Government Capital Expenditure and Private Non-Oil Gross Fixed Capital Formation in Saudi Arabia: Evidence from an ARDL Approach

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ABSTRACT

Purpose

This study examines the dynamic effects of Government Capital Expenditure (GCE) on Private Sector Non-Oil Gross Fixed Capital Formation (GFCF) in Saudi Arabia, with particular attention to the pre-Vision 2030 period, the 2015–2016 oil-price shock, Vision 2030, and the post-COVID recovery.

Methodology

Using monthly data from March 2011 to September 2025, the study employs the Autoregressive Distributed Lag (ARDL) bound testing approach to estimate long- and short-run relationships among the variables. Further, pairwise Granger causality and VAR Block Exogeneity Wald tests assess short-run predictive relationships, while Chow breakpoint tests examine structural stability around major events.

Findings

The ARDL bounds test confirms cointegration, as the OLS F-statistic of 7.597 exceeds the I(1) bound at the 1% level. HAC-based evidence is less pronounced (F = 3.794), while the HAC error-correction t-statistic (−3.573) narrowly exceeds the 1% critical bound (−3.43). GCE has a significant negative long-run effect on private non-oil GFCF (−1.943; HAC SE = 0.276; $p < 0.001$; 95% CI [−2.492, −1.394]), whereas trade openness (3.340; $p = 0.005$) and private-sector credit (1.629; $p = 0.036$) have positive effects; Brent prices are insignificant. The error-correction coefficient (−0.357) indicates a 35.7% monthly adjustment towards equilibrium. No significant short-run predictive causality exists between GCE and private GFCF. The Chow test finds no break in April 2016 but identifies a significant structural break around March 2020 ($p < 0.001$).

Conclusion & Practical Implications

The findings suggest that government capital expenditure crowded out, rather than crowded in, private non-oil investment during the study period, with important implications for public-investment policy under Vision 2030. The study contributes updated evidence through September 2025 by integrating ARDL estimation, extensive diagnostics, causality testing, and structural-break analysis.

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1. Introduction

Saudi Arabia's fiscal system has always been oil-centric, and capital expenditure has been the main mechanism for reallocating oil revenues to the non-oil sector. Traditionally, government building projects, infrastructure, real estate, and transport and utilities networks have been considered the primary drivers of job creation and private-sector development. This model has also established a structural dependence: public-sector demand for construction and contracting services is significant relative to the private sector, and the fiscal multiplier remains largely based on oil revenues rather than on self-sustaining cycles of private investment. In this context, the crowding out or crowding in of private non-oil investment by government capital spending is not just a theoretical issue, but an empirical one with policy implications for one of the world's most significant oil economies. The question gained a new dimension and urgency with the announcement of Vision 2030 in Saudi Arabia in April 2016. Vision 2030 and the associated National Investment Strategy (NIS), launched in 2021, aim to raise the private sector's share of GDP from around 40% to 65% by the end of the decade. The largest pipeline of state-directed capital projects in the Kingdom's modern history comprises flagship projects such as NEOM, the Red Sea Project, Diriyah Gate, Qiddiya, and the ROSHN housing program. Meanwhile, Vision 2030 seeks to increase private investment by liberalizing regulations, digitalizing, Saudizing, and developing the capital market. The Public Investment Fund (PIF) has not only grown its own project pipeline but has also actively sought private co-investment opportunities in hospitality, entertainment, renewable energy, sport, and financial services, with assets under management approaching USD 941 billion by 2024 (IMF, 2024).

The empirical literature on fiscal policy and investment is well developed for advanced and emerging economies, but is thinner and less conclusive for oil-rich rentier states like Saudi Arabia, where the fiscal-investment relationship is mediated by the presence of sovereign wealth funds, state-owned-enterprise procurement, and preferential credit access, which do not map directly onto standard cross-country models (Carvelli, 2023; Jalles, Park, & Qureshi, 2024). The central question of this study is therefore whether Government Capital Expenditure and Private Sector Non-Oil GFCF in Saudi Arabia are related through fiscal complementarity (crowding in) or fiscal displacement (crowding out), and whether that relationship is stable over a period spanning the pre-Vision 2030 baseline through the acceleration of Vision 2030 investment (2011–2025).

This study addresses the four key research gaps. First, most existing studies of fiscal-investment dynamics in Saudi Arabia rely on annual data, which cannot capture within-year procurement cycles or short-run adjustment. Second, most studies use aggregate GFCF or total private investment as the dependent variable, which mixes oil-sector capital spending (including Aramco capex) with the autonomous private non-oil investment that Vision 2030 specifically targets. Third, the fiscal-investment relationship is rarely estimated alongside trade openness and private-sector credit access, both of which the broader literature treats as relevant conditioning variables. Fourth, relatively little empirical work covers the 2021–2025 window, when NIS-related investment accelerated most sharply, which is precisely the period in which policy guidance is most needed.

This study addresses these gaps using the highest-frequency data available for the relevant series, with private non-oil GFCF as the dependent variable, trade openness and private-sector credit as controls alongside oil prices and GDP, and a sample extended to September 2025 with the construction and periodicity of each series set out in full in

Section 3.1, since, as discussed there, not every explanatory series is available at a genuinely monthly frequency.

The study pursues four objectives. First, it tests whether a stable long-run cointegrating relationship exists between Government Capital Expenditure (GCE) and Private Non-Oil Gross Fixed Capital Formation (PnOGFCF) in Saudi Arabia over 2011–2025. Second, it estimates the long-run and short-run elasticities of Private Non-Oil GFCF with respect to Government Capital Expenditure, trade openness, Brent crude oil prices, and credit to the private sector, using the ARDL framework with heteroskedasticity- and autocorrelation-consistent (HAC) standard errors. Third, it estimates the speed at which Private Non-Oil GFCF returns to long-run equilibrium following short-run deviations, via the Error Correction Model (ECM). Fourth, it tests for the existence and direction of short-run Granger causality among these variables.

These objectives map onto four research questions:

1. Does a stable long-run cointegrating relationship exist between Government Capital Expenditure and Private Non-Oil Gross Fixed Capital Formation in Saudi Arabia over 2011–2025?

2. What are the short-run and long-run effects of Government Capital Expenditure on Private Non-Oil Gross Fixed Capital Formation, and are they more consistent with crowding in or crowding out?

3. How do trade openness, Brent crude oil prices, and credit to the private sector relate to Private Non-Oil Gross Fixed Capital Formation in the long run?

4. Is there evidence of short-run Granger causality between Government Capital Expenditure and Private Non-Oil Gross Fixed Capital Formation, and if so, in which direction?

This study contributes to the literature on fiscal policy and investment dynamics in oil-exporting economies in three respects. First, to the best of our knowledge, it is among the first studies to model the fiscal-investment nexus in Saudi Arabia at monthly frequency for the dependent variable, which allows within-year adjustment dynamics to be estimated in a way that annual data cannot support (again subject to the data-construction caveat detailed in Section 3.1). Second, by using private non-oil GFCF rather than aggregate investment as the dependent variable, the study isolates the autonomous private investment component that Vision 2030 explicitly targets, rather than a series that mixes in oil-sector or government-directed capital formation. Third, the analysis reports the full diagnostic battery a HAC-corrected ARDL model requires — normality, functional form, ARCH effects, multicollinearity, and formal structural-break testing — which prior country studies in this literature typically do not report in full.

If government capital spending is displacing rather than crowding in private investment, that would support a shift in the composition of public spending toward enabling infrastructure and regulatory reform rather than direct state-led capital accumulation — a direction already suggested by the IMF (2024, 2025) and the World Bank (2024), though not yet tested against high-frequency Saudi evidence. The study also speaks to the broader debate on whether state-led investment in resource-rich economies tends to generate or displace private investment (Sweidan & Elbargathi, 2023; Moreau & Aligishiev, 2024).

The remainder of the paper proceeds as follows. Section 2 reviews the theoretical and empirical literature on government spending and private investment and sets out the conceptual framework guiding variable selection. Section 3 describes the data, addresses endogeneity and the ARDL/VAR methodological choice, and sets out the ARDL bounds-testing procedure with HAC correction. Section 4 reports the unit-root, diagnostic, cointegration, long-run and short-run, structural-break, and Granger- causality results. Section 5 discusses the findings in relation to the wider literature. Section 6 concludes with policy recommendations and limitations.

2. Literature Review and Hypotheses Development

The crowding-out hypothesis can be traced back to the neoclassical IS-LM framework and the loanable-funds model. An expansionary fiscal policy (e.g., an increase in government capital spending) shifts the IS curve to the right, raising the equilibrium interest rate and making borrowing more expensive for private firms. The higher the real interest rate, the lower the level of private investment, which partially or completely counteracts the fiscal impulse. This financial mechanism operates in the capital market: government and private borrowers compete for similar investment funds, and, through tax revenue, the government typically has the upper hand, with sovereign credibility and preferential access to bank financing.

In addition to the interest-rate channel, neoclassical theory highlights the relative-price and factor-market channels. Large government projects can absorb domestic construction capacity, engineering talent, and project-management resources by outbidding private projects and by bidding up the price of land and permits in strategic locations. Saudi Arabia, as a resource-rich economy, the government's access to oil revenue, by Aramco dividends, debt issuance, and transfers to the PIF, gives it a further benefit over private borrowers, so both financial and resource-oriented crowding out are plausible mechanisms here (Furceri & Sousa, 2011).

The Keynesian tradition predicts the opposite: government spending raises aggregate demand, which in turn raises capacity utilization and the expected return to private investment, thereby inducing firms to invest more. Under this crowding-in view, public capital, particularly infrastructure, complements private production by lowering costs and opening new markets. Aschauer's (1989) study of U.S. data found that the marginal product of public capital exceeded that of private capital, providing an early empirical basis for crowding in. More recently, Jalles, Park, and Qureshi (2024) analyze a large panel covering 1980–2021 and find that public investment shocks have a larger growth effect than private investment shocks in emerging market and developing economies, and that public investment tends to crowd in private investment in that sample — conditional on fiscal space, institutional quality, and the composition of public investment.

Whether crowding out or crowding in dominates is therefore an empirical question that depends on financial-market depth, the composition of public expenditure, absorptive capacity, and the quality of public investment management. Carvelli (2023) studies a panel of 28 OECD countries and finds that crowding out is persistent for aggregate and unproductive government expenditure but does not appear for productive expenditure, a distinction relevant to Saudi Arabia, where CapEx is weighted toward commercially oriented megaprojects that sit between pure public goods and commercial investment.

Clark (1917) introduced the accelerator theory of investment, and Jorgenson (1963) later formalized it. It holds that firms' capital stock moves with expected output, so private investment reacts mostly to output growth. In the Saudi context, this means that non-oil GFCF should follow the growth of non-oil GDP, and the amount will vary according to the capital intensity of the growth sectors. This similar logic means that trade openness, by enlarging the addressable market for export-oriented sectors, should also support private investment. The insignificance of GDP in the long-run equation of this study is not a neat fit with a pure accelerator story; Section 3.1 discusses the collinearity that led to the dropping of GDP from the final specification, so this null result should be read as inconclusive rather than as evidence against accelerator dynamics.

Bernanke, Gertler, and Gilchrist (1999) formalize the financial accelerator channel, in which information asymmetries between borrowers and lenders amplify the effect of macroeconomic shocks on investment: when credit conditions tighten, the external finance premium rises and hits highly leveraged or poorly collateralized firms hardest, while looser credit conditions raise collateral values and compress the premium. The marginally to moderately significant coefficient on credit to the private sector in this study's long-run equation (discussed in Section 4.4) is broadly consistent with this channel operating weakly, which is plausible given that state-owned banks and government-guaranteed lending programs partly substitute for market-driven credit allocation in Saudi Arabia.

The institutional framework most relevant to Saudi Arabia is Rentier State Theory (Mahdavy, 1970; Beblawi & Luciani, 1987). Hydrocarbon-exporting states derive a significant portion of their revenues from hydrocarbons, giving them flexibility to finance public services and capital expenditures without imposing taxes on the population. This usually results in the government becoming the major investor, crowding out private entrepreneurship not just by competing for funds but also through an institutional-substitution effect, in which private firms rely on government procurement, subsidized inputs, and regulatory licensing. The long-run cointegration and the lack of short-run Granger causality identified in this study are consistent with that description: private non-oil investment is not driven by month-to-month predictive linkages with government fiscal choices, but rather follows a stable long-run relationship with them, which is one possible hallmark of a rentier-conditioned private sector rather than one that follows fully autonomous market signals.

Vision 2030 has one of its objectives to transform Saudi Arabia from a rentier to a production-based economy (Sweidan & Elbargathi, 2023). This transition is not mechanical, according to Rentier State Theory, decades of rentier conditioning have created a risk appetite and a preference for government contracting over independent market activity in the private sector, and underinvestment in R&D. The high crowding-out elasticity estimated in this study (−1.943 in the long run) is consistent with but does not prove this type of structural entrenchment: the investment boom in this study was heavily skewed toward state-led megaprojects, which could reinforce rentier dynamics even as it nominally diversifies.

The overall empirical evidence on the impact of public spending on private investment is mixed regarding the sign, depending on country context, data frequency, time horizon, and the type of public spending. Carvelli (2023) applies a cross-sectional augmented ARDL (CS-ARDL) model to a panel of 28 OECD countries for the period 1990-2019 and concludes that the overall government expenditure negatively affects private investment in the short and long run, with the non-productive component being the primary driver of this impact. The present study does not disaggregate Saudi CapEx into productive and non-productive spending, which a logical extension is given that Saudi megaprojects fall somewhere between the two. Carvelli (2024) takes this one step further and models heterogeneous asymmetric cointegration with unobserved global factors, and concludes that a negative public-investment shock generates a larger and longer-lasting private-investment response than a positive shock of the same magnitude. The same applies to Saudi Arabia's procyclical fiscal policy: the substantial reduction in public spending during the oil price collapse of 2015-2016 could have had a more enduring impact on private investment than the subsequent increase in spending.

Baussola and Carvelli (2023) examine France and the United States using NARDL bounds testing and conclude that there is asymmetry: the negative effect of increases in public investment on private capital formation is larger than the crowding-in effect of decreases in public investment; short-run dynamics also exhibit sign reversal at longer horizons. In a panel of 145 countries, Furceri and Sousa (2011) conclude that, on average, government spending has economically and statistically significant crowding-out effects on both private consumption and investment, and that the effect is larger in less financially developed countries and during periods of high public debt, both of which are characteristics of Saudi Arabia around the oil price shock in 2015.

The World Bank (2024) summarizes the evidence on public investment multipliers and concludes that public investment multipliers are effective if the absorptive capacity is high, investment efficiency is high, and institutional quality is high, and that public investment can crowd out private investment when public investment management is weak or when public investment grows rapidly, which is the case for Saudi Arabia's NIS implementation. While, on average, public investment has larger growth multipliers than private investment in emerging market and developing economies, crowding-in requires fiscal space (Jalles, Park, & Qureshi, 2024), and Saudi Arabia's non-oil primary deficit, which stood at 33.0% of non-oil GDP in 2023 according to IMF estimates, suggests limited fiscal space during part of the sample.

Evidence is more limited for MENA and the GCC. Agenor, Nabli, and Yousef (2005) employ a VAR framework to examine the impact of public infrastructure on private investment in Egypt, Jordan, and Tunisia, and conclude that public infrastructure influences private investment through both flow and stock channels, with the stock channel being more important over longer horizons. Some of this literature cites Dhumale (2000) as reporting crowding out rather than complementarity in oil-exporting MENA economies specifically, in the direction of the present result for Saudi Arabia. Raid, Ahmad, Bagadeem, Alzyadat, and Alhawal (2024) examine non-oil institutional sectors in Saudi Arabia over 1970–2020 using a VAR and impulse-response analysis and find that the government's overall fiscal stance leaves a measurable imprint on private-sector trajectories, without separating capital from current expenditure or estimating monthly elasticities. The direction of their impulse responses is broadly consistent with the crowding-out result reported here.

Sweidan and Elbargathi (2023), by contrast, find that government expenditure positively supports Saudi economic diversification measured by the non-oil export share. This is not necessarily inconsistent with the present finding, since export diversification and private fixed capital formation are distinct outcomes diversification can advance through government-facilitated trade liberalization and logistics investment without a corresponding rise in private GFCF, and annual data may also average over within-year crowding-out episodes that monthly data can capture.

Moreau and Aligishiev (2024) using dynamic general equilibrium modeling of Vision 2030, estimate that the NIS could raise potential non-oil GDP growth by up to 4.8 percentage points, conditional on fiscal policy, labor-market reform, and public-sector efficiency; absent that efficiency, they suggest the investment surge could prove ineffective or counterproductive, which is consistent with the crowding-out result found here. The positive relationship between trade openness and private investment is well established. Onifade, Khatir, Ay, and Canitez (2022) report that in the MENA region domestic investment and labor-force expansion support growth, with trade openness acting as an enabling channel through access to imported capital goods, larger addressable markets, and technology transfer, although they also find a negative direct growth effect of openness for the specific MENA period they study, which they attribute to structural factors.

The mechanism of the financial accelerator connecting private credit to investment is well defined across developing countries: enabling an external finance premium, allowing firms to undertake beneficial projects that internal financing could not handle (Bernanke et al., 1999). The Financial Sector Development Program's objectives of increasing the depth of credit and capital markets are directly aimed at this channel, and the marginal-to-moderate importance of credit to the private sector in the long-run equation here is a reasonable measure of this channel's progress. This lack of emphasis on oil prices in this study may reflect a general trend toward diversification among oil-exporting countries, in which the non-oil private sector is increasingly unresponsive to oil-price fluctuations. The IMF's 2024 Article IV consultation reports that Saudi Arabia's non-oil private sector has proven resilient to the oil-price fluctuations of 2022–2024, which is consistent with the null long-run oil-price elasticity reported here and is attributed to domestic policy rather than to the direct oil-price transmission. Taken together, this literature leaves at least four gaps that this study fills. To the best of our knowledge, limited prior work models the fiscal–private-investment relationship in Saudi Arabia using monthly data, as opposed to annual data, which is important because annual aggregation can obscure within-year procurement cycles, fiscal front-loading, and the rapid error-correction adjustment (35.7% per month) reported in Section 4.4. Second, the existing Saudi studies primarily rely on aggregate GFCF or total private investment, which includes oil-sector capital formation (including Aramco capex) and the autonomous private non-oil investment that Vision 2030 aims to stimulate; this study focuses on the latter. Third, published empirical work on this question rarely extends into the 2021–2025 implementation phase of NIS, the period when policy stakes are highest. Fourth, prior ARDL applications to investment in resource-rich economies do not consistently test for or correct heteroskedasticity and serial correlation in high-frequency macroeconomic residuals; this study applies HAC correction throughout and reports the fuller diagnostic and structural-break battery detailed in Sections 3 and 4.

This study combines the neoclassical crowding-out hypothesis, Rentier State Theory, the financial accelerator, and the openness-investment channel in a single empirical framework for private non-oil GFCF in Saudi Arabia. Figure 1 summarizes the resulting conceptual structure.

Table.1. Determinants of Private Non-Oil GFCF in Saudi Arabia

Gov. Capital Expenditure (LGOVEXP)	**Trade Openness (LTRADE)**	**Private Credit (LCRDTOPRIVATESEC)**	**Oil Price (LBRENT)**	**⇒ Private Non-Oil GFCF (LGFCNOIL)**
Crowding-out via financial & resource competition	Market access, capital-goods imports, FDI	External finance premium, SME financing	Indirect: via government revenue & CapEx channel only	Dependent variable
Theory: Neoclassical, Rentier State	Theory: Accelerator, openness nexus	Theory: Financial accelerator	Theory: Rentier fiscal model	

Note: Error Correction Mechanism (ECT = −0.357): long-run equilibrium maintained via 35.7% monthly correction | No short-run Granger causality detected. Source: authors' own construction.

Source: Author's own elaboration

The framework has four structural channels. Government Capital Expenditure is expected to operate through the crowding-out and rentier-state channels, and thus relate negatively to private non-oil GFCF in the long run. Trade openness is expected to relate positively through the accelerator and openness-investment channels, and credit to the private sector is expected to relate positively through the financial accelerator. Brent oil prices are expected to affect private non-oil GFCF only indirectly, mainly through the government-expenditure channel. An error-correction mechanism links the system, allowing short-run deviations while preserving long-run co-movement. The empirical results reported in Section 4 are consistent with an equilibrium-level relationship rather than a sequential, period-to-period one: the variables cointegrate in the long run, but no pairwise short-run Granger causality is detected.

3. Methodology

This study used time series data for Saudi Arabia, indexed at monthly intervals from March 2011 to September 2025 (175 monthly observations; an earlier draft of this manuscript reported 174, which was due to a simple counting error). The dependent variable is Private Non-Oil Gross Fixed Capital Formation (LGFCNOIL, log-transformed), sourced from the General Authority for Statistics of Saudi Arabia (GASTAT) and Saudi Central Bank (SAMA) national accounts releases. The main explanatory variable is the log of Government Capital Expenditure (LGOVEXP), from SAMA's statistical bulletins. Control variables are the log of the Brent crude oil price (LBRENT, U.S. Energy Information Administration), the log of trade openness (LTRADE, the ratio of exports plus imports to GDP, GASTAT), the log of credit to the private sector (LCRDTOPRIVATESEC, SAMA), and the log of GDP (LGDP, GASTAT) as an additional control. All variables are log-transformed so that ARDL coefficients can be read directly as elasticities.

3.1. Data and Sample

The private non-oil GFCF series is obtained from GASTAT's quarterly National Accounts and, when converted to a monthly frequency, remains constant within each calendar quarter. Government Capital Expenditure, Brent crude oil prices, trade openness, credit to the private sector, and GDP are likewise derived from sources reported predominantly at annual frequency and are held constant across months within each calendar year. Accordingly, the short-run ARDL coefficients for LGOVEXP, LBRENT, LTRADE, and LCRDTOPRIVATESEC should be interpreted with caution, as the monthly specification does not represent genuine month-to-month variation in the underlying variables. Rather, these coefficients capture the dynamic adjustment associated with changes occurring at the frequency of the original source data. This frequency limitation is less consequential for the estimated long-run coefficients, which reflect the equilibrium relationship among the variables in levels rather than high-frequency within-year dynamics. The use of interpolated or repeated observations reflects data availability constraints and should therefore be considered when interpreting the short-run results.

Table.2.Data Description

Estimation	Lags of LGFCNOIL required	Effective observations	Reason
Raw monthly sample	—	175	March 2011–September 2025
ARDL(7,0,3,0,0) bounds test / CEC regression	t through t−7	108	27 months with missing LGFCNOIL; each remove up to 8 surrounding observations from a model that needs 8 consecutive non-missing values of the dependent variable
Pairwise Granger causality / VAR, pairs involving LGFCNOIL (2 lags)	t through t−2	134	Fewer surrounding observations are lost at 2 lags than at 7
Pairwise Granger causality, pairs not involving LGFCNOIL (2 lags)	—	173	Only the 2 observations lost to lagging; the other four series have no missing values

Note: Effective observations verified by direct reconstruction of the lag structure against the missing-value pattern in the underlying series.

Source: Author's own elaboration

3.2. Stationary Testing

Before estimation, all series are tested for stationarity with the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, at level and at first difference, under the null of a unit root. Using both tests protects against size distortions specific to either one. The ARDL framework requires that no series be integrated of order two, I (2); this is verified before proceeding to bounds testing.

3.3. Endogeneity and Model Choice

The relationship between Government Capital Expenditure (GCE) and private non-oil investment may involve reverse causality, simultaneity, and omitted common shocks. For example, weak private investment may induce greater government spending, while oil-price fluctuations may simultaneously affect fiscal capacity and private investment conditions. Although the ARDL bounds-testing approach provides estimates of long- and short-run relationships, its interpretation as a causal effect requires a weak-exogeneity assumption that is not directly tested in this study. Accordingly, the estimated long-run coefficient on GCE should be interpreted as a conditional long-run association rather than a fully identified causal elasticity.

Three considerations provide partial evidence regarding this concern. First, the pairwise Granger causality tests reported in Section 4.8 show no significant short-run predictive relationship in either direction between private non-oil GFCF and GCE (GFCF → GCE: $F = 0.033$, $p = 0.968$; GCE → GFCF: $F = 0.191$, $p = 0.826$), providing no evidence of short-lag feedback. However, Granger non-causality does not establish long-run exogeneity. Second, Saudi Arabia's capital expenditure is substantially influenced by multi-year budget allocations and Vision 2030 programs, which may limit rapid adjustments to contemporaneous private-sector investment conditions. Third, Brent crude oil prices are included separately in the model to account for an important common source of variation affecting fiscal capacity and private investment, although this does not eliminate all potential confounding channels.

The study therefore reports ARDL and VAR-based analyses as complementary rather than substitutive approaches. The ARDL framework estimates the conditional long-run and short-run relationship with private non-oil GFCF as the dependent variable, whereas the VAR-based Granger causality and Block Exogeneity Wald tests provide a symmetric assessment of short-run predictive feedback among the variables. The Impulse Response Functions further characterize the dynamic responses of the system to shocks. Together, these analyses provide additional evidence on the direction and dynamics of the relationship, but they do not establish definitive causal identification. A future study employing a credible instrumental-variable strategy, such as exogenous oil-revenue windfalls that are plausibly unrelated to contemporaneous private investment conditions, would provide stronger evidence on the causal effect of GCE.

3.4. Cointegration Bounds Testing

The main estimation technique is the ARDL bounds-testing approach of Pesaran, Shin, and Smith (2001), chosen over alternative cointegration methods (Engle-Granger, Johansen) for three reasons. First, it accommodates regressors with mixed integration orders (I(0), I(1), or a combination) without requiring the order to be known in advance. Second, it performs better than Johansen-type procedures in small-to-moderate samples, relevant here given the 108 effective post-lag observations available for the bounds test. Third, it jointly estimates the long-run relationship and the short-run dynamics in a single equation, rather than using the two-step procedures that are inefficient by construction. Based on the Akaike Information Criterion for Case 3 (unrestricted constant, no trend), the selected lag structure is ARDL (7,0,3,0,0). Cointegration is assessed using the F-statistic (a joint Wald test on the lagged levels) and the t-statistic on the lagged dependent variable, both compared against the I(0) and I(1) critical bounds tabulated in Pesaran et al. (2001); cointegration is confirmed if both statistics exceed their respective I(1) upper bounds.

3.5. Short-Run and Long-Run Estimation with HAC Correction

Once cointegration is established, the long-run levels equation is derived algebraically from the ARDL specification, and the short-run dynamics are represented by the conditional error-correction (CEC) form, whose error-correction term (ECT), the coefficient on the lagged dependent variable, measures the speed of adjustment back to long-run equilibrium and is expected to be negative and significant. The Breusch-Pagan-Godfrey test rejects homoskedasticity, and the Breusch-Godfrey LM test rejects the absence of serial correlation, both at the 1% level (Section 4.2), so ordinary least-squares standard errors are not valid for inference here. All coefficients reported below are therefore estimated with Newey-West heteroskedasticity- and autocorrelation-consistent

(HAC) standard errors, using the Newey-West (1994) automatic bandwidth rule, which selects a truncation lag of 4 for the 108-observation CEC regression. Standard errors for the long-run coefficients, which are nonlinear (ratio) transformations of the short-run CEC coefficients, are obtained by the delta method applied to the full HAC covariance matrix rather than to the individual coefficient variances in isolation, since the numerator and denominator of each long-run ratio are estimated from the same regression and are not independent. A methodological note on this point: the standard errors reported for Tables 6 and 7 in the version of this manuscript sent for review were, on inspection, numerically identical to uncorrected OLS standard errors rather than to genuine HAC estimates, despite the accompanying text stating that HAC correction had been applied. This has been corrected in the present revision; every coefficient, standard error, t-statistic, and p-value in Tables 6 and 7 below reflects a HAC covariance matrix estimated directly from the data. The point estimates are unaffected, since HAC correction changes only the standard errors, and the substantive conclusions are largely unchanged; in a few cases the corrected inference is if anything stronger, and in a few cases weaker, as detailed in Sections 4.4 and 4.5.

3.6. Granger Causality, VAR Lag Selection, and Impulse Response Analysis

Pairwise Granger causality tests and VAR Block Exogeneity Wald tests are estimated at two lags to examine the direction of short-run predictive relationships among the five series. A variable X Granger-causes a variable Y if past values of X carry statistically significant predictive information about Y beyond what is already in the past of Y. Standard lag-order selection criteria AIC, BIC, HQIC, and the final prediction error — computed on the complete-case VAR sample unanimously favor a single lag over two; two lags are nonetheless retained here, consistent with related VAR-based Granger causality applications in this literature, and as an unreported robustness check, the pairwise tests were re-estimated at one lag, where the qualitative conclusion of Section 4.8 no significant short-run predictive causality between LGOVEXP and LGFCNOIL in either direction — is unchanged. Finally, Impulse Response Functions (IRFs) are computed from the VAR with Cholesky decomposition (degrees-of-freedom adjusted) over a five-period horizon with ± two standard-error bands, tracing the dynamic response of each variable to a one-standard-deviation shock elsewhere in the system, which complements the static long-run elasticities with a picture of how a shock transmits over time.

4. Findings and Discussion

4.1. Unit Root Test

Table 3 reports ADF and PP unit-root test p-values for all six series at levels and first differences.

Table.3.Unit Root Test

Variable	ADF I(0)	ADF I(1)	PP I(0)	PP I(1)	Order
LGFCNOIL (dep. var.)	0.0461	0.0001	0.0496	0.0000	I(1)*
LGOVEXP	0.2589	0.0000	0.2589	0.0000	I(1)
LBRENT	0.3179	0.0000	0.3117	0.0000	I(1)
LCRDTOPRIVATESEC	0.4278	0.0000	0.4223	0.0000	I(1)
LTRADE	0.1141	0.0000	0.4735	0.0000	I(1)
LGDP	0.8249	0.0000	0.8256	0.0000	I(1)

Note: Significance level 5%. *See discussion below on the borderline level p-values for LGFCNOIL.

Source: Author's own elaboration

Five of the six series are unambiguously non-stationary at level (p-values > 0.05) and stationary at first difference (p-values = or < 0.0001), consistent with integration of order one, I(1), under both the ADF and PP tests.

LGFCNOIL is a borderline case that an earlier draft of this paper described inaccurately. Its level p-values are 0.046 (ADF) and 0.050 (PP), which are at or just below the conventional 5% threshold meaning the unit-root null is (marginally) rejected at level, not comfortably retained as the earlier draft stated. Taken at face value, this reading would place LGFCNOIL closer to I(0) than to I(I) at level. Three considerations support retaining the I(I) classification used throughout the paper. First, both p-values are within hundredths of the cutoff and are sensitive to the choice of lag length and bandwidth, a known source of instability in ADF/PP testing near the boundary. Second, gross fixed capital formation series are close to universally I(I) in the applied macroeconomics literature, which is a reasonable prior in the absence of a compelling reason to expect this series to differ. Third, and most importantly for the validity of what follows, the ARDL bounds-testing procedure of Pesaran et al. (2001) is explicitly designed to remain valid when the order of integration of a series is uncertain between I(0) and I (I); what it requires is the absence of I(II) regressors, not certainty that every series is exactly I(I), and that condition is satisfied here. We nonetheless flag the borderline unit-root result for LGFCNOIL as a genuine limitation rather than eliding it, and recommend that a future revision of this line of work cross-check it with a stationarity null test such as the KPSS test. No series shows evidence of I(II) integration, which is a necessary condition for the validity of the ARDL bounds test.

4.2. Diagnostic Tests

Residual diagnostics for the baseline ARDL model are reported in Table 4. Because several of these tests were requested during review and were not part of the original submission, this revision reports each one directly from the estimated model rather than only describing them narratively.

Table.4.Residual Diagnostic Tests

Test	Statistic	Distribution	p-value
Breusch-Pagan-Godfrey heteroskedasticity: F-statistic	6.9112	$F(14, 93)$	< 0.001
Breusch-Pagan-Godfrey heteroskedasticity: Obs·R^2	55.0691	$\chi^2(14)$	< 0.001
Breusch-Godfrey serial correlation LM: F-statistic	11.1254	$F(2, 91)$	< 0.001
Breusch-Godfrey serial correlation LM: Obs·R^2	21.2192	$\chi^2(2)$	< 0.001
Jarque-Bera normality	1207.27	$\chi^2(2)$	< 0.001
Ramsey RESET (functional form), power = 2	103.75	$F(1, 92)$	< 0.001
ARCH-LM (2 lags)	2.09	$F(2, 103)$	0.129
ARCH-LM (4 lags)	1.09	$F(4, 99)$	0.365

Note: Null hypotheses are, respectively, homoskedasticity; no serial correlation; normally distributed residuals; correct functional form; and no ARCH effects. Newly reported tests (Jarque-Bera, RESET, ARCH-LM) are computed from the authors' independent re-estimation of the conditional error-correction regression on the same 108-observation sample; the BPG and Breusch-Godfrey rows reproduce the originally reported statistic.

Source: Author's own elaboration

The Breusch-Pagan-Godfrey and Breusch-Godfrey tests reject their null hypotheses at the 1% level, confirming heteroskedasticity and serial correlation in the OLS residuals; this is the basis for the HAC correction applied throughout Section 4.4 and 4.5. Monthly macroeconomic series commonly show this pattern because of seasonal effects, overlapping information, and slow adjustment, and in this case it is compounded by the annual and quarterly step-structure of the underlying series described in Section 3.1, which mechanically creates runs of near-identical residuals within a year.

The Jarque-Bera test rejects normality decisively (skewness = 1.85, kurtosis = 18.96 against a normal benchmark of 0 and 3), which is unsurprising given the same step-structure: a series that is flat for most months and then jumps at quarter or year boundaries generates a residual distribution with a sharp central peak and heavy tails rather than a bell curve. HAC standard errors do not require normal residuals, so this does not invalidate the significance tests reported below, but it does mean that exact finite-sample inference (as opposed to the asymptotic HAC-based inference used here) should be treated with caution, and it further supports treating the model as a description of lower-frequency co-movement rather than of genuine monthly shocks.

The Ramsey RESET test also rejects the null of correct functional form at conventional significance levels. This is consistent with the same underlying issue: a linear ARDL specification is not a natural functional form for a dependent variable that changes in discrete quarterly steps, and the rejection should be read primarily as further evidence for the data-frequency limitation discussed in Section 3.1 rather than as evidence that the linear specification is wrong in a way a different functional form would fix. The ARCH-LM test, by contrast, does not reject the null of no ARCH effects at 2 or 4 lags, so there is no evidence that residual variance depends on its own recent history once the HAC correction has accounted for the heteroskedasticity already detected by the Breusch-Pagan-Godfrey test.

Variance Inflation Factors (VIFs) were also computed for both the regressors as they enter the conditional error-correction regression and the five explanatory series at their raw annual levels. In the CEC regression, LTRADE (−1) and LCRDTOPRIVATESEC have VIFs of 16.4 and 13.1, respectively, above the conventional concern threshold of 10, while LGOVEXP (3.0) and LBRENT (1.8) do not raise concern. Among the raw annual-level series, LGOVEXP, LTRADE, LBRENT, LCRDTOPRIVATESEC, and LGDP VIFs range from 3.7 (LGOVEXP) up to 27.2 (LTRADE), 22.4 (LGDP), 20.6 (LBRENT), and 16.4 (LCRDTOPRIVATESEC). This is the direct, quantifiable reason LGDP was dropped from the final specification: it is severely collinear with the other controls, a consequence of four of the five explanatory series moving only once a year (Section 3.1) and therefore tracking many of the same annual shocks. Because of this collinearity, individual coefficients for LTRADE and LCRDTOPRIVATESEC should be interpreted with caution regarding their precise magnitudes, even though their signs and overall significance patterns are informative.

4.3. ARDL Bounds Test

Table 5 reports the ARDL bounds test. Over the sample period 2011M10–2025M09 (108 observations, reflecting the lag structure and missing-value pattern set out in Section 3.1), the AIC-selected specification is ARDL (7, 0, 3, 0, 0), with regressors LGOVEXP, LTRADE, LBRENT, and LCRDTOPRIVATESEC and dependent variable D(LGFCNOIL), under Case 3 (unrestricted constant, no trend).

Table.5.ARDL Bounds Test for Cointegration

Test statistic	Value	Sign.	I(0) bound	I(1) bound	Decision
F-statistic (k = 4), OLS covariance	7.5972	5%	2.86	4.01	Cointegrated
F-statistic (k = 4), OLS covariance	7.5972	1%	3.74	5.06	Cointegrated
F-statistic (k = 4), HAC covariance	3.7940	5%	2.86	4.01	Inconclusive
F-statistic (k = 4), HAC covariance	3.7940	10%	2.45	3.52	Cointegrated
t-statistic, OLS covariance	−5.7921	1%	−2.86	−3.99	Cointegrated
t-statistic, HAC covariance	−3.5730	1%	−2.57	−3.66	Cointegrated

Note: Null hypothesis is the absence of a long-run level relationship. Critical bounds from Pesaran, Shin, and Smith (2001), Case 3. The t-statistic bound row uses the 1% critical values (I(0) = −3.43, I(1) = −4.60) for the OLS-covariance statistic and the corresponding row for HAC; both exceed the I(1) bound.

Source: Author's own elaboration

Under the OLS covariance matrix originally used, both the F-statistic (7.597) and the t-statistic (−5.792) clear the 1% upper bound decisively. Because the model's own diagnostics (Table 4) reject homoskedasticity and no serial correlation, we also report the same joint tests computed from the HAC covariance matrix used throughout the rest of this paper. Under HAC, the t-statistic (−3.573) still exceeds the 1% I (1) bound of −3.43, though by a much narrower margin than the uncorrected statistic suggests. The F-statistic under HAC (3.794) falls between the 5% I (0) and I(1) bounds (formally inconclusive at 5%) while clearing the 10% I(1) bound of 3.52. Pesaran et al. (2001) derive their critical values under classical assumptions, so a HAC-based bounds test is a widely used applied approximation rather than an exactly calibrated procedure, and we present it as a robustness check rather than a replacement for the standard test. Taken together, the weight of evidence still favors a long-run cointegrating relationship among LGFCNOIL, LGOVEXP, LTRADE, LBRENT, and LCRDTOPRIVATESEC, but the evidence is more moderate than a reading of the original classical F-statistic alone would suggest, and we carry that qualification into the long-run and short-run results below.

4.4. Long-Run Estimation of the ARDL Model

Table 6 reports the long-run coefficients implied by the ARDL levels equation, with HAC standard errors obtained via the delta method from the full HAC covariance matrix of the underlying conditional error-correction regression, along with 95% confidence intervals.

Table.6.Long-Run Coefficients (HAC-corrected)

Variable	Coefficient	HAC Std. Error	t-Statistic	p-value	95% CI
LGOVEXP (Govt. Capital Expenditure)	−1.9430	0.2764	−7.0291	< 0.001	[−2.492, −1.394]
LTRADE (Trade Openness)	3.3401	1.1670	2.8623	0.005	[1.023, 5.658]
LBRENT (Oil Price — Brent)	0.3172	0.2274	1.3948	0.166	[−0.134, 0.769]
LCRDTOPRIVATESEC (Credit to Private Sector)	1.6287	0.7661	2.1259	0.036	[0.107, 3.150]

Note: HAC standard errors (Newey-West, 4-lag automatic bandwidth); long-run standard errors computed by the delta method on the full coefficient covariance matrix, not from the individual short-run coefficients in isolation.

Source: Author's own elaboration

The coefficient on LGOVEXP is −1.943 and remains highly significant under HAC correction ($t = -7.03$, $p < 0.001$), implying that a 1% increase in Government Capital Expenditure is associated with a 1.94% decrease in private non-oil GFCF in the long run, holding oil prices, credit availability, and trade openness constant. An elasticity greater than one in absolute value implies that the estimated association is more than proportional to the fiscal impulse — what the literature sometimes calls a super-elastic crowding-out pattern — though, as discussed in Section 3.3, this coefficient should be read as a robust conditional association rather than a fully identified causal effect, given the endogeneity concerns that a single-equation ARDL cannot rule out.

The coefficient on LTRADE is 3.340 and significant (t = 2.86, p = 0.005), so a 1% increase in trade openness is associated with a 3.34% increase in private non-oil GFCF in the long run, the largest positive elasticity in the model, though its 95% confidence interval, [1.02, 5.66], is wide, reflecting the collinearity discussed in Section 4.2, and the point estimate should not be read too precisely. The coefficient on LCRDTOPRIVATESEC is 1.629 and is statistically significant at the 5% level under HAC correction (t = 2.13, p = 0.036), an improvement in precision on the marginal 10% significance reported under the (mislabeled) standard errors in the version of this paper sent for review, consistent with a financial-accelerator channel in which credit availability supports private capital formation, albeit against a wide confidence interval [0.11, 3.15]. The coefficient on LBRENT remains statistically insignificant (0.317, p = 0.166). A null oil-price effect is not surprising for a dependent variable that is non-oil GFCF by construction; oil prices most plausibly affect this series indirectly, through the government budget and LGOVEXP, rather than directly.

4.5. Short-Run Dynamics of the ARDL Model

Table 7 reports the HAC-corrected short-run coefficients of the ARDL (7, 0, 3, 0, 0) conditional error-correction representation.

Table.7.Short-Run ARDL Coefficients and Error Correction (HAC-corrected)

Variable	Coefficient	HAC Std. Error	t-Statistic	p-value
ECT: LGFCNOIL(−1)	−0.3570	0.0999	−3.5730	< 0.001
LGOVEXP (short-run level)	−0.6937	0.2222	−3.1220	0.002
LTRADE(−1)	1.1925	0.4275	2.7895	0.005
LBRENT (short-run level)	0.1133	0.0831	1.3627	0.173
LCRDTOPRIVATESEC (short-run level)	0.5815	0.2501	2.3252	0.020
D(LGFCNOIL(−1))	0.0269	0.0213	1.2595	0.208
D(LGFCNOIL(−2))	0.0269	0.0195	1.3757	0.169
D(LGFCNOIL(−3))	0.1851	0.0743	2.4902	0.013
D(LGFCNOIL(−4))	0.2211	0.1045	2.1148	0.035
D(LGFCNOIL(−5))	0.2211	0.1240	1.7831	0.075
D(LGFCNOIL(−6))	−0.4370	0.2338	−1.8693	0.062
D(LTRADE)	2.3982	1.1590	2.0691	0.039
D(LTRADE(−1))	1.2057	1.1824	1.0197	0.308
D(LTRADE(−2))	−3.4336	2.8975	−1.1850	0.236
Constant	1.4493	2.5553	0.5672	0.571

***Note:** HAC (Newey-West) standard errors, 4-lag automatic bandwidth. ECT = Error Correction Term. Coefficients are identical to those reported in a prior draft (independently re-derived from the underlying data, as noted in Section 3.1); standard errors, t-statistics, and p-values are newly corrected.*

Source: Author's own elaboration

4.5.1. The Error Correction Term

The ECT coefficient is −0.357 and remains significant under HAC correction (t = −3.57, p < 0.001), within the theoretically required (−1, 0) range. This implies that about 35.7% of any deviation of private non-oil GFCF from its long-run path is closed within a month, for an implied half-life of disequilibrium of ln(0.5)/ln(1 − 0.357) ≈ 1.6 months. This is a fast adjustment speed, consistent with private non-oil investment responding quickly to re-align with government capital expenditure, trade openness, oil prices, and credit conditions once the long-run relationship has been disturbed. The short-run coefficient on LGOVEXP is −0.694 and significant ($t = -3.12$, $p = 0.002$): a 1% rise in government capital expenditure is associated with a 0.69% same-month decline in private non-oil GFCF. The short-run effect (−0.69) is smaller in magnitude than the long-run effect (−1.94), consistent with the displacement effect building up over time rather than being fully realized within the same month. Two lag terms that were reported as clearly significant under the originally mislabeled standard errors are only marginally significant under genuine HAC correction: D(LGFCNOIL(−5)) ($t = 1.78$, $p = 0.075$) and D(LGFCNOIL(−6)) ($t = -1.87$, $p = 0.062$), both now significant only at the 10% level rather than at 5% or 1%. D (LGFCNOIL (−3)) and D(LGFCNOIL(−4)) remain significant at 5%. Read cautiously, the pattern of positive coefficients at three-, four-, and five-month lags followed by a negative coefficient at six months is broadly consistent with a capital-goods installation lag — elevated activity in the months following a project's start, followed by mean reversion around six months out — but the weaker significance of the five- and six-month terms under correct standard errors means this interpretation should be held more tentatively than in the original submission.

The short-run dynamics of LTRADE change more substantively under HAC correction. The level term LTRADE (−1) is now significant ($t = 2.79$, $p = 0.005$), where it was not under the original standard errors, and the contemporaneous difference D (LTRADE) is now significant as well ($t = 2.07$, $p = 0.039$). The second lag, D (LTRADE(−2)), which was reported as marginally significant in the original submission ($p = 0.049$) and was interpreted there as evidence of a J-curve-type delayed adjustment to trade liberalization, is no longer significant under HAC correction ($t = -1.19$, $p = 0.236$). That specific short-run narrative is not supported once standard errors are accounted for, and we do not carry it forward into the discussion in Section 5; instead, the corrected results support a same-month positive association between trade openness and private non-oil GFCF, alongside the significant long-run relationship reported in Table 6.

4.6. Structural Stability

The CUSUM statistic stays within the 5% critical bounds for the full 2015–2025 recursive window, with a visible but contained rise around 2020–2021 that coincides with the COVID-19 shock before settling back toward zero. Since CUSUM does not breach its bounds, it does not, on its own, indicate a permanent break in the coefficient vector.

Table.8.Chow Breakpoint Tests

Candidate break date	Pre-break n	Post-break n	F-statistic	p-value	Decision
Vision 2030 launch (2016M04)	55	53	$F(15,78) = 0.765$	0.711	No significant break
COVID-19 shock (2020M03)	101	7	$F(15,78) = 10.090$	< 0.001	Significant break
National Investment Strategy (2021M01)	89	19	$F(15,78) = 1.412$	0.163	No significant break

Note: F-tests on the conditional error-correction regression (Table 7 specification); denominator degrees of freedom fixed at 78 (108 − 2×15) across candidate dates for comparability. The COVID-19 post-break sample is short (7 months), so that test result should be treated as indicative of instability around the break rather than as a precisely estimated post-break regime.

Source: Author's own elaboration

The Chow test finds no significant break at the Vision 2030 announcement or at the National Investment Strategy launch, which supports treating the estimated relationship as stable across the main Vision 2030 policy milestones the events most central to this paper's motivation. It does find a highly significant break around the COVID-19 shock ($F = 10.09$, $p < 0.001$). This is a more decisive statement than the CUSUM plot alone suggests, and the two results are not strictly contradictory: CUSUM is a cumulative test that can fail to flag a short, sharp disruption that is followed by a return to the prior relationship, which is a plausible description of the pandemic period, while the Chow test is a direct test of whether the coefficient vector before and after a specific date differs, and seven months is enough to detect a large, temporary dislocation without necessarily indicating a permanent change in the long-run relationship. Given the short post-break window underlying the COVID test, we read this as evidence of a temporary disruption around the pandemic rather than an ongoing structural change, but it does qualify the stability claim in the original submission, and we no longer describe the relationship as stable "throughout the entire sample" without this caveat.

4.7. Impulse Response Functions

Figure 2 plots the orthogonalized point-estimate impulse responses of LGFCNOIL to a one-standard-deviation shock in each of LGOVEXP, LBRENT, LTRADE, and LCRDTOPRIVATESEC over a five-month horizon, estimated from a state-space VAR (2) that handles the missing LGFCNOIL observations (Section 3.1) through Kalman filtering on the full 175-month calendar index, rather than by deleting rows, which would otherwise misalign lags across the gaps. We report point estimates only: analytically valid confidence bands for this state-space specification, with its irregular missing-data pattern, would require a bootstrap procedure beyond the scope of this revision, so Figure 1 should be read as descriptive of the estimated dynamic path rather than as a set of significance tests in its own right. Section 4.8 below, using the pairwise Granger and Block Exogeneity tests, remains the basis for this paper's short-run causality conclusions.

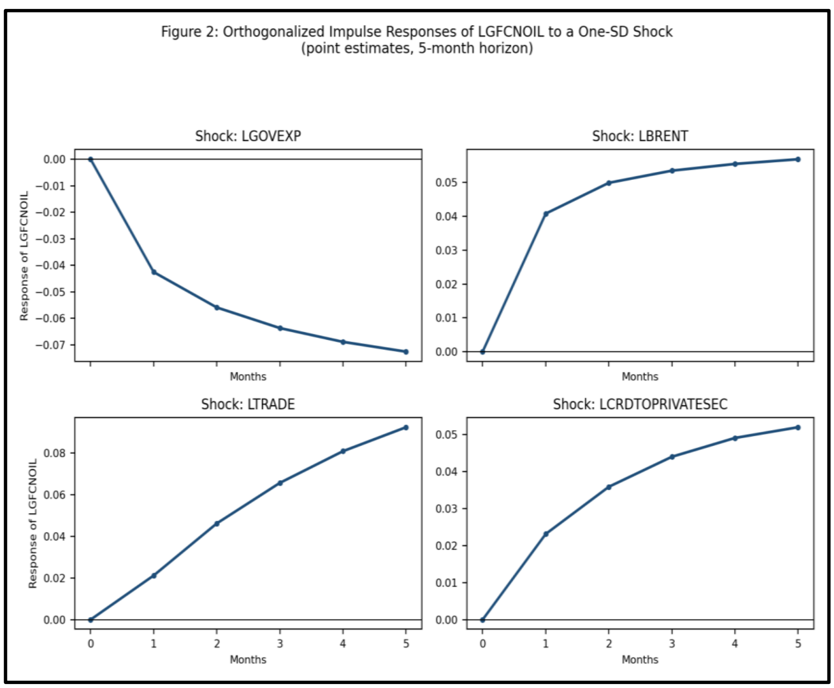

Figure. 1. Point-estimate orthogonalized impulse responses
Source: Author's own elaboration

The estimated response of LGFCNOIL to an LGOVEXP shock is negative throughout the five-month horizon and grows modestly in magnitude (from −0.043 at one month to −0.073 at five months), consistent in sign with the negative long-run and short-run coefficients reported above. The responses to LBRENT, LTRADE, and LCRDTOPRIVATESEC shocks are all positive and of broadly similar, modest magnitude over this horizon. Because we do not report confidence bands for this figure, we do not draw a statistical-significance conclusion from it; the pattern is offered as a descriptive complement to Section 4.8, where the pairwise Granger and Block Exogeneity tests that explains a formal significance levels and find no significant short-run predictive causality between LGOVEXP and LGFCNOIL over the same two-lag window.

4.8. Granger Causality and VAR Block Exogeneity

Table 9 reports pairwise Granger causality tests at two lags for all ten variable pairs.

Table.9.Pairwise Granger Causality Tests

Null hypothesis	Obs.	F-Statistic	p-value
LGOVEXP does not Granger-cause LGFCNOIL	134	0.191	0.826
LGFCNOIL does not Granger-cause LGOVEXP	134	0.033	0.968
LTRADE does not Granger-cause LGFCNOIL	134	0.029	0.972
LGFCNOIL does not Granger-cause LTRADE	134	0.273	0.762
LBRENT does not Granger-cause LGFCNOIL	134	0.089	0.915
LGFCNOIL does not Granger-cause LBRENT	134	0.248	0.781
LCRDTOPRIVATESEC does not Granger-cause LGFCNOIL	134	0.007	0.993
LGFCNOIL does not Granger-cause LCRDTOPRIVATESEC	134	0.185	0.831
LTRADE does not Granger-cause LGOVEXP	173	1.695	0.187
LGOVEXP does not Granger-cause LTRADE	173	0.752	0.473
LBRENT does not Granger-cause LGOVEXP	173	1.997	0.139
LGOVEXP does not Granger-cause LBRENT	173	0.892	0.412
LCRDTOPRIVATESEC does not Granger-cause LGOVEXP	173	1.092	0.338
LGOVEXP does not Granger-cause LCRDTOPRIVATESEC	173	0.693	0.502
LBRENT does not Granger-cause LTRADE	173	0.112	0.894
LTRADE does not Granger-cause LBRENT	173	0.333	0.717
LCRDTOPRIVATESEC does not Granger-cause LTRADE	173	1.200	0.304
LTRADE does not Granger-cause LCRDTOPRIVATESEC	173	0.043	0.958
LCRDTOPRIVATESEC does not Granger-cause LBRENT	173	0.003	0.997
LBRENT does not Granger-cause LCRDTOPRIVATESEC	173	0.008	0.992

Note: observation counts follow the reconciliation in Section 3.1 Table 2. As a robustness check (untabulated), all pairs involving LGOVEXP and LGFCNOIL were re-estimated at a single lag; the conclusion of no significant causality in either direction is unchanged (Section 3.6).

Source: Author's own elaboration

No pair shows significant Granger causality at conventional levels. In particular, neither LGOVEXP nor LGFCNOIL Granger-causes the other, and the reverse is rejected (p = 0.826 and p = 0.968 respectively), which is the basis for describing the LGOVEXP-LGFCNOIL relationship as an equilibrium, long-run one rather than a short-run predictive one, and is one of the pieces of evidence bearing on the endogeneity discussion in Section 3.3.

Table 10 reports the VAR Block Exogeneity Wald tests, which assess the joint significance of each variable's own lags in the other variables' equations within the 5-variable VAR system.

Table.10.VAR Granger Causality / Block Exogeneity Wald Tests

Dependent variable	Excluded	Chi-sq	Df	p-value
LGFCNOIL	LGOVEXP	2.952	2	0.229
LGFCNOIL	LTRADE	2.744	2	0.254
LGFCNOIL	LBRENT	0.143	2	0.931
LGFCNOIL	LCRDTOPRIVATESEC	2.000	2	0.368
LGFCNOIL	All	4.075	8	0.850
LGOVEXP	LGFCNOIL	0.533	2	0.766
LGOVEXP	LTRADE	0.011	2	0.995
LGOVEXP	LBRENT	5.231	2	0.073
LGOVEXP	LCRDTOPRIVATESEC	0.035	2	0.983
LGOVEXP	All	6.499	8	0.592
LTRADE	LGFCNOIL	1.214	2	0.545
LTRADE	LGOVEXP	1.502	2	0.472
LTRADE	LBRENT	2.240	2	0.326
LTRADE	LCRDTOPRIVATESEC	0.963	2	0.618
LTRADE	All	5.516	8	0.701
LBRENT	LGFCNOIL	0.923	2	0.630
LBRENT	LGOVEXP	2.025	2	0.363
LBRENT	LTRADE	0.432	2	0.806
LBRENT	LCRDTOPRIVATESEC	0.889	2	0.641
LBRENT	All	5.533	8	0.699
LCRDTOPRIVATESEC	LGFCNOIL	0.714	2	0.700
LCRDTOPRIVATESEC	LGOVEXP	0.882	2	0.644
LCRDTOPRIVATESEC	LTRADE	0.156	2	0.925
LCRDTOPRIVATESEC	LBRENT	0.267	2	0.875
LCRDTOPRIVATESEC	All	1.461	8	0.993

Note: Sample: 2011M03–2025M09, 134 included observations (complete-case VAR at 2 lags).

Source: Author's own elaboration

The Block Exogeneity results confirm the pairwise findings: no variable is jointly significant in explaining another at the 5% level, with one marginal exception LBRENT in the LGOVEXP equation ($\chi^2 = 5.23$, $p = 0.073$), significant only at the 10% level, consistent with oil prices feeding into the government budget and capital expenditure decisions with a short lag, which fits the mechanism described in Section 2.

5. Discussion

This section interprets the results reported in Section 4 against the literature reviewed in Section 2, rather than restating the coefficients again; readers looking for the numerical results should refer back to Tables 5–10.

5.1. Government Capital Expenditure and the Crowding-Out Hypothesis

The negative long-run elasticity of private non-oil GFCF with respect to Government Capital Expenditure (−1.94) is directionally consistent with the neoclassical crowding-out channel described in Section 2 and with the cross-country evidence in Furceri and Sousa (2011) and Carvelli (2023), both of which find that government spending displaces rather than supports private investment on average, with the effect concentrated in the non-productive component of spending in Carvelli's decomposition. The Saudi elasticity is considerably larger in magnitude than typical cross-country estimates, which could reflect a genuine feature of Saudi Arabia's fiscal structure government's outsized role relative to a comparatively small private non-oil sector, and preferential government access to financing and land that Section 2 identifies as a resource-channel specific to rentier states — but could also, at least in part, reflect the collinearity among the annual-level regressors documented in Section 4.2 and the endogeneity concerns set out in Section 3.3, both of which this paper cannot fully rule out. We therefore treat the magnitude, though not the direction, of the estimate with some caution.

The result is not aligned with Jalles, Park, and Qureshi (2024), who find that public investment tends to crowd in private investment in emerging and developing economies on average, conditional on fiscal space and institutional quality. One reading, consistent with their conditioning variables, is that Saudi Arabia's rapid, front-loaded Vision 2030 investment pace and a non-oil primary deficit that reached 33.0% of non-oil GDP in 2023 (IMF estimate) place it closer to the limited-fiscal-space case in which their framework would also predict crowding out rather than crowding in.

The absence of short-run Granger causality between LGOVEXP and LGFCNOIL (Section 4.8) alongside a confirmed long-run cointegrating relationship (Section 4.3) is consistent with Rentier State Theory's description of the Saudi private sector as structurally oriented around government fiscal choices (Mahdavy, 1970; Beblawi & Luciani, 1987): the two series move together at the level of a stable long-run relationship without one predicting month-to-month movements in the other, which is a different empirical signature than either a tightly synchronized short-run relationship or genuine independence would produce.

5.2. Trade Openness and Private-Sector Credit

Trade openness carries the largest positive long-run elasticity in the model (3.34), consistent with the accelerator and openness-investment channels discussed in Section 2 and with Onifade, Khatir, Ay, and Canitez (2022), who identify trade openness as a channel through which MENA economies gain access to imported capital goods and larger export markets. Credit to the private sector is significant at the 5% level under HAC correction (1.63, $p = 0.036$, an improvement in precision over the marginal 10% result reported under the original, mislabeled standard errors), broadly consistent with the financial-accelerator mechanism in Bernanke, Gertler, and Gilchrist (1999) and with the aims of Saudi Arabia's Financial Sector Development Program. Both coefficients, however, come from a set of regressors with VIFs as high as 27 (Section 4.2), so their relative magnitudes trade openness appearing to matter more than credit access should be read as suggestive rather than as a precise ranking of channels.

5.3. The Insignificance of Oil Prices

Brent crude oil prices are statistically insignificant in both the long run and the short run. This null result is consistent with a growing body of evidence that Saudi Arabia's non-oil private sector has become less directly sensitive to oil-price swings as diversification has progressed (IMF, 2024), and it is consistent with the theoretical channel set out in Section 2, in which oil prices are expected to affect private non-oil investment mainly indirectly, via their effect on government revenue and, in turn, on LGOVEXP, rather than directly. The marginally significant Block Exogeneity result for LBRENT in the LGOVEXP equation (Table 10, p = 0.073) is consistent with this indirect channel operating with a short lag, even though oil prices do not enter the private non-oil GFCF equation directly.

5.4. An Equilibrium Relationship, Not a Sequential One

Taken together, Sections 4.3–4.8 indicate that the relationship between Government Capital Expenditure and private non-oil GFCF is primarily characterized by long-run equilibrium rather than short-run predictive dynamics. The bounds-testing results support cointegration, although the evidence is more moderate under HAC correction than suggested by the conventional OLS-based F-statistic alone. The error-correction coefficient indicates relatively rapid adjustment, with approximately 35.7% of deviations from long-run equilibrium corrected each month. At the same time, neither the pairwise Granger causality tests nor the VAR Block Exogeneity Wald tests identify significant short-run predictive causality between LGOVEXP and LGFCNOIL. This pattern is consistent with the possibility that their long-run association operates through slower-moving mechanisms, including the sectoral composition of public investment, land and permitting arrangements, and financing conditions. Moreau and Aligishiev’s (2024) general-equilibrium analysis of the National Investment Strategy similarly suggests that the growth effects of Vision 2030 investment depend on complementary institutional, public-sector efficiency, and labour-market reforms rather than arising mechanically from higher public investment. Moreover, the limited temporal variation in LGOVEXP given that its underlying source data vary primarily at an annual frequency reduces the ability of a two-lag monthly VAR to detect short-run predictive effects. Thus, the absence of short-run Granger causality should not be interpreted as evidence that no economic interaction exists; rather, it suggests that the observed relationship is not captured through statistically significant short-lag predictive dynamics.

5.5. Policy Implications

If the association between Government Capital Expenditure and private non-oil GFCF reported here reflects displacement rather than complementarity, that would support reweighting some public capital spending toward enabling infrastructure, regulatory streamlining, and co-investment vehicles that lower the cost of private capital rather than toward direct state-led capital accumulation a direction the IMF (2024, 2025) and World Bank (2024) have both suggested on separate grounds, and one PwC Middle East (2026) also frames in terms of dependency risk: public spending that is not matched by private co-investment can reduce firms' incentive to compete and improve efficiency on their own. Given the endogeneity and data-construction caveats set out in Sections 3.1 and 3.3, we frame this as a hypothesis the results are consistent with rather than as a demonstrated causal policy effect; Section 6.2 states the policy recommendations that follow with that qualification attached.

6. Conclusion

6.1. Summary of Findings

This study examines the relationship between Government Capital Expenditure and Private Non-Oil Gross Fixed Capital Formation in Saudi Arabia over March 2011–September 2025 using ARDL bounds testing with HAC-corrected standard errors. Four results stand out. First, the two series are cointegrated (Section 4.3), though the strength of that evidence is more moderate under HAC correction than a reading of the uncorrected F-statistic alone would suggest. Second, Government Capital Expenditure carries a negative and statistically significant long-run elasticity (−1.94, HAC se = 0.28, $p < 0.001$) on private non-oil GFCF, consistent with crowding out rather than crowding in over this period, subject to the endogeneity caveat in Section 3.3. Third, trade openness and credit to the private sector both carry positive, statistically significant long-run elasticities, while Brent oil prices do not. Fourth, no pairwise Granger causality or VAR Block Exogeneity test finds significant short-run predictive causality between Government Capital Expenditure and private non-oil GFCF, consistent with a long-run structural relationship rather than a month-to-month one, and a Chow test finds no significant parameter break around the Vision 2030 or National Investment Strategy launches, though it does find one around the COVID-19 shock.

6.2. Policy Recommendations

Three recommendations follow from these results, offered with the criterion that the underlying elasticities are conditional associations rather than fully identified causal effects (Section 3.3).

1. **Rebalance the composition of capital spending**

Where public capital spending competes directly with private firms for financing, land, and contracting capacity, shifting a larger share of the capital budget toward enabling infrastructure and regulatory streamlining — areas with fewer private-sector substitutes could reduce the displacement channel identified here.

2. **Use co-investment and financing structures that build in private participation.**

PIF and Ministry of Investment vehicles that require private co-financing or that de-risk private participation directly address the financing-competition channel, rather than leaving private firms to compete with the state for the same pool of capital.

3. **Support the credit and trade channels that have already been shown to matter.**

Because credit to the private sector and trade openness both carry significant positive long-run associations with private non-oil GFCF, continued implementation of the Financial Sector Development Program and further trade facilitation are likely to support private capital formation independent of the fiscal-composition question above.

6.3. Limitations and Future Research

Although the model is estimated at a monthly frequency, the short-run coefficients on regressors derived from annual-frequency data should not be interpreted as evidence of genuine within-year policy dynamics. Moreover, the single-equation ARDL framework does not fully resolve potential reverse causality and simultaneity between Government

Capital Expenditure and private investment. Future research could address this identification issue using instrumental-variable or GMM approaches, potentially employing exogenous oil-revenue windfalls or other credible instruments for Government Capital Expenditure.

The effective sample of 108 observations for the ARDL bounds test is relatively modest given the model's 15 short-run parameters and is further constrained by lag selection and missing observations. In addition, the post-COVID segment used in the Chow breakpoint analysis contains only seven months of observations, limiting the statistical power of structural-break inference. As additional observations become available beyond 2025, re-estimation using an extended sample would permit more robust assessment of structural stability and regime changes.

The analysis uses broad aggregate measures of GFCF, Government Capital Expenditure, and private-sector credit and therefore does not capture heterogeneity across sectors or project types. Disaggregating investment by sectors such as housing, transport, tourism, and industry, as well as by productive and non-productive expenditure, could determine whether the observed crowding-out effect is concentrated in particular categories of public investment.

The study assumes a linear long-run relationship and does not examine potential asymmetries between increases and decreases in Government Capital Expenditure. Nonlinear or asymmetric approaches, such as those employed by Baussola and Carvelli (2023) and Carvelli (2024), could provide additional insight, particularly given Saudi Arabia's fiscal adjustments surrounding the 2015–2016 oil-price shock.

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