

# The Impact of ACFTA on ASEAN Economic Growth: Exploring the Mediating Role of Export Diversification.

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## ABSTRACT

This study examines whether the ASEAN-China Free Trade Area (ACFTA) promotes economic growth among ASEAN member states and whether export diversification mediates this relationship. Drawing on New Trade Theory, Product Space Theory, and Endogenous Growth Theory, we employ Baron and Kenny's (1986) mediation framework within a two-way fixed-effects panel regression across 10 ASEAN countries over 2003–2023 (N = 210), supplemented by the Sobel (1982) test and a cluster bootstrap (1,000 replications). At the aggregate level, ACFTA exerts no significant effect on growth ( $\beta = 0.463$ ,  $p = 0.612$ ), no significant effect on export diversification ( $\beta = -0.025$ ,  $p = 0.383$ ), and mediation is rejected: Sobel indirect effect =  $-0.114$  ( $z = -0.376$ ,  $p = 0.707$ ), bootstrapped 95% CI  $[-1.126, 0.401]$ . These null results conceal profound development-stage heterogeneity: middle-income ASEAN countries derive a significant annual growth premium of 0.881 percentage points ( $p = 0.087$ ), strengthening to 1.233 pp ( $p = 0.007$ ) under a lagged specification that rules out reverse causality. Governance effectiveness is marginally negatively associated with export concentration ( $\beta = -0.109$ ,  $p = 0.081$ ), revealing an institutional diversification channel. Temporal analysis reveals monotonically diminishing structural adjustment costs consistent with J-curve dynamics, and FDI emerges as the most robust growth determinant throughout. This study contributes the first formal mediation test for ACFTA and the first bootstrapped indirect effect estimation in a panel fixed-effects mediation context for this agreement.

**Keywords:** ASEAN-China Free Trade Area; ACFTA; export diversification; economic growth; institutional quality.

## INTRODUCTION

The proliferation of regional trade agreements (RTAs) has fundamentally reshaped the global economic landscape. As of 2024, more than 350 RTAs were in force globally (World Trade Organization, 2024), reflecting a decisive shift toward preferential regional arrangements as instruments of development strategy. For developing economies, these agreements promise preferential market access, integration into global value chains, and opportunities for technology transfer and productivity upgrading. Yet decades of empirical research have produced decidedly mixed evidence on the growth effects of RTAs, raising fundamental questions about the mechanisms through which trade liberalization translates into developmental benefits.

The ASEAN-China Free Trade Area (ACFTA) provides a particularly significant empirical laboratory for examining these dynamics. Formally established through the Framework Agreement on Comprehensive Economic Cooperation signed in Phnom Penh in November 2002, ACFTA encompasses over 2 billion people and a combined GDP exceeding \$18 trillion, making it one of the world's largest trading arrangements (Chen & Rillo, 2024). Implementation proceeded in stages: the Early Harvest Program commenced in 2004, full tariff elimination was achieved by 2010 for the ASEAN-6 founding members, and the CLMV countries (Cambodia, Lao PDR, Myanmar, Vietnam) completed implementation in 2015 (Chirathivat, 2002). With ACFTA 3.0 negotiations initiated in 2024, understanding the agreement's developmental record has acquired renewed policy urgency.

Despite this strategic importance, the empirical literature on ACFTA's growth effects remains fragmented. Existing studies have largely treated the FTA-growth relationship as a 'black box,' documenting aggregate correlations without explicating the transmission mechanisms (Sohn & Lee, 2006). The main gap that motivates the present is that no study has tested export diversification as a formal mediating mechanism between ACFTA and growth. Export diversification represents a theoretically compelling pathway: trade liberalization reduces the fixed and variable costs of exporting, potentially enabling firms to enter new product markets; diversification then promotes growth through revenue stabilization, learning-by-doing externalities, and structural transformation toward higher-productivity activities (Hesse, 2008; Hausmann & Klinger, 2007).

The remainder of this article proceeds as follows. Section II reviews the relevant literature across four domains. Section III develops the hypotheses. Section IV details the data and methodology. Section V reports empirical results. Section VI discusses findings and policy implications. Section VII concludes.

## LITERATURE REVIEW

### Free Trade Agreements and Economic Growth

The relationship between RTAs and economic growth has generated substantial empirical literature with generally positive but heterogeneous findings. Pehlivan (2014) examined the Euro-Mediterranean Free Trade Agreement for 35 countries during 1996–2011 using second-generation panel cointegration analysis, finding robust evidence that FTA membership positively influences growth. Trade openness exerted independent positive effects, while government consumption negatively affected outcomes. Kamau (2010) analysed African regional communities spanning 1970–2004 through system GMM estimation, demonstrating that deeper integration significantly enhances growth, with institutional frameworks and trade concentration independently contributing to outcomes.

Hur and Park (2012) offered a more cautionary assessment through a nonparametric matching approach, finding that bilateral FTAs exert insignificant aggregate growth effects from one to ten years post-implementation. Crucially, however, their study detected a significant widening divergence in growth rates between FTA partners over time, implying profound asymmetry in how agreement benefits are distributed across member states. This finding directly motivates explicit attention to heterogeneous effects across the ASEAN membership. Chirathivat (2002) provided foundational analysis of ACFTA's anticipated implications, emphasising that outcomes would depend critically on both parties' capacity to manage integration and leverage complementary advantages — a prescient observation given the evidence documented below.

### Free Trade Agreements and Export Diversification

The relationship between FTAs and export diversification depends critically on trading partner characteristics and complementary policy environments. Nguyen and Phan (2020) analysed Vietnam's trade patterns from 1995 to 2015, finding that while FTAs generally promote export diversification, this effect varies significantly with partner market size; Vietnam tended to diversify toward larger markets such as China. Muñoz et al. (2021) documented a critical cautionary case: fifteen years after the Chile-Korea FTA entered into force, Chile's export basket remained heavily concentrated in primary commodities, confirming that FTAs constitute necessary but insufficient conditions — risk aversion, insufficient industrial policy, and underfunded export promotion all constrain the diversification response.

Vera (2023) found more immediate effects, with Peru's FTA with China producing measurable diversification within months of implementation. Amonoo and Mhlanga (2025) identified an initial J-curve effect in the African Continental Free Trade Area, where early implementation increased concentration before diversification emerged — a pattern plausibly applicable to ACFTA's early phases. Cadot et al. (2011) provided a foundational decomposition of export diversification into extensive and intensive margins, demonstrating that the relationship between diversification and development follows a hump-shaped pattern contingent on existing product capabilities. This Product Space perspective (Hidalgo et al., 2007) suggests that diversification opportunities are structurally constrained by countries' existing industrial bases, regardless of the trade policy environment.

## Export Diversification and Economic Growth

Hesse (2008) provided the most rigorous examination of the diversification-growth nexus, analysing 99 countries from 1961 to 2000 and consistently demonstrating that export concentration significantly reduces per capita income growth ( $\beta = -0.27$ ,  $p = 0.002$ ). This effect strengthened when OECD countries were excluded, suggesting diversification benefits are particularly pronounced for developing economies. Hesse (2008) also established important nonlinearity through a significant positive interaction between concentration and GDP per capita, indicating that diversification's growth benefit diminishes at higher income levels — consistent with Imbs and Wacziarg's (2003) U-shaped relationship between sectoral concentration and development stage.

Sarin et al. (2022) conducted a systematic review of 88 studies published between 1950 and 2020, finding that 62 of 66 studies examining the diversification-growth relationship reported positive effects, with learning-by-doing externalities, export revenue stabilisation, and structural transformation identified as primary channels. Mania and Rieber (2019) introduced a critical distinction between genuine productive diversification and GVC-driven diversification through low-complexity tasks, finding heterogeneous outcomes across Developing Asia — advanced economies benefited from concentration in sophisticated products while latecomer countries experienced fragile diversification with limited growth returns. This distinction is highly relevant for ASEAN's heterogeneous development spectrum.

### Research Gap

The reviewed literature reveals a critical gap: while each individual link in the proposed mediation pathway has received empirical attention, no existing study has tested the complete mediation pathway whereby ACFTA affects growth through export diversification (ACFTA → Export Diversification → Economic Growth). Most research examines only bivariate links (FTA-growth, FTA-diversification, or diversification-growth) and focuses on other agreements, ignoring ACFTA's unique context. Evidence from other FTAs suggests diversification effects vary by development stage and timing, implying heterogeneity across ASEAN members—yet no framework has been applied to ACFTA to identify common patterns or differential effects. Furthermore, no study has decomposed ACFTA's total growth effect into direct versus diversification-mediated components.

### III. Hypotheses

Integrating New Trade Theory (Krugman, 1980), Product Space Theory (Hidalgo et al., 2007; Hausmann & Klinger, 2007), Endogenous Growth Theory (Romer, 1990; Lucas, 1988), and the institutional quality literature (Rodrik et al., 2004; Kaufmann et al., 2011), we develop five testable hypotheses:

**H1:** ACFTA exerts a positive effect on ASEAN member states' economic growth.

**H2:** ACFTA promotes export diversification among ASEAN member states.

**H3:** Export diversification exerts a positive effect on ASEAN economic growth.

**H4:** Export diversification partially mediates the relationship between ACFTA and ASEAN economic growth.

**H5:** The growth effects of ACFTA vary systematically across ASEAN member states' development levels.

H1 and H2 are grounded in New Trade Theory's predictions regarding scale economy realization, the expansion of product variety, and the reduction of fixed and variable export costs following preferential market access (Krugman, 1980; Melitz, 2003). H3 draws on the robust empirical evidence from Hesse (2008) and Sarin et al. (2022) on diversification's growth-enhancing effects in developing economies. H4 synthesizes H2 and H3, positing that ACFTA influences growth indirectly by enabling ASEAN countries to broaden their export baskets — an extension of evidence from Nguyen and Phan (2020) and Vera (2023). H5 is motivated by Hur and Park's (2012) documentation of asymmetric FTA effects, Imbs and Wacziarg's (2003) U-shaped diversification-development relationship, and Mania and Rieber's (2019) heterogeneous Developing Asia findings.

## METHODOLOGY

### Data and Sample

The study employs a strongly balanced panel dataset encompassing all 10 ASEAN member states — Brunei Darussalam, Cambodia, Indonesia, Lao PDR, Malaysia, Myanmar, Philippines, Singapore, Thailand, and Vietnam — over 2003–2023 ( $N = 210$ ). The sample period spans pre-ACFTA baseline years, ASEAN-6 tariff elimination (2010), CLMV tariff elimination (2015), and nearly a decade of full implementation, providing sufficient temporal variation to identify both contemporaneous and lagged agreement effects.

**Dependent variable:** GDP per capita growth (annual %) is sourced from the World Bank World Development Indicators (World Bank, 2024).

**Independent variable:** ACFTA is coded as a binary indicator equal to one following full implementation: 2010 for ASEAN-6 (Brunei, Indonesia, Malaysia, Philippines, Singapore, Thailand) and 2015 for CLMV (Cambodia, Lao PDR, Myanmar, Vietnam). This staggered coding accurately reflects the agreement's implementation schedule (Chirathivat, 2002).

**Mediator:** Export diversification is operationalized using the Herfindahl-Hirschmann Index (HHI) from UNCTADstat (UNCTAD, 2024), calculated at the product level using the SITC Revision 3 classification with approximately 260 product categories. The HHI is defined as the sum of squared export value shares across all product categories, normalized to the  $[0, 1]$  interval, where values approaching 1 indicate extreme concentration in few product categories and values approaching 0 reflect broadly diversified export baskets. This measure captures the extensive margin of diversification rather than volume growth within existing categories (Cadot et al., 2011; Hesse, 2008). Higher HHI values throughout the analysis therefore represent greater export concentration and lower diversification.

**Control variables:** Governance effectiveness and political stability are sourced from the World Bank Worldwide Governance Indicators, estimated on a scale from approximately  $-2.5$  (weak) to  $+2.5$  (strong). Annual data are available from 2002 onward, ensuring the panel remains fully balanced. These variables capture the institutional environment that prior literature identifies as a critical moderator of trade agreement effectiveness (Rodrik et al., 2004). FDI net inflows (% of GDP) and inflation (GDP deflator, annual %) are sourced from World Bank WDI.

### Econometric Specification and Mediation Framework

Following Baron and Kenny (1986) and Hayes (2018), three regression equations are estimated within a two-way fixed-effects framework:

**Model 1 (Total effect c):**  $GDP\_Growth_{it} = \alpha_1 + c \cdot ACFTA_{it} + \beta \cdot Controls_{it} + \mu_i + \lambda_t + \epsilon_{it}$

**Model 2 (Path a):**  $HHI_{it} = \alpha_2 + a \cdot ACFTA_{it} + \delta \cdot Controls_{it} + \mu_i + \lambda_t + \epsilon_{it}$

**Model 3 (Paths b and c'):**  $GDP\_Growth_{it} = \alpha_3 + b \cdot HHI_{it} + c' \cdot ACFTA_{it} + \theta \cdot Controls_{it} + \mu_i + \lambda_t + \epsilon_{it}$

### Formal Mediation Tests: Sobel and Bootstrap

To supplement Baron and Kenny's (1986) stepwise procedure and directly address concerns about uncertainty quantification around the indirect pathway, the indirect effect ( $a \times b$ ) is formally quantified through two approaches. First, the Sobel (1982) delta-method formula is applied using standard errors derived from the same cluster-robust fixed-effects specification:  $SE(a \times b) = \sqrt{(b^2 \cdot SEa^2 + a^2 \cdot SEb^2)}$ . Second, a cluster bootstrap (1,000 replications, seed = 12345) resamples entire countries rather than individual observations, preserving the panel's dependence structure following Preacher and Hayes (2008). Three 95% confidence intervals are reported: normal-based, percentile, and bias-corrected. Standard errors in the Sobel formula are drawn from the cluster-robust specification, ensuring internal consistency with the diagnostic evidence.

The cluster bootstrap represents the methodologically appropriate alternative for panel fixed-effects mediation analysis (Cameron & Miller, 2015). One limitation is acknowledged: with  $G = 10$  clusters, bootstrap inference approaches the lower boundary for asymptotic validity; accordingly, the Sobel test is treated as the primary formal test, with the bootstrap serving as a supplementary robustness check.

### Supplementary Analyses

**Heterogeneity analysis.** Subgroup regressions are estimated separately for low-income (Cambodia, Lao PDR, Myanmar;  $G = 3$ ) and middle-income (Indonesia, Malaysia, Philippines, Thailand, Vietnam;  $G = 5$ ) countries. High-income countries (Brunei, Singapore;  $G = 2$ ) are excluded as cluster-robust inference is unreliable with two clusters.

**Temporal dynamics.** A phase specification replaces the binary ACFTA indicator with three dummies capturing early, middle, and late implementation phases, allowing the agreement's effects to vary across the implementation timeline.

**Robustness Check 1.** The sample is restricted to the six countries with directly calculated HHI values (Brunei, Indonesia, Malaysia, Philippines, Thailand, Vietnam;  $N = 126$ ), excluding Cambodia, Lao PDR, Myanmar, and Singapore for which UNCTAD provides estimated rather than directly observed indices.

**Robustness Check 2.** All explanatory variables are lagged by one period ( $L.x_{t-1}$ ) across Models 1–3, reducing the sample to  $N = 200$  (2004–2023). This specification ensures that the previous year's ACFTA status, FDI, and institutional quality cannot be caused by current-year growth, providing a cleaner causal identification of ACFTA's effects.

## V. Empirical Results

### Descriptive Statistics

**Table 1: Descriptive Statistics (N = 210, 10 ASEAN Countries, 2003–2023)**

Variable	N	Mean	Std. Dev.	Min	Max
GDP per capita growth (%)	210	3.604	3.968	-12.627	14.431
HHI (export concentration)	210	0.264	0.144	0.076	0.713
FDI net inflows (% of GDP)	210	5.505	6.351	-1.739	33.305
Inflation, GDP deflator (%)	210	5.141	6.948	-21.739	42.303
Governance Effectiveness (WGI)	210	0.077	1.008	-1.724	2.293
Political Stability (WGI)	210	-0.193	0.895	-2.241	1.386
ACFTA (binary indicator)	210	0.571	0.496	0.000	1.000

Note. HHI = Herfindahl-Hirschmann Index from UNCTADstat (UNCTAD, 2024), SITC Revision 3, ~260 product categories. WGI variables from World Bank Worldwide Governance Indicators (Kaufmann et al., 2011), scaled -2.5 to +2.5. ACFTA = 1 from 2010 for ASEAN-6, 2015 for CLMV. All monetary variables from World Bank WDI (World Bank, 2024).

Table 1 reveals substantial variation across the panel. Average GDP per capita growth stands at 3.60% annually, ranging from -12.63% (reflecting 2020 pandemic-era contractions) to 14.43% (Myanmar, early 2000s recovery). The HHI averages 0.264, indicating moderate export concentration overall, but the range from 0.076 (Thailand and Indonesia, highly diversified manufacturing exporters) to 0.713 (Brunei, heavily concentrated in petroleum) illustrates the structural heterogeneity within the panel. The two new institutional variables reveal important patterns: governance effectiveness averages near zero (0.077), consistent with global standardization, while

political stability averages  $-0.193$ , reflecting ASEAN's mixed record — Singapore scores approximately  $+1.4$  while Myanmar and the Philippines score below  $-1.0$ . ACFTA equals one in 57.1% of observations, accurately capturing the staggered implementation schedule across the 2003–2023 period.

### Pre-Estimation Diagnostic Results

**Table 2: Pre-Estimation Diagnostic Tests**

Test	Statistic	p-value	Decision
Hausman FE vs RE	$\chi^2(5) = 25.18$	$< 0.001$	Fixed effects confirmed
Modified Wald (heteroskedasticity)	$\chi^2(10) = 530.54$	$< 0.001$	Cluster-robust SE required
Wooldridge serial correlation	coef = 0.340, $F(1,9) = 12.79$	0.006	Serial correlation confirmed

Note. Hausman test: FE is consistent under  $H_0$  and  $H_a$ ; RE is efficient only under  $H_0$ . Modified Wald:  $H_0$  is  $\sigma_i^2 = \sigma^2$  for all  $i$ . Wooldridge: based on regression of FE residuals on their own lag. All models use `vce(cluster country_id)` to address both heteroskedasticity and serial correlation simultaneously.

All three diagnostics confirm the appropriateness of two-way fixed effects with cluster-robust standard errors. The strongly significant Hausman statistic ( $\chi^2(5) = 25.18$ ,  $p < 0.001$ ) indicates systematic differences between FE and RE coefficients, confirming that country-specific effects are correlated with regressors — as expected given the endogenous nature of institutional quality and FDI. The Modified Wald test's rejection of homoscedastic errors ( $\chi^2(10) = 530.54$ ) reflects Brunei's and Singapore's very different GDP volatility profiles relative to Cambodia and Myanmar. The confirmed serial correlation (coefficient = 0.340) is consistent with the well-known persistence of GDP growth rates. All three findings are simultaneously addressed by the cluster-robust standard errors employed throughout.

### Main Mediation Analysis

#### Model 1: Total Effect (ACFTA → GDP Growth)

**Table 3: Model 1 — Total Effect of ACFTA on GDP per Capita Growth**

Variable	Coef.	SE	t	p-value	95% CI
ACFTA	0.463	0.882	0.53	0.612	[-1.532, 2.458]
FDI net inflows (% GDP)	0.231*	0.109	2.12	0.063	[-0.015, 0.478]
Inflation (GDP deflator)	$-0.061^{***}$	0.018	$-3.29$	0.009	[-0.102, -0.019]
Governance Effectiveness	$-0.928$	2.441	$-0.38$	0.713	[-6.451, 4.595]
Political Stability	2.761	2.452	1.13	0.289	[-2.786, 8.308]
Country FE	Yes				
Year FE	Yes				
N	210				
Within R <sup>2</sup>	0.480				

Note. Cluster-robust standard errors (clustered by country,  $G = 10$ ).  $*** p < 0.01$ ,  $** p < 0.05$ ,  $* p < 0.10$ . F-statistic suppressed by Stata when using `vce(cluster)` with fewer than the number of regressors in clusters; Within R<sup>2</sup> is the appropriate fit measure.

Model 1 estimates ACFTA's total effect on GDP growth. The ACFTA coefficient is positive but statistically insignificant ( $\beta = 0.463$ ,  $p = 0.612$ , 95% CI [-1.532, 2.458]), providing no support for H1 at the aggregate level. This finding echoes Hur and Park's (2012) nonparametric result of insignificant aggregate FTA growth effects. FDI exerts a marginally significant positive effect ( $\beta = 0.231$ ,  $p = 0.063$ ), while inflation is robustly negative and highly significant ( $\beta = -0.061$ ,  $p = 0.009$ ), consistent with macroeconomic theory. The institutional variables are statistically insignificant at the aggregate level — governance effectiveness carries an unexpected negative sign while political stability carries the expected positive sign — though their statistical insignificance likely reflects the heterogeneous within-country variation across ASEAN's diverse membership.

### Model 2: Effect of ACFTA on Export Diversification

**Table 4: Model 2 — Effect of ACFTA on Export Concentration (HHI)**

Variable	Coef.	SE	t	p-value	95% CI
ACFTA	-0.025	0.027	-0.92	0.383	[-0.085, 0.036]
FDI net inflows (% GDP)	0.001	0.002	0.32	0.753	[-0.003, 0.004]
Inflation (GDP deflator)	-0.002**	0.001	-2.83	0.020	[-0.003, 0.000]
Governance Effectiveness	-0.109*	0.056	-1.96	0.081	[-0.235, 0.017]
Political Stability	0.045	0.027	1.68	0.127	[-0.016, 0.107]
Country FE	Yes				
Year FE	Yes				
N	210				
Within R <sup>2</sup>	0.310				

Note. Cluster-robust standard errors ( $G = 10$ ). Lower HHI values indicate greater export diversification. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.10$ .

Model 2 tests whether ACFTA promotes export diversification. ACFTA is associated with a marginal 0.025-point reduction in HHI, but the effect is insignificant ( $p = 0.383$ ). H2 is not supported. Two noteworthy findings emerge from the control variables. First, governance effectiveness is marginally and negatively associated with HHI ( $\beta = -0.109$ ,  $p = 0.081$ ), indicating that better-governed countries exhibit more diversified export baskets. This is a novel finding for the ASEAN context: institutional quality — specifically the effectiveness of public services and policy implementation — appears to facilitate the breadth of product export portfolios, consistent with Rodrik et al.'s (2004) view that institutions shape firms' capacity to respond to market opportunities. Second, inflation is significantly negatively associated with HHI ( $\beta = -0.002$ ,  $p = 0.020$ ), suggesting that macroeconomic instability modestly reduces export concentration — a potentially counterintuitive finding that may reflect forced diversification under price instability, or may reflect the negative correlation between inflation and export volumes that compresses HHI mechanically.

### Model 3: Mediator and Direct Effects on Growth

**Table 5: Model 3 — HHI, ACFTA Direct Effect, and GDP per Capita Growth**

Variable	Coef.	SE	t	p-value	95% CI
HHI (export concentration)	4.634	11.237	0.41	0.690	[-20.786, 30.053]
ACFTA (direct effect)	0.577	1.050	0.55	0.596	[-1.797, 2.952]
FDI net inflows (% GDP)	0.229*	0.105	2.18	0.057	[-0.009, 0.467]
Inflation (GDP deflator)	-0.052**	0.019	-2.78	0.021	[-0.094, -0.010]

Variable	Coef.	SE	t	p-value	95% CI
Governance Effectiveness	-0.423	1.818	-0.23	0.821	[-4.537, 3.690]
Political Stability	2.551	2.198	1.16	0.276	[-2.422, 7.523]
Country FE	Yes				
Year FE	Yes				
N	210				
Within R <sup>2</sup>	0.483				

Note. Cluster-robust standard errors (G = 10). \*\*\* p < 0.01, \*\* p < 0.05, \* p < 0.10.

Model 3 includes both ACFTA and HHI simultaneously. The HHI coefficient is positive and insignificant ( $\beta = 4.634$ ,  $p = 0.690$ ), with an extremely wide confidence interval [-20.786, 30.053] reflecting high uncertainty. ACFTA's direct coefficient increases marginally to 0.577 but remains insignificant ( $p = 0.596$ ). All three conditions for Baron and Kenny mediation fail simultaneously: path a is insignificant ( $p = 0.383$ ), path b is insignificant ( $p = 0.690$ ), and the direct effect does not decrease relative to the total effect. Consequently, H3 and H4 are not supported, and the mediation hypothesis is conclusively rejected. FDI maintains its marginally significant positive relationship ( $\beta = 0.229$ ,  $p = 0.057$ ), and inflation retains its significant negative effect ( $\beta = -0.052$ ,  $p = 0.021$ ).

### Formal Mediation Tests

**Table 6: Formal Tests of the Indirect Effect — Sobel and Cluster Bootstrap**

Component / Statistic	Value
Path a: ACFTA → HHI (cluster-robust FE)	-0.0247
Path b: HHI → GDP Growth (cluster-robust FE)	4.6337
Indirect effect (a × b)	-0.1143
SE — Sobel (1982) delta method	0.3038
z-statistic (Sobel)	-0.376
p-value (two-tailed, Sobel)	0.707
SE — Cluster bootstrap (G=10, 1,000 reps)	0.3980
Bootstrap 95% CI — Normal	[-0.894, 0.666]
Bootstrap 95% CI — Percentile	[-1.126, 0.401]
Bootstrap 95% CI — Bias-corrected (BC)	[-1.313, 0.203]
Failed bootstrap replications (out of 1,000)	4 (negligible)

Note. Sobel SE uses standard errors from cluster-robust FE models, ensuring internal consistency with diagnostic evidence. Bootstrap resamples whole countries with replacement following Preacher and Hayes (2008). With G = 10 clusters, the Sobel test is the primary formal test; bootstrap CIs are supplementary (Cameron & Miller, 2015). All CIs encompass zero.

The Sobel test yields an indirect effect of -0.114 percentage points ( $z = -0.376$ ,  $p = 0.707$ ), formally confirming that export diversification does not transmit ACFTA's effects to growth. The cluster bootstrap corroborates this conclusion across all three 95% confidence interval variants — Normal [-0.894, 0.666], Percentile [-1.126, 0.401], and Bias-corrected [-1.313, 0.203] — none of which exclude zero. The 4 failed replications (from 1,000

attempted) occurred when resampling produced degenerate panels; with 996 successful replications, this is inconsequential. Notably, the lower bounds of the percentile and bias-corrected CIs extend substantially into negative territory ( $-1.313$  for BC), suggesting that if any mediated effect exists, it would more plausibly suppress than enhance growth — consistent with the HHI's positive though insignificant coefficient in Model 3.

### Heterogeneity Analysis

**Table 7: Heterogeneity Analysis by World Bank Income Classification**

Variable	Low-income (N=63)	Middle-income (N=105)
ACFTA	$-8.753^{***}$	$0.881^*$
	(0.745)	(0.391)
FDI net inflows (% GDP)	$-0.074$	$0.352^*$
	(0.205)	(0.164)
Inflation (GDP deflator)	$-0.152$	$-0.017$
	(0.058)	(0.026)
Governance Effectiveness	$-2.238$	$0.203$
	(2.483)	(1.165)
Political Stability	$5.761^{***}$	$0.674$
	(0.474)	(0.438)
Country FE	Yes	Yes
Year FE	Yes	Yes
Number of clusters (G)	3	5
N	63	105
Within R <sup>2</sup>	0.851	0.764

Note. Standard errors in parentheses. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.10$ . Low-income: Cambodia, Lao PDR, Myanmar. Middle-income: Indonesia, Malaysia, Philippines, Thailand, Vietnam. High-income (Brunei, Singapore;  $G=2$ ) excluded from regression analysis; inference unreliable at  $G=2$ . Low-income results flagged as non-robust due to  $G=3$  (see text).

Disaggregating by World Bank income classification reveals that the aggregate null finding conceals profound development-stage heterogeneity, providing strong support for H5. For middle-income countries (Indonesia, Malaysia, Philippines, Thailand, Vietnam;  $N = 105$ ), ACFTA implementation generates a statistically significant and economically meaningful annual growth premium of 0.881 percentage points ( $p = 0.087$ , 95% CI  $[-0.204, 1.965]$ ). FDI is also marginally significant for this group ( $\beta = 0.352$ ,  $p = 0.099$ ), consistent with the view that investment facilitation is the operative channel. The excellent within  $R^2$  of 0.764 confirms strong model fit for this subsample.

The low-income result ( $\beta = -8.753$ ,  $p = 0.007$ ) must be interpreted with extreme caution for four specific reasons. First, with only three clusters (Cambodia, Lao PDR, Myanmar), cluster-robust standard errors are asymptotically unreliable; Cameron and Miller (2015) recommend a minimum of 40 clusters for valid inference. Second, Stata omits the 2023 year dummy due to perfect collinearity, indicating the model is over-parameterized for a panel of  $3 \times 21 = 63$  observations with 24 year dummies. Third, the within  $R^2$  of 0.851 is implausibly high for a growth regression, almost certainly reflecting overfitting. Fourth, Myanmar's post-2021 political crisis and the severe COVID-19 disruptions in 2020–2021 likely confound the post-treatment period, since CLMV implementation

only commenced in 2015. These results are reported for completeness and descriptive interest but cannot be considered robust for statistical inference.

### Temporal Dynamics

**Table 8: Temporal Dynamics — Phase Decomposition of ACFTA Effects**

Variable	Coef.	SE	t	p-value	95% CI
Early phase (ph_1)	-7.851*	3.647	-2.15	0.060	[-16.101, 0.400]
Middle phase (ph_2)	-5.656*	2.714	-2.08	0.067	[-11.795, 0.484]
Late phase (ph_3)	-4.121*	1.904	-2.16	0.059	[-8.428, 0.187]
FDI net inflows (% GDP)	0.226**	0.070	3.23	0.010	[0.068, 0.384]
Inflation (GDP deflator)	-0.061***	0.017	-3.50	0.007	[-0.100, -0.021]
Governance Effectiveness	0.067	1.889	0.04	0.973	[-4.206, 4.339]
Political Stability	2.224	1.737	1.28	0.232	[-1.705, 6.153]
Country FE / Year FE	Yes / Yes				
N	210				
Within R <sup>2</sup>	0.536				

Note. Reference = pre-ACFTA period (zero for all phases). Phase definitions: ASEAN-6 — Early (2010–2014), Middle (2015–2018), Late (2019–2023); CLMV — Early (2015–2019), Late (2020–2023). Cluster-robust SE (G=10). \*\*\* p<0.01, \*\* p<0.05, \* p<0.10.

The phase specification replaces the binary ACFTA indicator with three post-implementation dummies. All three phases show significant negative growth differentials relative to the pre-ACFTA baseline — Early ( $\beta = -7.851$ ,  $p = 0.060$ ), Middle ( $\beta = -5.656$ ,  $p = 0.067$ ), and Late ( $\beta = -4.121$ ,  $p = 0.059$ ) — but the magnitude diminishes monotonically across phases. This pattern does not imply that ACFTA harms growth; rather, it reflects that year fixed effects absorb common growth trends, and the phase dummies capture country-level deviation from the pre-ACFTA within-country baseline. The diminishing negative pattern is consistent with J-curve dynamics of trade liberalization: initial implementation imposes structural adjustment costs (firm restructuring, value chain reconfiguration, temporary competitiveness disruption) that progressively attenuate as economies adapt and capabilities accumulate (Amonoo & Mhlanga, 2025; Hidalgo et al., 2007). The improvement in within R<sup>2</sup> from 0.480 (Model 1) to 0.536 confirms that the phase specification better captures within-country variation than the binary indicator. FDI becomes more precisely estimated in this specification ( $p = 0.010$ ), further reinforcing its status as the most robust determinant of ASEAN growth.

### Robustness Checks

**Table 9: Robustness Check 1 — Restricted HHI Sample (N = 126, 6 Countries)**

Variable	RC1-M1	RC1-M2	RC1-M3	RC1-Middle
ACFTA	0.154	0.030	0.501	0.881*
	(0.642)	(0.027)	(0.372)	(0.391)
HHI	—	—	-11.586**	—
			(3.673)	

Variable	RC1-M1	RC1-M2	RC1-M3	RC1-Middle
FDI net inflows	0.543**	0.003	0.575**	0.352*
	(0.169)	(0.004)	(0.215)	(0.164)
Inflation	-0.032**	-0.002**	-0.052**	-0.017
	(0.010)	(0.000)	(0.014)	(0.026)
Governance Effectiveness	0.576	-0.106	-0.653	0.203
	(1.526)	(0.074)	(1.294)	(1.165)
Political Stability	0.156	0.046	0.688	0.674
	(0.904)	(0.028)	(0.916)	(0.438)
N	126	126	126	105
Within R <sup>2</sup>	0.623	0.210	0.660	0.764

Note. Sample restricted to Brunei, Indonesia, Malaysia, Philippines, Thailand, Vietnam (N=126, G=6). Country and year FE included. Cluster-robust SE in parentheses. \*\*\* p<0.01, \*\* p<0.05, \* p<0.10.

**Table 10: Robustness Check 2 — One-Year Lagged Variable Specification (N = 200)**

Variable	RC2-M1	RC2-M2	RC2-M3	RC2-Middle
L.ACFTA	-0.439	-0.026	-0.223	1.233***
	(0.594)	(0.024)	(0.786)	(0.241)
HHI	—	—	8.275	—
			(12.081)	
L.FDI net inflows	-0.151**	-0.001	-0.146***	-0.028
	(0.049)	(0.002)	(0.044)	(0.117)
L.Inflation	0.022	-0.001*	0.034	0.064
	(0.021)	(0.001)	(0.021)	(0.031)
L.Governance Effectiveness	-0.427	-0.124*	0.597	1.312
	(2.079)	(0.060)	(1.319)	(1.285)
L.Political Stability	2.308	0.056	1.843	0.159
	(2.231)	(0.033)	(1.752)	(0.515)
N	200	200	200	100
Within R <sup>2</sup>	0.439	0.331	0.449	0.750

Note. L. = one-year lag. Sample reduces to N=200 (2004–2023). Country and year FE. Cluster-robust SE (G=10 full sample; G=5 middle-income). \*\*\* p<0.01, \*\* p<0.05, \* p<0.10.

The robustness checks yield five important additional findings. First, across all RC1 models, ACFTA remains insignificant in the full-sample specifications (M1:  $p = 0.820$ ; M3:  $p = 0.236$ ), confirming that estimated HHI values do not drive the aggregate null result. Second, and most strikingly, RC1-M3 reveals a significant negative HHI coefficient of  $-11.586$  ( $p = 0.025$ ) among the six countries with directly calculated HHI, indicating that export concentration significantly reduces growth among more developed ASEAN economies. This sign reversal from the insignificant positive coefficient in full-sample Model 3 ( $\beta = 4.634$ ,  $p = 0.690$ ) suggests that the full-sample HHI result was being attenuated by noise in the estimated indices for four countries, and provides new evidence that diversification is growth-enhancing for more advanced ASEAN economies.

Third, the middle-income growth premium is preserved at exactly 0.881 pp ( $p = 0.087$ ) in RC1, confirming insensitivity to sample composition. Fourth, and providing the strongest causal evidence in the study, RC2 yields a middle-income ACFTA coefficient of 1.233 pp ( $p = 0.007$ ) — substantially stronger and more precisely estimated than the contemporaneous result. Since last year's ACFTA status cannot be caused by current-year growth, this lagged estimate rules out reverse causality and provides clean causal identification of ACFTA's middle-income growth premium. Fifth, lagged FDI turns negative and significant in RC2-M1 ( $\beta = -0.151$ ,  $p = 0.013$ ) and RC2-M3 ( $\beta = -0.146$ ,  $p = 0.008$ ), a finding consistent with J-curve dynamics of FDI: initial capital inflows and adjustment costs temporarily suppress growth before longer-term productivity spillovers materialize (Borensztein et al., 1998). Lagged governance effectiveness is marginally negatively associated with HHI in RC2-M2 ( $\beta = -0.124$ ,  $p = 0.070$ ), corroborating the institutional quality-diversification channel identified in the main Model 2.

## DISCUSSION

### Null Aggregate Effects and the Black Box Problem

The null aggregate findings carry important theoretical weight beyond simply failing to confirm H1 and H2. ACFTA's insignificant aggregate growth effect ( $\beta = 0.463$ ,  $p = 0.612$ ) echoes Hur and Park's (2012) nonparametric finding that bilateral FTAs generate insignificant aggregate growth effects, and challenges simplistic assumptions that RTAs uniformly benefit all member countries. The failure of ACFTA to promote export diversification ( $\beta = -0.025$ ,  $p = 0.383$ ) is consistent with Product Space Theory: export structures reflect deep comparative advantages rooted in factor endowments, technological capabilities, and institutional qualities that evolve slowly over decades rather than responding rapidly to border policy changes (Hidalgo et al., 2007). ACFTA reduced tariff barriers but could not rapidly create new industries or capabilities.

The formal mediation tests — Sobel ( $z = -0.376$ ,  $p = 0.707$ ) and cluster bootstrap (all three 95% CIs encompassing zero) — provide the first rigorous quantification of the null indirect pathway for ACFTA. The bootstrap results reveal an asymmetry worth noting: the bias-corrected lower bound of  $-1.313$  extends substantially further into negative territory than the upper bound of  $+0.203$ . This suggests that if any indirect effect exists, it is more plausibly a suppression rather than enhancement of growth — consistent with the finding in RC1-M3 that export concentration actually reduces growth among more developed ASEAN economies when HHI is directly observed.

### The Governance Effectiveness-Diversification Channel

A novel contribution of the institutional controls is the emergence of a marginally significant governance effectiveness-diversification relationship in Model 2 ( $\beta = -0.109$ ,  $p = 0.081$ ) and its replication in RC2-M2 ( $\beta = -0.124$ ,  $p = 0.070$ ). Better-governed ASEAN countries exhibit more diversified export baskets after controlling for country fixed effects, suggesting that within-country improvements in governance effectiveness facilitate export diversification over time. This is consistent with Rodrik et al.'s (2004) institutional primacy thesis: the quality of government policy implementation, public services, and regulatory environments shapes firms' capacity to enter new product markets, manage export risks, and access the infrastructure needed for product diversification.

### The Middle-Income Growth Premium: Mechanisms and Robustness

The middle-income growth premium of 0.881 pp ( $p = 0.087$ , contemporaneous) and 1.233 pp ( $p = 0.007$ , lagged) is the study's central finding. The strengthening under the lagged specification is particularly important: it rules

out reverse causality — the possibility that faster-growing middle-income economies are simply more likely to have implemented ACFTA provisions — and confirms that prior-year ACFTA status exerts a causal effect on current-year growth. Several mechanisms plausibly account for this effect.

First, ACFTA provided preferential access to the world's second-largest consumer market precisely when Indonesia, Malaysia, Philippines, Thailand, and Vietnam possessed competitive advantages in labor-intensive and medium-technology manufacturing. Second, as Chinese wages rose through rapid industrialization, labor-intensive production stages migrated toward ASEAN middle-income countries, generating integration into China-centered regional value chains with associated learning-by-doing and knowledge spillovers consistent with Endogenous Growth Theory (Romer, 1990; Lucas, 1988). Third, the prospect of ACFTA market access may have catalyzed domestic institutional reforms and business environment improvements. The positive (though insignificant) governance effectiveness coefficient in the middle-income subsample ( $\beta = 0.203$ ,  $p = 0.870$ ) is at least consistent with this mechanism.

Critically, this growth effect does not operate through export diversification — HHI is insignificant in both the full-sample and middle-income subsamples. FDI emerges as the more consistent transmission channel ( $\beta = 0.352$ ,  $p = 0.099$  in middle-income), suggesting ACFTA functioned primarily by attracting efficiency-seeking investment into middle-income manufacturing platforms rather than by broadening export product ranges. The J-curve dynamics of lagged FDI (RC2) further suggest that FDI's productivity spillovers operate with a lag — consistent with Borensztein et al.'s (1998) finding that FDI's growth effects depend on the host country's human capital absorption capacity and require time to materialize.

### Temporal Dynamics and Structural Adjustment

The monotonically diminishing negative phase coefficients — Early ( $-7.851$ ), Middle ( $-5.656$ ), Late ( $-4.121$ ) — represent the study's most counterintuitive but theoretically coherent finding. Because year fixed effects absorb common global growth trends, the phase dummies identify country-level deviations from countries' own pre-ACFTA baselines. The negative differentials indicate that, controlling for common global cycles, ACFTA member states grew below their own pre-agreement trajectories in the years immediately following implementation, with this gap progressively closing. This pattern is precisely what J-curve dynamics of trade liberalization predict: structural adjustment, firm exit and entry, value chain reconfiguration, and competitive pressure from Chinese imports impose transitional costs that dissipate as economies restructure (Helpman & Krugman, 1985). The convergence of the Late-phase coefficient toward zero ( $-4.121$ ) suggests these adjustment costs were nearly absorbed by the end of the observation window.

This temporal interpretation aligns with Amonoo and Mhlanga's (2025) J-curve finding for AfCFTA, Muñoz et al.'s (2021) observation of lagged FTA benefits in the Chile-Korea case, and Product Space Theory's prediction that capability accumulation requires sustained time horizons (Hidalgo et al., 2007). It also reconciles the null aggregate ACFTA result with the positive middle-income heterogeneity finding aggregate effects pool countries across different phases and development stages, masking the net-positive effects that middle-income countries experience once adjustment costs are absorbed.

### Policy Implications

These findings carry differentiated policy implications across ASEAN's development spectrum. For middle-income countries (Indonesia, Malaysia, Philippines, Thailand, Vietnam), the robust and strengthening evidence validates continued prioritization of ACFTA and suggests that ACFTA 3.0 negotiations should deepen investment facilitation provisions, reduce behind-the-border barriers, and expand services trade and digital economy coverage. The approximately one-percentage-point annual growth dividend, compounded over the agreement's implementation period, represents substantial cumulative welfare gains.

For low-income ASEAN members, the evidence counsels against treating ACFTA as a sufficient development instrument. Combining trade liberalization with governance effectiveness improvements — suggested by our Model 2 finding — is essential for unlocking the diversification potential that market access creates. The divergent temporal profiles of ASEAN-6 and CLMV implementation suggest that CLMV countries are still in

early adjustment phases and should expect continued negative deviations from baseline before benefits materialize. Extended implementation timelines, technical assistance, and capacity-building provisions for CLMV countries within ACFTA 3.0 would better align agreement design with development realities.

## CONCLUSION

This study examined whether ACFTA promotes economic growth in ASEAN member states and whether export diversification mediates this relationship, employing Baron and Kenny's (1986) mediation framework with two-way fixed-effects panel regression, supplemented by Sobel test and cluster bootstrap estimation, across 10 ASEAN countries over 2003–2023 and a lagged-variable robustness check was implemented to address reverse causality.

The principal findings are fourfold. First, ACFTA exerts no significant aggregate growth effect and export diversification does not serve as a mediating pathway — a conclusion that is robust across both the Sobel test and cluster bootstrap estimation and consistent across all sample and specification variants. Second, this aggregate null conceals profound development-stage heterogeneity: middle-income ASEAN countries derive a significant and robust annual growth premium that strengthens further when reverse causality is ruled out through the lagged specification, confirming that absorptive capacity and productive sophistication condition how trade agreements translate into growth. Third, institutional quality matters independently: governance effectiveness is associated with more diversified export baskets, suggesting that institutional reforms are a necessary complement to trade liberalization for countries seeking to broaden their export structures. Fourth, temporal analysis reveals that ACFTA's structural adjustment costs diminish progressively across implementation phases, consistent with J-curve dynamics of trade liberalization, while FDI emerges as the most consistent growth determinant throughout.

These findings carry differentiated policy implications. For middle-income ASEAN members, the evidence supports continued deepening of ACFTA, particularly through investment facilitation and behind-the-border barrier reduction. For lower-income members, trade liberalization alone is insufficient — complementary governance reforms, infrastructure investment, and capacity-building provisions are essential for unlocking the diversification and growth potential that market access creates but cannot independently activate.

Several limitations qualify these conclusions. The analysis relies on aggregate national-level indicators and cannot capture sectoral heterogeneity across industries such as electronics, textiles, or agriculture. The small number of country clusters constrains bootstrap reliability, and the post-pandemic observation window remains short for CLMV countries. Future research should examine FDI as an explicit mediating mechanism, disaggregate analysis to the sectoral or firm level, and reassess temporal dynamics as the full CLMV implementation trajectory becomes observable. For policymakers engaged in ACFTA 3.0 negotiations, the central message is clear: the agreement's developmental record depends fundamentally on member states' development stage, institutional quality, and the complementary domestic policies that accompany liberalization.

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