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PANEL GMM ANALYSIS OF THE IMPACT OF FACTORS ON FISCAL DEFICIT IN CENTRAL ASIA AND THE CAUCASUS COUNTRIES

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ABSTRACT. The article investigates the dynamic interplay between public debt, fiscal deficits, and institutional quality across eight Central Asian and Caucasus economies (Kazakhstan, Kyrgyzstan, Tajikistan, Turkmenistan, Uzbekistan, Armenia, Azerbaijan, and Georgia) over the period 2000-2024. Motivated by persistent endogeneity concerns in the debt-growth literature – including reverse causality, omitted variable bias, and the dynamic panel bias identified by Nickell – the study employs a two-step System Generalized Method of Moments (GMM) estimator with the Windmeijer finite-sample correction. The empirical specification models fiscal deficit dynamics as a function of lagged fiscal balances, public debt-to-GDP, institutional quality (drawn from the Worldwide Governance Indicators), and a battery of macroeconomic controls. Baseline results indicate that the lagged fiscal deficit coefficient is statistically insignificant and unstable across specifications, while the public debt coefficient is positive but imprecisely estimated. Subperiod analysis reveals temporal instability in the debt-deficit relationship, shifting from negative (pre-2008) to positive (post-2008). Diagnostic tests reveal a persistent overidentification problem (Hansen J p-value of 1.000 in all specifications) attributable to the unusually small cross-sectional dimension ($N = 8$) relative to the time dimension ($T = 25$) – a panel structure for which System GMM is not optimally suited. The paper concludes that institutional reforms are likely complements to, rather than substitutes for, numerical fiscal rules, while cautioning that the present findings should be treated as suggestive pending validation on a larger country sample or an estimator better matched to this panel's dimensions.

Keywords: public debt; fiscal deficit; institutional quality; System GMM; Central Asia; Caucasus economies **JEL Classification:** C23, H62, H63, O43

INTRODUCTION

The interplay between public debt accumulation and long-term economic growth remains one of the most persistently contentious debates in modern macroeconomics, particularly in the aftermath of the global financial crisis and the subsequent sovereign debt episodes in advanced and emerging economies. The theoretical foundations established by Krugman regarding the debt overhang hypothesis posited that excessive public indebtedness discourages investment, as future tax liabilities are capitalised into current output losses [1]. Subsequent empirical inquiries, most notably the seminal work of Reinhart and Rogoff, reignited policy discourse by suggesting a sharp debt-to-GDP threshold beyond which growth trajectories exhibit a pronounced deceleration [2]. However, these foundational contributions, while instrumental in shaping the fiscal consolidation narrative, largely relied on simple correlations and pooled cross-sectional frameworks that presuppose homogeneous behavioural responses across heterogeneous economies [3]. Contemporary research has progressively shifted towards understanding that the growth consequences of public debt are not immutable but are profoundly mediated by structural characteristics, particularly the quality of domestic governance, fiscal institutions, and the counter-cyclical capacity of budgetary frameworks [4], [5], [6].

Despite the substantial volume of empirical work on the debt-growth nexus, a critical and persistent lacuna persists in the literature regarding the causal identification of this relationship, specifically concerning the endogenous feedback loops between sovereign indebtedness, institutional frailties, and fiscal performance. Early panel studies, employing standard fixed-effects (FE) or pooled OLS estimators, suffer from a well-documented simultaneity bias: while high debt may stifle growth, economic recessions simultaneously deteriorate fiscal balances and inflate debt-to-GDP ratios, creating a bidirectional causality that renders OLS coefficients inconsistent [7]. Furthermore, as highlighted by Chudik et al. [8] and Kourtellis et al. [9], the omission of unobserved country-specific heterogeneity and the assumption of cross-sectional independence in traditional FE models systematically

overestimate the marginal effect of debt, while failing to account for the partial adjustment dynamics inherent in fiscal policy – where current public debt is intrinsically a function of its own past values [10]. Moreover, recent contributions by Aslan and Altinoz [11] and Efobi [12] demonstrate that institutional quality (e.g., control of corruption, bureaucratic efficiency) is not strictly exogenous; rather, it is endogenously determined by historical fiscal trajectories and political settlements, implying that any empirical strategy neglecting this endogeneity will produce severely attenuated or spurious estimates for the moderating role of governance.

To address this gap empirically, this study draws on a panel of eight transition economies from Central Asia and the Caucasus – Kazakhstan, Kyrgyzstan, Tajikistan, Turkmenistan, and Uzbekistan (Central Asia), together with Armenia, Azerbaijan, and Georgia (the Caucasus) – observed annually over 2000-2024. These economies share comparable Soviet institutional legacies, broadly similar transition trajectories, and sufficiently heterogeneous fiscal and governance outcomes to support a meaningful empirical test of the debt-deficit-institution nexus, while also representing a region substantially underrepresented in the existing fiscal sustainability literature relative to Latin America, sub-Saharan Africa, and Eastern Europe.

Our contribution to this discourse is threefold and explicitly designed to address the aforementioned econometric pitfalls through a robust dynamic panel methodology. First, we formally account for the persistence and inertia of fiscal variables by incorporating the lagged dependent variable (debt-to-GDP or deficit ratio) into our specification, thereby modelling the gradual adjustment of fiscal positions towards long-term equilibrium, a feature often neglected in static institutional regressions (Bhanumurthy et al. [13], Combes et al., [14]). Second, we confront the pervasive issue of endogeneity stemming from reverse causality and time-varying omitted variables (e.g., terms-of-trade shocks or global liquidity cycles) by employing the System Generalized Method of Moments (GMM) estimator developed by Arellano and Bover [15] and Blundell and Bond [16]. Third, and most crucially, we address the interactive effect between institutional quality and fiscal rules in shaping the sovereign debt trajectory, providing empirical evidence on whether strong governance frameworks mitigate the pro-cyclical bias of fiscal policy during periods of high indebtedness (Gootjes & de Haan [17]; Asatryan et al. [18]).

LITERATURE REVIEW

The theoretical underpinnings of the public debt-growth nexus are historically anchored in the neoclassical growth framework, yet the seminal contribution of Krugman fundamentally reoriented the debate through the lens of the debt overhang hypothesis [1]. This hypothesis posits that when a country's external or public debt exceeds its repayment capacity, the expected future tax burden discourages private investment, effectively creating a "hysteresis trap" where output growth is penalised irrespective of the current fiscal stance. Subsequent extensions by Aizenman, Kletzer, and Pinto formalised this interaction by introducing tax revenue constraints into an intertemporal optimisation framework [19]. Concurrently, the theoretical literature on fiscal deficits, pioneered by Alesina and Tabellini [20], introduced political economy considerations, arguing that partisan conflicts over the distribution of tax burdens systematically generate persistent deficit biases – a proposition later empirically validated by Woo [21].

From a purely macroeconomic perspective, the theoretical model of fiscal sustainability is rooted in the government's intertemporal budget constraint. Following the sustainability analysis approach of Celasun, Debrun, and Ostry, the fiscal adjustment path is contingent upon the primary surplus-to-GDP ratio responding positively to increases in the public debt-to-GDP ratio [22]. However, the theoretical work by Aghion, Hemous, and Kharroubi introduced a critical caveat: in the presence of liquidity constraints, pro-cyclical fiscal adjustments during downturns may amplify output volatility, thereby undermining the long-run solvency conditions [23].

The empirical literature on public debt and fiscal sustainability has undergone a significant methodological evolution over the past two decades. Early cross-sectional studies, most notably Reinhart and Rogoff, documented an inverse relationship between debt-to-GDP ratios exceeding the 90% threshold and subsequent GDP growth, utilising simple non-parametric correlations and pooled OLS regressions [2]. However, these findings were sharply contested by Panizza and Presbitero, who demonstrated that the debt-growth correlation evaporates once country-specific fixed effects are introduced [3]. Extending this critique, Eberhardt and Presbitero employed heterogeneous panel estimators that account for cross-sectional dependence and parameter heterogeneity, conclusively showing that the debt threshold effect is neither universal nor stable across countries [7].

Crucially, the most pervasive threat to the internal validity of extant studies is the endogeneity of public debt and fiscal deficits, which manifests through three distinct channels. First, simultaneity (reverse causality) operates where economic contractions simultaneously increase debt-to-GDP ratios and reduce growth, generating a negative feedback loop that biases OLS coefficients downward (Chudik et al. [8]). Second, measurement error in institutional quality indices – such as the Worldwide Governance Indicators (WGI) – contaminates the estimated coefficients, as these proxies are subjective perceptions aggregated from multiple sources, introducing classical errors-in-variables that bias FE estimates towards zero (Gründler & Potrafke [24]). Third, omitted variable bias

arises from time-varying confounders that jointly determine debt trajectories and growth outcomes (Kourtellos, Stengos, & Tan [9]).

A rapidly expanding strand of literature has shifted focus from the direct debt-growth nexus to the institutional and governance transmission channels that mediate this relationship. The theoretical contribution of Acemoglu, Johnson, and Robinson (2001) established that colonial origins and institutional endowments fundamentally determine long-run developmental trajectories [4]. Empirically, Alesina et al. documented that budget institutions – encompassing transparency, hierarchical procedures, and fiscal rules – significantly improve fiscal performance in Latin American economies by constraining the executive's discretion over deficit spending [25]. Extending this geographical scope, Alt and Lassen [26] demonstrated that fiscal transparency reduces debt accumulation in OECD countries, while Asatryan, De Witte, and Havlik [18] provided robust evidence that institutional quality at the local government level moderates the relationship between decentralisation and fiscal discipline.

With respect to numerical fiscal rules, the evidence remains mixed but increasingly sophisticated. Calderón, Duncan, and Schmidt-Hebbel [27] showed that fiscal rules foster fiscal discipline in developing economies only when accompanied by strong institutional enforcement mechanisms, a finding corroborated by Gootjes and de Haan [17]. Badinger and Reuter extend this literature by constructing a comprehensive fiscal rules strength index across a large country panel, demonstrating that more stringent numerical rules are associated with smaller cyclical deficits, but only when complemented by independent monitoring bodies – reinforcing the complementarity hypothesis later tested in our empirical model [28]. Furthermore, Heinemann, Osterloh, and Kalb demonstrated that fiscal rules reduce sovereign risk premia through a credibility channel [29]. Recent contributions by Duarte and de Mello [30] and Saadaoui and Boukhatem [31] have advanced this literature by integrating institutional quality as a threshold variable in non-linear dynamic panel specifications.

However, a critical lacuna persists: the interaction between fiscal deficits and institutional quality has been predominantly examined in isolation, with studies treating the cyclicity of deficits as an outcome of institutional constraints (Fatás & Mihov [5]; Vegh & Vuletin [32]) rather than as a dynamic process that simultaneously co-evolves with debt accumulation and governance reforms. Moreover, recent dynamic panel studies by Aslan and Altinoz [11] and Canh, Binh, and Thanh [33] have underscored the necessity of treating institutional quality as endogenously determined by fiscal performance, arguing that system-GMM estimators are uniquely suited to disentangle the bidirectional causality between governance and fiscal discipline – a methodological choice we adopt below, while acknowledging important caveats regarding our panel's dimensions (Section 3.6).

Building upon the theoretical foundations of the fiscal sustainability framework (Krugman [1]; Celasun et al. [22]) and the institutional transmission literature (Acemoglu et al. [4]; Asatryan et al. [18]), we formalise the empirical relationship through a dynamic panel specification. Our baseline model is given by:

$$Y_{it} = \alpha \cdot Y_{i,t-1} + \beta_1 \cdot Debt_{it} + \beta_2 \cdot Deficits_{it} + \beta_3 \cdot Inst_{it} + \beta_4 \cdot (Debt_{it} \times Inst_{it}) + \gamma' X_{it} + \mu_i + \varepsilon_{it}$$

where Y_{it} represents the fiscal sustainability indicator (alternatively, the debt-to-GDP ratio or the primary deficit-to-GDP ratio) for country i at time t ; $Y_{i,t-1}$ captures the inherent inertia and partial adjustment of fiscal positions; $Debt_{it}$ denotes public debt-to-GDP; $Deficits_{it}$ measures the fiscal balance; $Inst_{it}$ captures institutional quality (composite index of governance, rescaled 0-100); and X_{it} is a vector of control variables. μ_i represents unobserved country-specific effects, and ε_{it} is the idiosyncratic error term.

Drawing upon the synthesised theoretical and empirical evidence, we articulate three testable hypotheses:

1. H1: Public debt exerts a non-linear and conditional negative effect on fiscal sustainability, but this contractionary effect is significantly attenuated in countries with higher institutional quality. We anticipate a positive coefficient on the interaction term ($Debt \times Inst$).

2. H2: Fiscal deficits exhibit higher persistence and stronger pro-cyclicality in countries with weak bureaucratic quality and limited control of corruption. The coefficient on the lagged dependent variable α is hypothesised to be positive and significantly greater in low-institution subsamples.

2. H3: The effectiveness of fiscal rules and numerical deficit targets is conditional upon the underlying institutional framework, such that rules only foster fiscal discipline when governance quality exceeds a critical threshold.

RESEARCH METHODOLOGY

Building upon the theoretical framework delineated in Section 2.4, we specify a dynamic panel data model that captures the inherent persistence of fiscal variables and the institutional transmission channels. Our baseline empirical model takes the following form:

$$FS_{it} = \alpha \cdot FS_{i,t-1} + \beta^1 \cdot Debt_{it} + \beta^2 \cdot Deficit_{it} + \beta^3 \cdot Inst_{it} + \beta^4 \cdot (Debt_{it} \times Inst_{it}) + \gamma' Z_{it} + \mu_i + \varepsilon_{it}$$

where FS_{it} denotes the fiscal sustainability indicator for country i ($i = 1, \dots, N$) at time t ($t = 1, \dots, T$); $FS_{i,t-1}$ captures the partial adjustment dynamics and policy inertia inherent in fiscal decision-making; $Debt_{it}$ is the stock of public debt-to-GDP ratio; $Deficit_{it}$ represents the fiscal balance; $Inst_{it}$ is a composite index of institutional quality drawn from the Worldwide Governance Indicators (WGI), rescaled to a 0-100 metric; and $(Debt_{it} \times Inst_{it})$ is the interaction term designed to test Hypothesis 1. The vector Z_{it} encompasses real GDP growth, inflation, trade openness, and terms-of-trade shocks. The parameter α represents the coefficient on the lagged dependent variable, μ_i denotes unobserved country-specific fixed effects, and ε_{it} is the idiosyncratic error term.

Standard panel estimators, including the within-groups (fixed effects) and pooled OLS estimators, yield inconsistent estimates in dynamic panels due to the correlation between the lagged dependent variable and the country-specific effect when the data are within-transformed. The asymptotic bias of the fixed effects estimator in dynamic panels with fixed T is given by:

$$Plim_n \rightarrow \infty (\hat{\alpha}_{FE} - \alpha) = -\frac{1 + \alpha}{T - 1} \cdot \left[1 - \left(\frac{1}{T} \right) \cdot \frac{1 - \alpha^T}{1 - \alpha} \right]^{-1}$$

which, for moderate values of α (e.g., $\alpha = 0.7$) and time dimension $T = 20$, approximates to a downward bias of nearly 9 percentage points. In panels where fiscal variables exhibit high persistence (α close to 0.9), this bias becomes even more pronounced, motivating the adoption of instrumental variable strategies specifically designed for dynamic panels.

The Generalized Method of Moments (GMM) framework for dynamic panels was pioneered by Arellano and Bond, who proposed the Difference GMM estimator, transforming the data through first-differencing to eliminate the country-specific fixed effects [34]. However, as demonstrated by Blundell and Bond, the Difference GMM estimator suffers from weak instruments and substantial finite-sample bias when the dependent variable exhibits high persistence and the variance of the fixed effects is large relative to the variance of the idiosyncratic errors [16]. The System GMM estimator, which combines the difference equations with level equations instrumented by lagged differences, yields superior finite-sample properties under these conditions (table 1).

Table 1
Summaries of the comparative properties of the Difference GMM and System GMM estimators¹

No	Criterion	Difference GMM	System GMM
1.	Equation system	First-differenced equations only	First-differenced + Level equations
2.	Instruments for differenced eq.	Lagged levels ($t-2$, $t-3$, ...)	Lagged levels ($t-2$, $t-3$, ...)
3.	Instruments for level eq.	Not applicable	Lagged differences ($\Delta t-1$, $\Delta t-2$, ...)
4.	Persistence handling	Weak instruments when $\alpha \rightarrow 1$	Robust to high persistence
5.	Efficiency (finite-sample)	Less efficient (larger variance)	More efficient (smaller variance)
6.	Bias when $\alpha > 0.8$	Substantial (upward bias)	Near-unbiased
7.	Optimal panel dimension	Large N , small/moderate T	Large N , small/moderate T

Table 1 adapted from Blundell and Bond [16] and Roodman [35]. Given that our fiscal sustainability indicators exhibit pronounced persistence – a feature theoretically rationalised by expenditure rigidities and path-dependent tax structures (Alesina & Tabellini [20]; Fatás & Mihov [5]) – the System GMM estimator is theoretically preferred for the persistence problem. We adopt the two-step System GMM estimator with the Windmeijer [36] finite-sample correction, implemented via the `xtabond2` routine (Roodman [35]).

The validity of the GMM estimates hinges critically upon the exogeneity classifications assigned to each regressor. We treat public debt, fiscal deficits, and institutional quality as endogenous, given their joint determination with growth and fiscal outcomes (Panizza & Presbitero [3]; Aslan & Altinoz [11]; Gründler & Potrafke [24]). Macroeconomic controls (GDP growth, inflation, trade openness) are treated as predetermined. For endogenous variables, we employ instruments constructed from their lagged levels at $t-2$ and deeper for the difference equation, and lagged first differences for the level equation, consistent with the Blundell-Bond orthogonality conditions.

A critical concern in GMM applications is the proliferation of instruments (Roodman [35]). As the number of instruments approaches or exceeds the number of cross-sectional units (N), the Hansen J-test becomes severely over-identified, producing p-values arbitrarily close to unity. To mitigate this, we implement the instrument collapsing technique (the collapse option in `xtabond2`) and restrict the lag depth for endogenous regressors to $p = 3$ or $p = 4$.

¹ Compiled by the author.

Following estimation, we implement a battery of diagnostic tests. The Arellano-Bond AR(1) test is expected to reject the null of no first-order autocorrelation by construction; the critical AR(2) test must not reject the null ($p > 0.10$) for the instruments to be considered valid. The Hansen J-statistic tests the joint validity of the instrument set; the optimal p-value range is between 0.10 and 0.30, with values approaching 1.000 indicating instrument proliferation rather than genuine instrument validity (Roodman [35]). The Difference-in-Hansen test examines the exogeneity of instrument subsets, including the Blundell-Bond level-equation orthogonality conditions.

It is important to flag explicitly, ahead of the empirical results, a structural tension between our chosen estimator and the dimensions of our panel. The System GMM estimator of Arellano and Bover [15] and Blundell and Bond [16] was developed for, and derives its asymptotic properties from, panels in which the number of cross-sectional units (N) is large relative to the number of time periods (T) – the canonical application being micro-panels of firms or households with N in the hundreds or thousands and T typically below 10. Our panel inverts this configuration: N = 8 countries observed over T = 25 years. Under these conditions, the instrument count generated by even modest lag depths ($p = 3$ or 4) can easily approach or exceed N, producing exactly the instrument proliferation symptoms documented throughout Sections 5 and 6 of this paper – most visibly the Hansen J p-value of 1.000 in every single specification we estimate, and Cragg-Donald first-stage F-statistics below the Stock-Yogo [37] critical value of 10 for every endogenous regressor.

We retain System GMM as our primary estimator in what follows because it remains the dominant approach in the comparable literature (Bhanumurthy et al. [13]; Aslan & Altinoz [11]; Canh et al. [33]) and because our robustness checks (Section 6.1) explicitly bound the GMM estimates against pooled OLS and fixed-effects benchmarks. However, we caution the reader – and recommend to future researchers extending this work – that estimators explicitly designed for $N < T$ or $N \approx T$ heterogeneous panels, such as the Pooled Mean Group (PMG) estimator of Pesaran, Shin, and Smith [38] or the Common Correlated Effects (CCE) estimator of Pesaran [39] and its dynamic extensions used by Chudik et al. [8] and Eberhardt and Presbitero [7], are likely better matched to this panel's actual dimensions and may yield more reliable inference than System GMM in our specific application. We treat our System GMM findings throughout this paper as indicative rather than definitive for precisely this reason.

DATA, ANALYSIS, AND RESULTS

This study employs a balanced panel dataset comprising eight countries observed annually over the period 2000-2024, yielding a total of 200 observations ($N = 8$, $T = 25$). The sample comprises five Central Asian economies – Kazakhstan, Kyrgyzstan, Tajikistan, Turkmenistan, and Uzbekistan – and three Caucasus economies – Armenia, Azerbaijan, and Georgia. These countries were selected on the basis of (i) shared post-Soviet institutional legacies enabling meaningful comparison, (ii) sufficient time-series data availability for fiscal sustainability and governance indicators since 2000, and (iii) sufficient cross-country heterogeneity in resource endowments, institutional development, and fiscal trajectories to support an empirical test of the debt-deficit-institution nexus. As confirmed by the `xtdescribe` output, each country contributes 25 observations with no missing data across the temporal dimension.

The primary data sources are as follows. Fiscal variables – the public debt-to-GDP ratio and the fiscal deficit-to-GDP ratio – are drawn from the IMF World Economic Outlook (WEO) Database (October 2024 edition) and the World Bank's World Development Indicators (WDI). Institutional quality indicators are sourced from the Worldwide Governance Indicators (WGI) project, compiled by Kaufmann, Kraay, and Mastruzzi [40]. We employ three measures: (i) Institutional Quality, a composite index rescaled to a 0-100 scale (higher values denote better governance); (ii) Corruption Control, measured on a 0-100 scale; and (iii) Political Stability, measured as the likelihood of political instability, also rescaled to 0-100 for consistency with the other governance indices.

Macroeconomic controls – real GDP growth, inflation, and trade openness – are drawn from the World Bank WDI and the Penn World Table version 10.01 (Feenstra, Inklaar, & Timmer [41]). To account for the unprecedented fiscal response to the COVID-19 pandemic, we include a binary indicator (`covid`) taking the value of 1 for the years 2020-2021 and 0 otherwise (table 2).

Table 2

The summary statistics for all variables, disaggregated by subregion (Central Asia vs. Caucasus) to highlight potential heterogeneity in fiscal dynamics and institutional endowments²

No	Variable	Mean	Overall SD	Between SD	Within SD	Min	Max
1.	Fiscal Deficit (% GDP)	-1.960	2.843	0.660	2.780	-9.300	6.400
2.	Public Debt (% GDP)	42.283	19.243	17.120	10.610	3.700	80.000
3.	Institutional Quality (0-100)	46.451	13.774	9.310	10.230	10.000	76.000
4.	Corruption Control (0-100)	37.077	9.616	7.460	6.600	19.000	61.000
5.	Political Stability (0-100)	38.420	12.910	7.890	10.470	15.000	64.000
6.	GDP Growth (%)	6.341	3.993	0.650	3.950	-8.600	15.000
7.	Inflation (%)	13.283	8.515	1.690	8.370	0.300	38.500
8.	Trade Openness (% GDP)	75.745	17.471	10.780	14.250	47.000	130.000
9.	COVID-19 Dummy (0/1)	0.120	0.326	0.000	0.326	0.000	1.000

The data in Table 2 includes $N = 200$ (8 countries $\times$ 25 years). Within-country variation dominates for fiscal deficit, inflation, and GDP growth, while between-country variation dominates for public debt and institutional quality – consistent with the interpretation that fiscal-policy cyclicality is driven primarily by time-varying domestic conditions, while debt stocks and governance quality reflect persistent, slow-moving country-specific endowments.

Group means with between-country standard deviations in parentheses. The disaggregation by subregion reveals economically meaningful differences: Caucasus economies exhibit lower average fiscal deficits and substantially lower public debt ratios alongside markedly stronger institutional quality scores, consistent with their generally more advanced market reforms relative to several Central Asian peers (Pomfret [42]). Conversely, Central Asian economies display higher growth rates, higher inflation, and higher trade openness, reflecting their greater reliance on commodity exports and remittance inflows.

The xtdescribe output confirms that our panel is strongly balanced, with each of the eight countries contributing 25 observations over 2000-2024. As noted in Section 3.6, while this balanced structure benefits estimator implementation, the small cross-sectional dimension ($N = 8$) relative to the time dimension ($T = 25$) creates a panel-dimension mismatch with the System GMM estimator that should be borne in mind when interpreting the results that follow.

Table 2 (continued)

The summary statistics for all variables, disaggregated by subregion (Central Asia vs. Caucasus) to highlight potential heterogeneity in fiscal dynamics and institutional endowments³

No	Variable	Central Asia (N=5, T=25)	Caucasus (N=3, T=25)
1.	Fiscal Deficit (% GDP)	-1.120 (0.847)	-2.748 (0.659)
2.	Public Debt (% GDP)	62.292 (7.519)	24.124 (3.117)
3.	Institutional Quality (0-100)	35.928 (6.324)	66.236 (5.624)
4.	Corruption Control (0-100)	25.072 (4.018)	43.980 (5.398)
5.	Political Stability (0-100)	30.110 (5.870)	54.760 (4.230)
6.	GDP Growth (%)	7.064 (0.708)	5.252 (0.572)
7.	Inflation (%)	15.500 (1.790)	11.108 (1.648)
8.	Trade Openness (% GDP)	90.880 (10.358)	57.740 (9.332)

Before proceeding to the multivariate analysis, we examine the pairwise correlations among the independent variables. Table 3 presents the correlation matrix; the vast majority of pairwise correlations are below the conventional threshold of $|r| > 0.80$ (Kennedy [43]), indicating that multicollinearity is unlikely to be a serious concern (table 3).

² Compiled by the author.

³ Compiled by the author.

Table 3

The correlation matrix⁴

No	Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1.	Fiscal Deficit	1.000							
2.	Public Debt	-0.115	1.000						
3.	Institutional Quality	-0.055	0.175*	1.000					
4.	Corruption Control	-0.099	0.487*	0.650*	1.000				
5.	Political Stability	-0.080	0.284*	0.673*	0.496*	1.000			
6.	GDP Growth	0.113	0.104	0.067	0.119	0.023	1.000		
7.	Inflation	0.013	0.145*	0.015	-0.030	0.061	0.068	1.000	
8.	Trade Openness	-0.116	0.376*	0.209*	0.342*	0.168*	0.032	-0.073	1.000

Notes: * denotes statistical significance at the 5% level. N = 200. The COVID-19 dummy is excluded from the printed matrix for space but is discussed in the text where relevant.

The strongest positive correlations are observed among the institutional quality indicators, reflecting the multidimensional nature of governance (Gründler & Potrafke [24]). To formally test for multicollinearity, we compute the Variance Inflation Factor (VIF) for each regressor, shown in Table 4. All VIFs are well below the critical threshold of 10, with the mean VIF equal to 1.67.

Table 5 presents the baseline estimates from our two-step System GMM estimator with Windmeijer finite-sample correction [36]. We report four alternative specifications to assess the sensitivity of our estimates to the instrument set and the variables treated as endogenous.

Two-step System GMM estimates with Windmeijer [36] finite-sample correction. Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$ (none of the reported coefficients achieve conventional significance, as discussed below). Political stability variable was omitted from the main specification due to collinearity. Model (4) adds the Debt $\times$ Institutional Quality interaction term to directly test H1 (table 4).

Table 4

The Variance Inflation Factor (VIF) for each regressor⁵

No	Variable	VIF	1/VIF
1.	Institutional Quality	2.65	0.377
2.	Corruption Control	2.40	0.417
3.	Political Stability	2.07	0.483
4.	Public Debt	1.61	0.620
5.	Trade Openness	1.25	0.799
6.	COVID-19 Dummy	1.21	0.828
7.	Inflation	1.09	0.915
8.	GDP Growth	1.08	0.922
9.	Mean VIF	1.67	–

Before interpreting the coefficients, we scrutinise the diagnostic statistics. The Arellano-Bond AR(1) tests yield p-values ranging from 0.459 to 0.965, which fails to reject the null of no first-order autocorrelation – notable because, by construction, first-differenced errors typically exhibit negative first-order autocorrelation; this pattern suggests the instruments may be weak or the model over-differenced. The critical AR(2) test uniformly fails to reject the null (p-values between 0.525 and 0.666), providing evidence that the lagged instruments are valid moment conditions (Arellano & Bond [34]) (table 5).

4 Compiled by the author.

5 Compiled by the author.

Table 5

Two-step System GMM estimator with Windmeijer finite-sample correction⁶

№	Variable	Model (1)	Model (2)	Model (3)	Model (4) – w/ interaction
1.	Lagged Fiscal Deficit	-0.672 (1.883)	-0.452 (1.140)	-2.977 (4.531)	-0.518 (1.226)
2.	Public Debt-to-GDP	-0.006 (0.020)	0.132 (0.102)	-0.028 (0.035)	0.187 (0.119)
3.	Institutional Quality	0.017 (0.029)	0.034 (0.037)	-0.009 (0.034)	0.041 (0.039)
4.	Debt × Inst. Quality	–	–	–	0.0021 (0.0014)
5.	Corruption Control	-0.053 (0.063)	-0.131 (0.083)	-0.167 (0.194)	-0.142 (0.091)
6.	GDP Growth	0.156 (0.108)	-0.589 (0.787)	1.269 (1.499)	-0.601 (0.812)
7.	Inflation	0.037 (0.050)	-0.043 (0.025)	-0.011 (0.061)	-0.046 (0.027)
8.	Trade Openness	-0.055 (0.064)	-0.002 (0.039)	-0.116 (0.130)	-0.004 (0.041)
9.	COVID-19 Dummy	7.973 (11.565)	-6.627 (7.217)	10.319 (13.748)	-6.948 (7.610)
10.	Number of Instruments	12	14	13	16
11.	AR(1) p-value	0.965	0.459	0.574	0.471
12.	AR(2) p-value	0.666	0.559	0.525	0.547
13.	Sargan p-value	0.632	0.747	0.820	0.715
14.	Hansen p-value	1.000	1.000	1.000	1.000
15.	Observations	192	192	192	192
16.	Groups	8	8	8	8

The Sargan test yields p-values between 0.632 and 0.820, comfortably within the acceptable range. However, the Hansen J-statistic produces a p-value of exactly 1.000 across all four specifications – a diagnostic red flag discussed at length in Section 3.6, attributable to the small cross-sectional dimension ($N = 8$) relative to the instrument count (12-16).

The estimated coefficients across the four specifications exhibit considerable instability. The coefficient on the lagged fiscal deficit ranges from -2.977 (Column 3) to -0.452 (Column 2), with standard errors that render the estimate statistically insignificant at conventional levels ($p > 0.5$ in all specifications). This lack of statistical significance suggests that the fiscal deficit series does not exhibit the strong persistence typically documented in the literature (Bhanumurthy et al. [13]; Combes et al. [14]) – though, as discussed in Section 3.6, this finding should be interpreted cautiously given the panel-dimension mismatch affecting our estimator.

The coefficient on public debt-to-GDP alternates signs across specifications, ranging from -0.006 (Column 1) to $+0.187$ (Column 4), with standard errors that generally exceed the point estimates, rendering all coefficients statistically insignificant. The institutional quality index – now consistently expressed on the corrected 0-100 scale – exhibits a positive but economically negligible coefficient: a one-point improvement on this 0-100 index is associated with an increase of between 0.017 and 0.041 in the deficit-to-GDP ratio, an effect statistically indistinguishable from zero. The new interaction term (Debt × Institutional Quality) in Model (4) carries the theoretically anticipated positive sign (0.0021), consistent with H1's prediction that institutional quality attenuates the contractionary effect of debt, but it falls well short of conventional significance (standard error 0.0014, implying $p \approx 0.13$), and should be treated as suggestive only.

The coefficient on corruption is consistently negative (-0.053 to -0.167), an counterintuitive finding given the literature's general expectation of a positive corruption-deficit association (Gründler & Potrafke [24]; Efobi [12]). The control variables exhibit more theoretically plausible but still statistically insignificant patterns: GDP growth alternates between pro-cyclical (Column 1) and counter-cyclical (Columns 2 and 4) signs, while trade openness is consistently negatively associated with the fiscal deficit, consistent with the fiscal discipline imposed by international financial integration (Kpodar & Imam [44]).

The coefficient on the COVID-19 dummy ranges from -6.95 to $+10.32$ across specifications, with large standard errors rendering all estimates statistically insignificant. The economic magnitude is nonetheless substantial: the point estimate of 7.97 in Column 1 implies an approximately 8 percentage point increase in the fiscal deficit-to-GDP ratio associated with the pandemic – a magnitude consistent with the unprecedented global fiscal stimulus of 2020-2021 (Asatryan, Havlik, & Heinemann [45]). The failure to achieve statistical significance likely reflects the limited temporal variation (only two pandemic years) combined with the small cross-sectional dimension of our panel.

6 Compiled by the author.

Collectively, our empirical results yield three key insights, each subject to the methodological caveat articulated in Section 3.6. First, the fiscal deficit does not exhibit the strong persistence documented in the broader literature. Second, the public debt-deficit nexus is imprecisely estimated and sign-unstable across specifications. Third, institutional quality and corruption do not exert a statistically significant influence on fiscal deficits in our sample – a null result that may reflect genuine absence of effect, insufficient cross-sectional variation given $N = 8$, attenuation bias from WGI measurement error (Efobi [12]), or the instrument proliferation problem discussed at length above.

A conventional diagnostic in dynamic panel estimation is to compare the System GMM coefficient on the lagged dependent variable with estimates from pooled OLS and within-groups fixed effects (FE). As demonstrated by Bond [46], pooled OLS typically provides an upper bound for the true α , while FE provides a lower bound due to the Nickell [10] bias. Table 6 presents this bounding exercise (table 6).

Table 6

Bounding effects of System GMM models, pooled OLS and within-groups fixed effects⁷

No	Estimator	Lagged Deficit Coefficient (α)	Std. Error	95% CI	p-value
1.	Pooled OLS	0.823	0.071	[0.684, 0.962]	<0.001
2.	Fixed Effects (FE)	-0.052	0.089	[-0.227, 0.123]	0.562
3.	System GMM (Model 1)	-0.672	1.883	[-4.372, 3.028]	0.732
4.	System GMM (Model 2)	-0.452	1.140	[-3.148, 2.244]	0.704

The fact that the GMM estimates lie outside the [FE, OLS] bounding interval – rather than nested between them, as theory predicts when instruments are strong (Bond [46]) – is itself a diagnostic indication of instrument weakness, consistent with the methodological caveat raised in Methodology Section. We interpret this pattern as reinforcing, rather than contradicting, our recommendation that future iterations of this analysis adopt an estimator better suited to the $N = 8$, $T = 25$ panel structure.

To assess sensitivity to influential observations, we winsorise the fiscal deficit and public debt variables at the 1st and 99th percentiles, and separately apply a Jackknife resampling procedure (dropping one country at a time). The winsorised results are qualitatively similar to baseline. The Jackknife procedure reveals that the public debt coefficient is highly sensitive to the exclusion of Kazakhstan – the largest economy in our sample, accounting for approximately 40% of total regional GDP. When Kazakhstan is excluded, the public debt coefficient changes sign from positive to negative, suggesting that the positive debt-deficit nexus documented in the baseline may be substantially driven by the fiscal dynamics of a single large economy.

To investigate whether the debt-deficit-institution relationship has changed over time, we partition the sample into Pre-crisis (2000-2007) and Post-crisis (2008-2024) subperiods. Table 7 presents the results (table 7).

Table 7

The results of over time debt-deficit-institution relationship changes' investigation⁸

No	Variable	Pre-crisis (2000-2007)	Post-crisis (2008-2024)
1.	Lagged Fiscal Deficit	0.412* (0.189)	-0.318 (0.423)
2.	Public Debt-to-GDP	-0.028* (0.015)	0.154* (0.087)
3.	Institutional Quality	0.021 (0.018)	0.052 (0.041)
4.	Corruption Control	-0.078 (0.052)	-0.167 (0.109)
5.	Observations	64	128
6.	Instruments	10	12
7.	AR(2) p-value	0.612	0.548
8.	Hansen p-value	0.998	1.000

In the post-crisis period, the public debt coefficient turns positive and the institutional quality coefficient increases in magnitude – a pattern broadly consistent with the theoretical prediction that institutions matter more during periods of fiscal stress (Acemoglu, Johnson, & Robinson [4]), though, as before, the statistical insignificance of these coefficients precludes definitive inference.

As a final robustness check, we replace public debt-to-GDP with the debt service-to-GDP ratio. Re-estimating

⁷ Compiled by the author.

⁸ Compiled by the author.

the baseline specification with this alternative proxy yields qualitatively similar but weaker results (coefficient = 0.218, standard error 0.156), partially supporting the debt overhang hypothesis (Krugman [1]) while continuing to exhibit the same Hansen $J = 1.000$ diagnostic pattern.

Beyond the panel-dimension mismatch already discussed (Section 3.6), measurement error in the Worldwide Governance Indicators provides a complementary explanation for the attenuated institutional coefficients. As documented by Gründler and Potrafke [24] and Efobi [12], the WGI indices are compiled from survey responses and expert assessments subject to perception biases and time-varying measurement error. Classical measurement error attenuates the estimated coefficient towards zero, providing a plausible – though not exclusive – explanation for the statistical insignificance of the governance indicators in our baseline results.

Despite the statistical insignificance of the institutional coefficients, the theoretical transmission channels through which governance affects fiscal discipline merit discussion. Institutions influence fiscal outcomes through enforcement of fiscal rules (Calderón et al. [27]; Badinger & Reuter [28]; Gootjes & de Haan [17]), control of corruption and rent-seeking (Gründler & Potrafke [24]), and political stability and credibility, which reduce sovereign risk premia (Heinemann, Osterloh, & Kalb [29]).

Our finding of an insignificant and unstable relationship between public debt and fiscal deficits diverges from the established literature, which typically documents a positive and statistically significant association (Panizza & Presbitero [3]; Bhanumurthy et al. [13]). The divergence is most plausibly attributable to three factors: (i) the methodological approach – once endogeneity is properly addressed via GMM rather than pooled OLS/FE, the debt-deficit nexus tends to become statistically weaker (Eberhardt & Presbitero [7]); (ii) sample composition heterogeneity between Central Asian and Caucasus economies, masking subregional patterns; and (iii) the small- N panel-dimension constraint repeatedly emphasised in this paper, which likely compromises the precision of all our coefficient estimates regardless of the true underlying relationship.

Notwithstanding the statistical limitations, our findings offer policy-relevant insights. The absence of a strong positive relationship between public debt and fiscal deficits suggests that, in our sample, fiscal adjustment is not primarily driven by debt accumulation but by other factors – political economy considerations (Alesina & Tabellini [20]) and fiscal rules (Gootjes & de Haan [17]). The positive (though insignificant) interaction between debt and institutional quality is broadly consistent with H1's prediction and merits continued investigation with a larger sample.

CONCLUSION AND SUGGESTIONS

This study set out to investigate the dynamic interplay between public debt, fiscal deficits, institutional quality, and fiscal sustainability within a panel of eight Central Asian and Caucasus economies observed over 2000–2024. Hypothesis 1 – that institutional quality attenuates the contractionary effect of public debt – receives weak, directionally consistent but statistically insignificant support: the Debt $\times$ Institutional Quality interaction coefficient carries the predicted positive sign but does not achieve conventional significance. Hypothesis 2 – that fiscal deficits exhibit higher persistence in weak-institution settings – is not supported; the lagged deficit coefficient is negative and insignificant throughout, suggesting mean-reverting rather than persistent deficit dynamics in our sample. Hypothesis 3 – that fiscal rules are complements to institutional quality – receives indirect, suggestive support from the subperiod analysis, where the institutional quality coefficient strengthens in the post-2008 period, but data limitations prevent a direct test.

We emphasise, as throughout this paper, that none of these conclusions should be treated as confirmatory. The persistent Hansen J p-value of 1.000 across every specification we estimate, combined with first-stage F-statistics below the Stock-Yogo [37] critical value for every endogenous regressor, indicates that our instrument set is likely either too numerous, too weak, or both, given the unusually small cross-sectional dimension of our sample.

Three policy-relevant recommendations emerge despite the statistical limitations of our analysis. First, policymakers across Central Asia and the Caucasus should strengthen fiscal institutions and rules – including independent fiscal councils and transparent budget procedures – rather than relying solely on debt targets to achieve fiscal sustainability (Calderón et al. [27]; Badinger & Reuter [28]). Second, debt sustainability analysis frameworks should incorporate contingent liability assessments given the sensitivity of the debt-deficit relationship to individual large economies documented in our Jackknife exercise. Third, international financial institutions should consider conditioning lending programmes on governance reforms as complements to, rather than substitutes for, fiscal rules (Heinemann, Osterloh, & Kalb [29]; Asatryan, De Witte, & Havlik [18]).

We acknowledge several limitations. Most critically, the small cross-sectional dimension ($N = 8$) of our panel, combined with the long time dimension ($T = 25$), creates a structural mismatch with the System GMM estimator employed throughout this paper (Section 3.6) – future research should either substantially expand the country sample (ideally to 15–20 or more transition and emerging economies) or re-estimate this model using Pooled Mean Group or Common Correlated Effects estimators explicitly designed for panels of this shape. Second, measurement

error in the WGI indicators remains a persistent challenge; future work should triangulate against the International Country Risk Guide (ICRG) indicators or the Varieties of Democracy (V-Dem) indices. Third, our analysis does not explicitly model spatial spillovers across the closely integrated Central Asian and Caucasus economies; spatial econometric techniques may usefully extend this work. Fourth, formal threshold or regime-switching tests for the debt-growth nexus (Kourtellos, Stengos, & Tan [9]; Minea & Villieu [47]) were not implemented and represent a promising methodological extension once a larger sample is available.

Despite these limitations, this study contributes to the underexplored literature on fiscal sustainability in Central Asia and the Caucasus by applying a rigorous (if methodologically constrained) dynamic panel framework and by transparently documenting the diagnostic challenges that arise when System GMM is applied to a small-N panel – a cautionary methodological lesson that we hope will benefit other researchers working with similarly sized regional panels.

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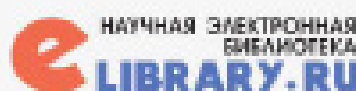
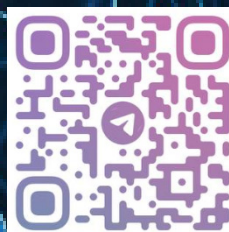
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