165 and highly significant (t = −4.22), so about 16.5% of any deviation of output from its long-run path is corrected within a year, and the half-life of a deviation is about 3.8 years. Because the selected model contains no lagged differences, the coefficients on the current levels of the regressors are their impact (short-run) elasticities. The short-run effects of FDI (−0.0005, p = 0.790) and FR (−0.0064, p = 0.240) are again small and insignificant, whereas capital formation has a significant positive impact effect (0.160, p < 0.001). Hypothesis H3 is therefore not supported either.

Table 8. Error-correction model (Model 2): dependent variable $\Delta LGDP_t$

Regressor	**Coefficient**	**Std. error**	**t-statistic**	**p-value**
$LGDP_{t-1}$ (error-correction coefficient, ρ)	−0.1654***	0.0392	−4.216	<0.001
$LFDI_t$	−0.0005	0.0020	−0.269	0.790
LFR_t	−0.0064	0.0053	−1.200	0.240
$LGFCF_t$	0.1604***	0.0354	4.528	<0.001
LLF_t	−0.1268**	0.0604	−2.098	0.045
Constant	2.8203***	0.8349	3.378	0.002

Model statistics	**Value**
R^2	0.587

Model statistics	Value
Adjusted R^2	0.516
S.E. of regression	0.0080
Durbin–Watson	2.069
Half-life of a deviation (years)	3.83
Observations	35

Note: Estimated by OLS on 1991–2025. The model is the exact error-correction re-parameterization of ARDL(1, 0, 0, 0, 0) (Equation 3); the coefficients on the regressors' current levels are impact (short-run) elasticities. ***, ** and * denote significance at 1%, 5% and 10%.

Diagnostic and stability test results

Table 9 shows that Model 2 passes every diagnostic test at the 5% level. There is no evidence of serial correlation (Breusch–Godfrey p = 0.777 and 0.849 for one and two lags), heteroskedasticity (Breusch–Pagan p = 0.646; White p = 0.377), ARCH effects (p = 0.204) or non-normality (Jarque–Bera p = 0.642), and the RESET test does not reject correct functional form at 5% (p = 0.067). The CUSUM and CUSUM-of-squares statistics remain within their 5% bounds throughout (Figure 4), indicating stable coefficients and error variance over 1997–2025, including the pandemic years. Model 1, in contrast, shows heteroskedasticity (Breusch–Pagan p = 0.013), ARCH effects (p = 0.005) and a CUSUM-of-squares path that crosses its bounds, further evidence that it is misspecified.

Table 9. Residual diagnostic and stability tests

Test	Null hypothesis	Model 2: statistic [p-value]	Model 1: statistic [p-value]
Breusch–Godfrey LM, 1 lag	No serial correlation	0.080 [0.777]	0.523 [0.470]
Breusch–Godfrey LM, 2 lags	No serial correlation	0.327 [0.849]	2.374 [0.305]
Breusch–Pagan	Homoskedasticity	3.349 [0.646]	12.647 [0.013]
White	Homoskedasticity	21.346 [0.377]	16.734 [0.271]
ARCH LM, 1 lag	No ARCH effects	1.612 [0.204]	7.839 [0.005]
Jarque–Bera	Normal residuals	0.888 [0.642]	0.077 [0.962]
Ramsey RESET (F)	Correct functional form	3.641 [0.067]	0.362 [0.552]
CUSUM (5%)	Parameter stability	Within bounds	Within bounds
CUSUM of squares (5%)	Variance stability	Within bounds	Crosses bounds

Note: LM = Lagrange multiplier. The CUSUM and CUSUM-of-squares tests use recursive residuals with 5% significance bounds (Brown et al. 1975).

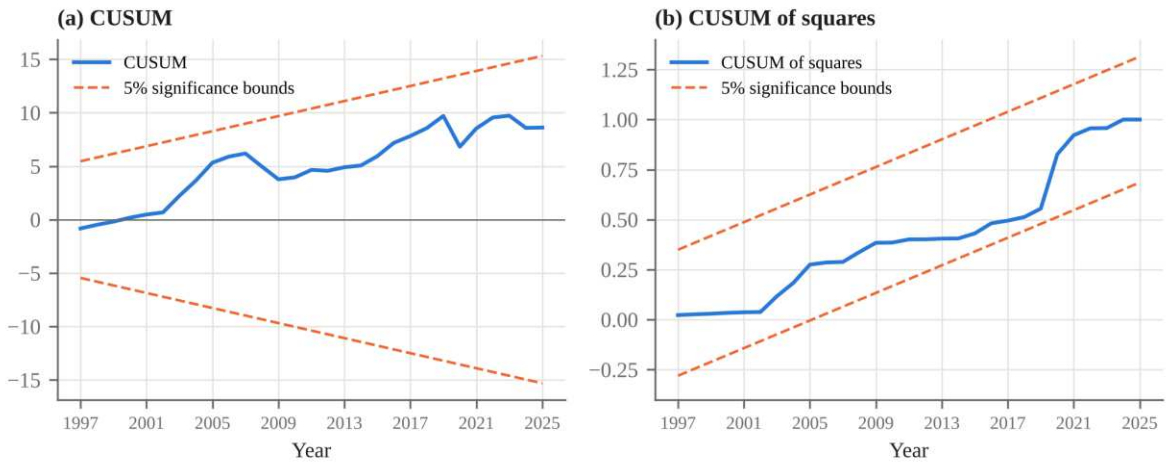


Figure 4. CUSUM and CUSUM-of-squares tests for Model 2. Source: Authors' calculations.

Robustness checks

Table 10 shows that the central finding, namely that FDI and FR have no positive long-run effect on output once domestic investment is accounted for, holds across specifications and estimators. Adding a COVID-19 dummy for 2020 strengthens the evidence of cointegration (F = 10.92, t = −4.57); the dummy itself is negative and significant (−0.0228, p = 0.004), but the FDI and FR elasticities barely change. Adding trade openness leaves the FDI and FR elasticities insignificant, and trade openness is itself insignificant (−0.053, p = 0.491). Dropping the labor force produces a capital elasticity of 0.83 and, again, insignificant FDI and FR elasticities. FMOLS, CCR and DOLS give capital elasticities between 0.85 and 0.91 and remittance elasticities that are close to zero and insignificant. These estimators find a small negative FDI elasticity of about −0.03 that is statistically significant. Its sign agrees with the ARDL point estimate, but its magnitude is economically negligible: even a doubling of FDI would lower output by only about 2% in the long run.

A natural objection is that GFCF is a flow rather than a stock. A capital stock built by the perpetual inventory method from real GFCF since 1980, with depreciation rates of 4–6%, behaves as an I(2) series over the sample (ADF p-values of 0.79–0.80 for its first difference and 0.002 or less for its

second difference), so it cannot enter the bounds framework. With that caveat, an ARDL model that uses the constructed stock leads to the same conclusion, with insignificant long-run elasticities for FDI (−0.0029, p = 0.542) and FR (−0.0001, p = 0.995).

Two caveats emerge from the robustness checks. When labor is dropped or trade openness is added, the t-statistic of the bounds test becomes inconclusive or fails, although the F-statistic remains large. On the pre-pandemic sample 1990–2019, the error-correction coefficient is essentially zero, so cointegration is not detected and the long-run coefficients are not identified. The evidence of a levels relationship therefore depends on the full 1990–2025 sample, whose post-2019 years provide the variation that identifies the equilibrium. In none of the eight specifications, however, is the FDI or FR elasticity positive and significant.

Table 10. Robustness of the long-run elasticities

Specification	LFDI	LFR	LGFCF	LLF	Bounds F / t	Cointegration (5%)
(1) Baseline, ARDL(1, 0, 0, 0, 0)	−0.003	−0.039	0.970***	−0.767*	8.25 / −4.22	Yes
(2) + COVID-19 dummy (2020), ARDL(1, 0, 0, 0, 0)	−0.005	−0.043	0.981***	−0.749**	10.92 / −4.57	Yes
(3) + trade openness (LTRADE), ARDL(1, 0, 0, 0, 0, 1)	0.002	−0.028	1.032***	−1.049***	11.55 / −3.81	Inconclusive (t)
(4) Without labour force, ARDL(1, 0, 0, 1)	−0.006	−0.035	0.826***	–	5.88 / −1.55	No (t)
(5) Sample 1990–2019, ARDL(1, 0, 0, 0, 0)	0.082	−0.304	3.976	−5.045	9.44 / −0.08	No (t)
(6) FMOLS	−0.032***	−0.005	0.864***	−0.361**	–	–
(7) CCR	−0.033***	−0.001	0.849***	−0.301*	–	–
(8) DOLS (1 lead, 1 lag)	−0.027***	−0.005	0.909***	−0.624***	–	–

Note: Specifications (1)–(5) are ARDL models estimated as described in the Methods section, with lag orders selected by AIC except (2), where the SBC-selected order is used; specification (3) uses a maximum lag of one for the regressors because of the larger number of parameters. Specifications (6)–(8) use a Bartlett kernel; DOLS uses robust standard errors. Cointegration is assessed with the asymptotic 5% bounds of Pesaran et al. (2001) for the relevant k; 'No (t)' means that the t-statistic lies above its I(0) bound and 'Inconclusive (t)' that it lies between the bounds. ***, ** and * denote significance at 1%, 5% and 10%.

Figure 5 presents recursive estimates of the long-run FDI and FR elasticities from the baseline model, estimated on samples that start in 1990 and end in each year from 2010 to 2025. The remittance elasticity is positive, at about 0.07–0.10, in samples ending between 2010 and 2015,

although it is imprecisely estimated (p-values between 0.09 and 0.24). It falls to 0.02 for the sample ending in 2016 and lies between −0.02 and −0.04, never significant, for every sample ending from 2020 onwards. The FDI elasticity stays close to zero throughout. The fading of the remittance effect coincides with the sharp fall in the remittance-to-GDP ratio between 2012 and 2017, a period in which GDP growth remained above 6% a year. Because the t-test does not confirm cointegration in samples ending before 2020, the early estimates should be read as indicative; even so, the pattern helps reconcile the positive remittance effects found with data ending in the 2000s with the null or negative effects reported in more recent studies.

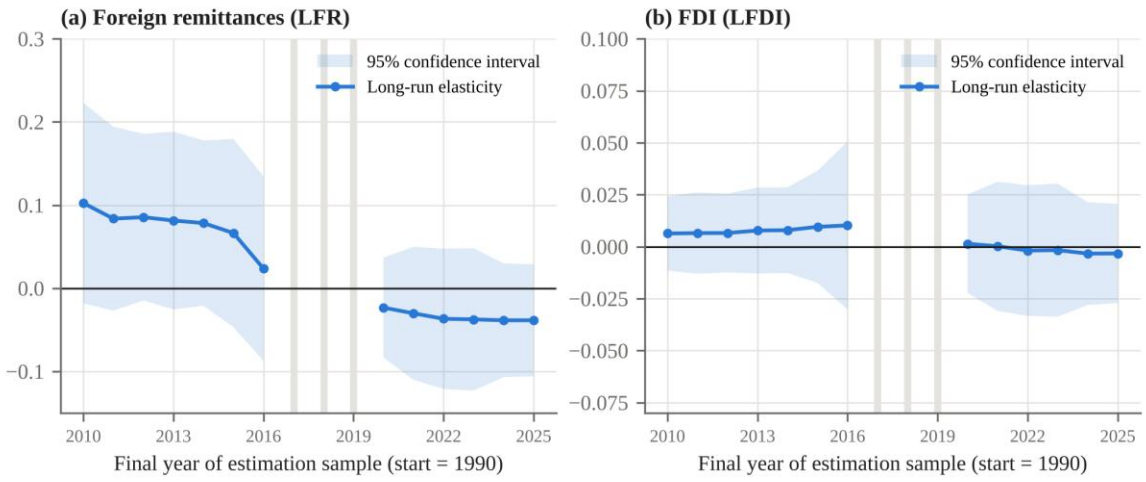


Figure 5. Recursive long-run elasticities of FR and FDI. Estimates from the baseline ARDL(1, 0, 0, 0, 0) model on samples starting in 1990 and ending in each year from 2010 to 2025, with 95% confidence intervals. Source: Authors' calculations.

Toda–Yamamoto causality tests

The SBC selects one lag for the levels VAR in LGDP, LGFCF, LFDI and LFR, so the Toda–Yamamoto tests are based on a VAR(2) with Wald tests on the first lag (34 observations). Table 11 shows that none of the twelve non-causality hypotheses is rejected at the 10% level. In particular, neither FDI ($p = 0.993$) nor FR ($p = 0.250$) Granger-causes output, and neither FDI ($p = 0.874$) nor FR ($p = 0.159$) Granger-causes fixed capital formation. At annual frequency, there is

therefore no evidence that FDI or FR operate even indirectly, through the capital-accumulation channel. With 34 observations these tests have limited power, so the absence of causality should be read as an absence of evidence rather than proof of no effect.

Table 11. Toda–Yamamoto Granger non-causality tests

Null hypothesis	Wald χ^2	df	p-value	Decision (10%)
LFDI does not Granger-cause LGDP	0.000	1	0.993	Not rejected
LFR does not Granger-cause LGDP	1.322	1	0.250	Not rejected
LGFCF does not Granger-cause LGDP	0.717	1	0.397	Not rejected
LFDI does not Granger-cause LGFCF	0.025	1	0.874	Not rejected
LFR does not Granger-cause LGFCF	1.983	1	0.159	Not rejected
LGDP does not Granger-cause LGFCF	0.977	1	0.323	Not rejected
LGDP does not Granger-cause LFDI	0.713	1	0.398	Not rejected
LGDP does not Granger-cause LFR	0.606	1	0.436	Not rejected
LGFCF does not Granger-cause LFDI	1.585	1	0.208	Not rejected
LGFCF does not Granger-cause LFR	0.058	1	0.810	Not rejected
LFR does not Granger-cause LFDI	0.333	1	0.564	Not rejected
LFDI does not Granger-cause LFR	0.007	1	0.932	Not rejected

Note: Levels VAR with k = 1 (selected by SBC; HQ also selects 1, AIC selects 2) plus dmax = 1 additional lag; Wald tests on the first k lags only (Toda and Yamamoto 1995).

Hypothesis tests

Table 12 summarizes the outcome of the hypothesis tests. None of the three hypotheses is supported.

Table 12. Summary of hypothesis tests

Hypothesis	Evidence	Decision
H1: FDI has a positive long-run effect on real GDP	$\theta = -0.003$ ($p = 0.781$); 95% CI [−0.027, 0.021]; FMOLS, CCR and DOLS: small negative	Not supported
H2: FR has a positive long-run effect on real GDP	$\theta = -0.039$ ($p = 0.252$); 95% CI [−0.106, 0.029]; insignificant in all robustness checks	Not supported
H3: FDI and FR have positive short-run effects	Impact elasticities −0.0005 ($p = 0.790$) and −0.0064 ($p = 0.240$); no Granger causality	Not supported

Note: Based on Model 2 (Tables 7, 8, 10 and 11).

Discussion

The central result is that, once domestic investment and labor are accounted for, neither FDI nor FR has a detectable positive effect on Bangladesh's real output, in the long run or the short run,

while domestic fixed investment anchors the long-run growth path. This contrasts with the widespread view that remittances have driven Bangladesh's growth and with studies that report large positive remittance elasticities (Wadood and Hossain 2015; Golder et al. 2023). It agrees, however, with the null or negative findings of Sutradhar (2020), Chowdhury et al. (2023) and Orchi and Ahmed (2025), and with the meta-analytic conclusion that remittance effects shrink once other sources of finance are controlled for (Cazachevici et al. 2020). The comparison of Models 1 and 2 shows why specification matters: without domestic inputs there is no cointegrating relationship at all, so any long-run remittance elasticity derived from such a model reflects the common upward trend of the series rather than a genuine equilibrium relationship.

Why FDI does not register

The most direct explanation for the absence of an FDI effect is scale. Net FDI inflows averaged 0.58% of GDP over 1990–2025 and only about 0.36% over 2020–2025, a small fraction of domestic fixed investment, which was close to 30% of GDP. Inflows of this size are unlikely to move aggregate output measurably, whatever their effect on individual firms or sectors. The absorptive-capacity literature also predicts weak aggregate effects where human capital and financial depth are limited (Borensztein et al. 1998; Alfaro et al. 2004) and where foreign firms are poorly linked to domestic ones (Newman et al. 2020), and Bangladesh's FDI regime has been criticized for prioritizing investor protection over development linkages (Nasrullah and Rahim 2025). The small negative elasticity found by FMOLS, CCR and DOLS is consistent with the weak aggregate FDI–growth evidence for developing countries (Carkovic and Levine 2005; Herzer et al. 2008), and the absence of causality from FDI to growth matches panel evidence that the relationship runs mainly the other way (Choe 2003); the estimated effect is, however, too small to be economically meaningful. The difference from Sarker (2024), who finds positive FDI effects,

is plausibly explained by specification and sample: the present model conditions on domestic investment and labor and includes the post-2020 years in which FDI fell as a share of GDP.

Why foreign remittances do not register

Three explanations, which are not mutually exclusive, are consistent with the remittance results. First, remittances in Bangladesh predominantly finance household consumption rather than productive investment (Khondker and Majid 2025). Such spending supports aggregate demand and household welfare, and smooths consumption against shocks (Combes and Ebeke 2011), but in a production-function framework that conditions on capital it would not be expected to raise output independently. Even where remittances are saved, as household survey evidence for Bangladesh suggests (Salahuddin et al. 2022), much of that saving takes forms such as deposits and insurance whose conversion into fixed investment depends on financial intermediation. The Toda–Yamamoto tests find no evidence that FR raises fixed capital formation, which is consistent with the compensatory view of remittances (Chami et al. 2005).

Second, recorded remittances are an imperfect measure of the underlying flows. A substantial share of remittances has historically moved through informal *hundi* channels, and the surge of 2024–2025 partly reflects a shift into formal banking channels following exchange-rate adjustments, higher incentives and enforcement against informal operators (Khondker and Majid 2025). When recorded growth reflects changes in recording rather than new resources, the measured remittance series contains noise unrelated to output, which biases estimated effects towards zero (Clemens and McKenzie 2018).

Third, the relationship has changed over time. The recursive estimates show a positive, if imprecise, remittance elasticity in samples ending before 2016, which disappears as later years are added. Between 2012 and 2017 the remittance-to-GDP ratio more than halved while GDP growth

remained above 6% a year, decoupling the two series. This helps explain why studies based on data from the 1970s to the early 2010s found remittance-led growth (Siddique et al. 2012; Wadood and Hossain 2015), while studies with more recent data do not.

Theoretical implications

The findings are consistent with the augmented neoclassical view in which capital accumulation is the proximate driver of output (Solow 1956; Mankiw et al. 1992) and in which external inflows matter only insofar as they finance additions to capital or raise productivity. For Bangladesh, neither channel is visible in the aggregate data. The results therefore support neither the optimistic narrative of remittance-led growth nor the pessimistic dependency perspective (Dos Santos 1970): there is no evidence that external inflows have harmed long-run output, but none that they have raised it independently of domestic investment. Methodologically, the study shows that conclusions about externally financed growth can hinge on whether domestic inputs are controlled for and on whether the error-correction t-test is reported; studies that omit either step risk mistaking common trends for causal effects.

Practical implications

Three implications follow. First, policy towards remittances should shift from maximizing recorded volumes to converting inflows into productive investment. Incentives for formal-channel transfers remain worthwhile because remittances support foreign-exchange reserves and macroeconomic stability, but the growth dividend depends on channeling a larger share into savings and investment, for example through remittance-linked savings and investment products, credit for small enterprises run by migrant households, and skills training that raises migrants' earnings (Khondker and Majid 2025).

Second, FDI policy should focus on scale and linkages rather than on headline commitments. With net inflows below half a percent of GDP, FDI cannot be a material driver of aggregate growth; raising it requires a more predictable regulatory regime oriented towards development outcomes (Nasrullah and Rahim 2025), together with supplier-development programs that connect foreign firms with domestic producers. Third, and most importantly, domestic investment is the variable that anchors Bangladesh's long-run output. The fall in the fixed-investment ratio from 32.2% of GDP in 2019 to 28.5% in 2025 is therefore a greater concern for long-run growth than fluctuations in either external inflow, and measures that improve financial intermediation, infrastructure and the investment climate deserve priority.

Limitations and future research

The study has several limitations. First, with 35 usable annual observations, the bounds and causality tests have limited power, and the evidence of cointegration depends on the full sample, as it is not detected on the 1990–2019 subsample. Second, GFCF is an investment flow used as a proxy for the capital stock; a constructed stock behaves as an I(2) series over this sample, and official capital-stock estimates would allow a more structural interpretation of the capital elasticity. Third, the labor-force series consists of modeled estimates and is almost perfectly collinear with capital formation, which explains its implausible sign. Fourth, the analysis uses aggregate national data, combines fiscal-year national accounts with calendar-year balance-of-payments flows, and relies on recorded remittances that omit informal transfers. Finally, the model is linear and symmetric. Future research could use quarterly data to increase power, allow for asymmetric or threshold effects with nonlinear ARDL models, exploit sectoral FDI data to test for spillovers within industries, and use household survey data to trace how remittances are used.

Conclusions

This study re-examined whether FDI and foreign remittances drive economic growth in Bangladesh, using actual World Development Indicators data for 1990–2025 in constant prices, an augmented production-function specification and the complete ARDL bounds-testing procedure. A parsimonious model relating GDP to FDI and FR alone shows no valid long-run relationship. When domestic fixed investment and labor are included, a stable long-run relationship emerges, with about 17% of any disequilibrium corrected each year, but it is anchored by domestic investment; the long- and short-run elasticities of FDI and FR are small and statistically insignificant. No alternative estimator, specification or causality test reveals a positive effect of either inflow, and recursive estimates show that the modest positive remittance effect visible in earlier samples has disappeared. The hypotheses that FDI and FR raise Bangladesh's output in the long run (H1, H2) and in the short run (H3) are therefore not supported. For policymakers, the lesson is that the volume of external inflows is not, by itself, a reliable engine of growth; what matters is how much of it is converted into productive domestic investment.

Supplementary information

Additional file 1. Format: Microsoft Excel workbook (.xlsx). Title: Dataset. Description: the World Development Indicators extract for Bangladesh, 1970–2025, the estimation sample for 1990–2025 and the variable definitions.

Additional file 2. Format: ZIP archive (.zip) containing Python scripts (.py), a Stata do-file (.do), a README file (.md), the data files (.csv) and the estimation outputs (.json, .csv, .png). Title: Replication code. Description: code that builds the dataset from the World Development Indicators extract and reproduces every table and figure in the article, with instructions and the outputs it produces.

Abbreviations

ADF: Augmented Dickey–Fuller

AI: Artificial intelligence

AIC: Akaike information criterion

ARCH: Autoregressive conditional heteroskedasticity

ARDL: Autoregressive distributed lag

BoP: Balance of payments

CCR: Canonical cointegrating regression

CI: Confidence interval

COVID-19: Coronavirus disease 2019

CUSUM: Cumulative sum of recursive residuals

DOLS: Dynamic ordinary least squares

FDI: Foreign direct investment

FMOLS: Fully modified ordinary least squares

FR: Foreign remittances

GDP: Gross domestic product

GFCF: Gross fixed capital formation

GLS: Generalized least squares

GMM: Generalized method of moments

HQ: Hannan–Quinn information criterion

JB: Jarque–Bera

KPSS: Kwiatkowski–Phillips–Schmidt–Shin

LFDI: Natural log of real FDI net inflows

LFR: Natural log of real foreign remittances

LGDP: Natural log of real GDP

LGFCF: Natural log of real gross fixed capital formation

LLF: Natural log of the labor force

LM: Lagrange multiplier

LTRADE: Natural log of trade openness (% of GDP)

NARDL: Nonlinear autoregressive distributed lag

OLS: Ordinary least squares

PP: Phillips–Perron

RESET: Regression equation specification error test

SBC: Schwarz Bayesian criterion

SE: Standard error

VAR: Vector autoregression

VECM: Vector error-correction model

WDI: World Development Indicators

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- Additionalfile1Dataset.xlsx
- Additionalfile2Replicationcode.zip