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Institutional Quality and the Resilience of Islamic Banking to Growth Shocks: Conditional Evidence on Non-Performing Financing

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Abstract
This study examines whether national institutional quality moderates the effect of adverse growth shocks on Non-Performing Financing (NPF) in Islamic banking. NPF is the dependent variable, a growth shortfall is the independent variable, a composite Worldwide Governance Indicators (WGI) index is the moderating variable, and inflation and income are control variables. The data are fourth-quarter Islamic Financial Services Board (IFSB) aggregates and World Bank indicators for 25 jurisdictions over 2013–2025. The analysis uses two-way fixed-effects models with a shock–governance interaction, complemented by two-step system Generalized Method of Moments (GMM) estimation. The interaction is negative but not significant under small-cluster inference, and the GMM estimates do not confirm it. Regulatory quality and control of corruption show negative interactions in exploratory tests, but the relationship disappears outside the pandemic years and among consistently reporting jurisdictions. A general institutional shock-absorber effect is therefore not established. For supervisors, low reported NPF during crises should not be read as resilience, and restructured and moratorium-affected financing should be disclosed separately. The study is limited by a short, unbalanced panel, slow-moving governance measures and aggregate data that cannot separate loss prevention from deferred recognition.

Keywords: Islamic Banking, Institutional Quality, Non-Performing Financing, Financial Resilience, Macroeconomic Shocks, Panel Data

JEL Classification Code: G21, G28, G01, O43, C23

1. Introduction

Islamic banking has become a systemically relevant segment of the financial sector in a growing number of jurisdictions, particularly in the Gulf Cooperation Council (GCC), Southeast Asia and parts of South Asia and Africa (Hassan & Aliyu, 2018; IFSB, 2021). Its intermediation model relies predominantly on sale-based, lease-based and, to a lesser extent, partnership-based contracts, and the quality of these financing assets is the principal determinant of its solvency (Aggarwal & Yousef, 2000; Hassan & Aliyu, 2018; Khan, 2010). Previous comparative work shows that Islamic banks differ from conventional banks in business orientation, efficiency, capitalisation and risk exposure, but that these differences are neither uniform across countries nor stable over time (Abedifar et al., 2013; Beck et al., 2013; Čihák & Hesse, 2010). Asset quality, measured by the ratio of Non-Performing Financing (NPF) to total financing, therefore offers a direct window on whether Islamic banking systems

can withstand adverse macroeconomic conditions (Ghosh, 2015; Louzis et al., 2012).

In jurisdictions where Islamic banks hold a large share of domestic banking assets, a deterioration in their financing portfolios can transmit directly to deposit holders, investment account holders and the payment system (Čihák & Hesse, 2010; Hassan & Aliyu, 2018). Supervisors and standard-setters, including the Islamic Financial Services Board (IFSB), have invested considerably in harmonised prudential indicators precisely so that the resilience of Islamic banking can be monitored across countries (IFSB, 2019). If national institutions systematically dampen the impact of downturns on asset quality, then institutional reform would be a complementary instrument of financial stability policy for Islamic banking (Barth et al., 2004; Bermpei et al., 2018); if the apparent dampening instead reflects slower recognition of losses, then low reported NPF in well-governed jurisdictions would convey a misleading sense of safety (Bushman & Williams, 2015; Huizinga & Laeven, 2012). Distinguishing between these possibilities requires an explicit empirical framework in which institutions moderate, rather than merely accompany, the transmission of shocks.

The resilience of Islamic banking systems depends on the quality of financing decisions, the ability to restructure exposures and the speed with which impaired financing is recognised (Abedifar et al., 2013; Bushman & Williams, 2015). A large body of evidence on conventional banking documents that economic downturns raise credit risk by weakening borrowers' cash flows and collateral values (Beck et al., 2015; Louzis et al., 2012; Salas & Saurina, 2002). Evidence for Islamic banks points in a similar direction, although the magnitude of the response depends on the contractual structure of financing and on the market in which banks operate (Alandejani & Asutay, 2017; Kabir et al., 2015). Yet the same macroeconomic setback need not generate the same observed NPF response in all countries. Contract enforcement, administrative capacity, the quality of regulation and the control of corruption plausibly alter how a growth shock is transmitted to financing portfolios (Djankov et al., 2003; Jappelli et al., 2005; Qian & Strahan, 2007).

The direction of this institutional influence on measured NPF is, however, theoretically ambiguous (Ashraf, 2017). Stronger institutions may prevent losses by improving screening, recovery and supervisory follow-through; they may equally cause losses to be identified and reported more promptly (Barth et al., 2004; Huizinga & Laeven, 2012). Studies of bank regulation and supervision show that the effects of formal rules depend on the institutional environment in which they are applied (Barth et al., 2004; Bermpei et al., 2018; Laeven & Levine, 2009), while research on loan-loss recognition and forbearance shows that weak incentives can suppress reported impairment for extended periods (Bushman & Williams, 2015; Peek & Rosengren, 2005). A cross-country analysis of Islamic banking must therefore treat observed NPF as the joint outcome of economic losses and of the institutions that govern their recognition.

The COVID-19 episode makes this measurement problem especially consequential (Demirgüç-Kunt et al., 2021; Elnahass et al., 2021). The IFSB noted that regulatory forbearance and payment moratoria may have flattened reported NPF during part of the pandemic (IFSB, 2021). Cross-country studies of Islamic banking during the pandemic report heterogeneous policy responses and performance outcomes (Mansour et al., 2022; Nugrohowati & Fakhrunnas, 2024). Danlami et al. (2022) examine institutional quality within a bank-level stability framework for six member countries of the Organisation of Islamic Cooperation (OIC), and Mollah et al. (2017) show that governance structures inside Islamic banks shape their risk-taking. At the country level, Yunan (2021) finds that corruption affects both the profitability and the soundness of Islamic banks in OIC countries, and Yunan et al. (2024) show that political stability and law and order are positively associated with the health of Islamic banks. What remains largely unexamined is whether country-level institutions moderate the relationship between macroeconomic shocks and asset quality for aggregated Islamic banking systems across a broad set of jurisdictions (Bermpei et al., 2018; Danlami et al., 2022; Narayan & Phan, 2019).

This study addresses that gap and makes four contributions. Conceptually, it frames governance as a potential shock-absorption mechanism rather than merely a direct correlate of banking outcomes, and it makes explicit that a shock-absorber effect and a recognition effect imply opposite signs for the same interaction term (Bermpei et al., 2018; Brambor et al., 2006). Empirically, it estimates the interaction between a composite institutional quality index and a precisely defined growth shortfall using the prudential and structural Islamic financial indicators compiled by the IFSB (IFSB, 2026). In terms of heterogeneity, it distinguishes government effectiveness, regulatory quality, rule of law and control of corruption, and examines whether the pattern depends on the pandemic years and on the composition of reporting jurisdictions (Kaufmann et al., 2011; Langbein & Knack, 2010). Methodologically, it adopts a sequential design: country and year fixed effects form the main specification, while dynamic system Generalized Method of Moments (GMM) estimation is reported as a diagnostic complement whose weight in the conclusions depends on its specification tests (Blundell & Bond, 1998; Roodman, 2009b).

The analysis is deliberately conducted at the level of national Islamic banking systems. Because the moderator of interest, national institutional quality, varies only across jurisdictions and over time, sector-level aggregates align the level of measurement of the outcome with that of the moderator and avoid the overstated precision that arises when an aggregate regressor is matched to many micro units (Moulton, 1990). The IFSB aggregates are also compiled under a common methodology, which improves cross-country comparability (IFSB, 2019). This choice has a cost: aggregation conceals heterogeneity within each jurisdiction, for example between retail-oriented and corporate-oriented Islamic banks, between large and small institutions, and

between portfolios dominated by sale-based and by partnership-based contracts (Abedifar et al., 2013; Čihák & Hesse, 2010). The estimates therefore describe how system-wide asset quality responds to shocks, not how an individual bank responds, and they should not be read as bank-level behavioural parameters (Moulton, 1990; Robinson, 1950).

The results place clear limits on a broad resilience claim. The fixed-effects interaction between the growth shortfall and institutional quality is negative, consistent with a weaker NPF response in better-governed jurisdictions (Table 4), but it does not meet the five per cent threshold under country-clustered inference and is weaker still under a wild cluster bootstrap (Cameron et al., 2008; Cameron & Miller, 2015). The dynamic specifications do not provide independent confirmation, and the aggregate interaction disappears when 2020 and 2021 are excluded or when the sample is restricted to jurisdictions that report consistently. Dimension-specific interactions for regulatory quality and control of corruption are negative, but they are exploratory and unadjusted for multiple testing (Holm, 1979) (Table 5). Reporting these boundary conditions transparently is itself informative for Islamic finance research, which frequently relies on short, unbalanced country panels and on reported asset-quality ratios (Hassan & Aliyu, 2018; Narayan & Phan, 2019).

Accordingly, the objective of this study is to assess whether, and under which conditions, national institutional quality moderates the association between adverse growth shocks and NPF in Islamic banking systems, and to identify which institutional dimensions, periods and samples drive any observed moderation. The remainder of the paper is organised as follows. Section 2 reviews the literature, sets out the theoretical framework and develops the hypotheses. Section 3 describes research methods. Section 4 presents results and discussion, and Section 5 concludes.

2. Literature Review

The link between the business cycle and bank credit risk is one of the most consistently documented relationships in banking research. Salas and Saurina (2002) show that real Gross Domestic Product (GDP) growth, together with bank-specific characteristics, explains problem loans among Spanish commercial and savings banks, and that the institutional form of the lender changes the strength of this relationship. Louzis et al. (2012) find that non-performing loans in Greece are driven mainly by real GDP growth, unemployment, interest rates and public debt, with differences across consumer, business and mortgage portfolios. Using a global sample, Beck et al. (2015) confirm that real GDP growth is the principal driver of non-performing loan ratios, while share prices, exchange rates and lending interest rates also matter, depending on the structure of the economy (Beck et al., 2015). These studies share the premise that an adverse growth

realisation lowers borrowers' income and collateral values and therefore raises default rates, typically with a lag.

The same mechanism applies to Islamic banks, but the contractual form of financing modifies it (Abedifar et al., 2013; Kabir et al., 2015). Sale-based contracts such as *murabahah* fix the bank's mark-up at origination, so repayment obligations do not adjust to adverse conditions, while partnership contracts such as *musharakah* and *mudarabah* allow losses to be shared rather than recorded as default (Aggarwal & Yousef, 2000; Khan, 2010). The Islamic prohibition of compounding penalties on late payment may also weaken the deterrent against strategic delay. Abedifar et al. (2013) find that small Islamic banks that are leveraged or based in predominantly Muslim countries have lower credit risk than conventional banks, which they attribute partly to borrower religiosity. Kabir et al. (2015) show that conclusions about comparative credit risk depend on the measure used, Islamic banks exhibit lower market-based credit risk but higher accounting-based credit risk than conventional banks. Alandejani and Asutay (2017) document that non-performing loans in GCC banking systems respond to macroeconomic and bank-specific factors and that Islamic banks display distinct NPL dynamics. Hassan et al. (2019) further show that credit risk and liquidity risk interact in shaping the stability of both Islamic and conventional banks. Together, this evidence suggests that NPF is sensitive to growth conditions but that the sensitivity is mediated by contracts, markets and measurement practices.

The contractual channel also implies that institutions matter in a specific way for Islamic banks (Hasan & Asutay, 2011). Because many Islamic financing contracts are asset-backed, the ability to repossess, lease or sell the underlying asset depends on the enforceability of property rights and on the efficiency of courts, which are features captured by the rule of law (Djankov et al., 2003; Jappelli et al., 2005; Kaufmann et al., 2011; Qian & Strahan, 2007). In several jurisdictions, however, enforcement operates within a dual or mixed legal system. In Malaysia, Islamic financing disputes are heard by the civil courts, which are required to refer Shariah questions to the central bank's Shariah Advisory Council (Hasan & Asutay, 2011), whereas in the GCC states banking disputes are typically settled by civil or commercial courts or specialised banking-dispute committees applying codified laws that were not designed for the distinctive features of Islamic financial products (Al-Shibli, 2018). Where the Shariah compliance or enforceability of a contract is contested, the outcome depends on how these systems interact, and cross-border financing is frequently governed by English law (Hasan & Asutay, 2011). Such arrangements can lengthen and add uncertainty to repossession and recovery. Because the speed of contract enforcement differs widely across legal systems (Djankov et al., 2003) and shapes the terms on which banks lend (Qian & Strahan, 2007), enforcement delays are likely to matter most in downturns, when many contracts become distressed at the same time.

Partnership-based financing further relies on reliable information about the borrower's business, which depends on accounting standards, disclosure and the absence of corruption (Aggarwal & Yousef, 2000; Khwaja & Mian, 2005). Where these conditions are weak, Islamic banks may favour short-term sale-based contracts that shift risk back to the client, so that adverse shocks translate more directly into repayment difficulties (Aggarwal & Yousef, 2000; Khan, 2010). Where they are strong, banks may be better able to restructure financing, for example by rescheduling instalments or converting obligations into lease arrangements, without recording an impairment (IFSB, 2021; Mansour et al., 2022). These considerations suggest that the transmission of macroeconomic shocks to Islamic banking asset quality is unlikely to be homogeneous across institutional environments.

A second strand asks whether Islamic banks are more resilient than conventional banks when crises occur. Čihák and Hesse (2010) find that small Islamic banks tend to be financially stronger than small conventional banks, whereas large Islamic banks tend to be weaker than large conventional banks, a result they attribute to the difficulty of adapting credit risk management as Islamic banks grow. Beck et al. (2013) report that Islamic banks were better capitalised and had higher asset quality during the global financial crisis, although they find few significant differences in business orientation. Bourkhis and Nabi (2013), by contrast, find no significant difference between the soundness of Islamic and conventional banks during the 2007–2008 crisis. The pandemic has renewed this debate. Mansour et al. (2022) review the regulatory measures adopted for Islamic banking during the outbreak of COVID-19 and show that relief packages, moratoria and prudential adjustments varied considerably across jurisdictions. Nugrohowati and Fakhrunnas (2024) find that the pandemic affected Islamic banking performance across countries, with differing effects on profitability and financing quality.

Elnahass et al. (2021) show that the pandemic affected the stability of conventional and Islamic banking systems differently, and Ashraf et al. (2022) find that stock market investors in the GCC did not regard Islamic banks as more resilient than conventional banks during the COVID-19 shock. Evidence from earlier stress episodes is similarly nuanced: Ibrahim and Rizvi (2018) find that Islamic banks in dual-banking countries sustained financing growth during crises without taking on excessive risk. The crisis literature therefore suggests that resilience is conditional and that the policy environment in which a shock occurs matters for the observed outcome.

A third strand examines how the institutional and regulatory environment shapes bank risk. Barth et al. (2004) show that regulatory strategies that force accurate information disclosure and empower private monitoring are associated with better banking performance, whereas strengthening official supervisory power does not necessarily improve bank development or stability. Laeven and Levine (2009) demonstrate that the effect of a given regulation on

bank risk-taking depends on the ownership structure of each bank, implying that identical rules can produce different outcomes. Boudriga et al. (2009) find that Capital Adequacy Ratio (CAR) and provisioning policies relate to lower non-performing loans across countries, and that the institutional environment conditions the effectiveness of supervisory tools. Bermpei et al. (2018) show directly that institutional quality conditions the effect of regulation and supervision on bank stability in emerging and developing economies. Ashraf (2017) demonstrates that political institutions affect bank risk-taking, but not necessarily in a stabilising direction, because stronger institutions may also expand credit and risk appetite.

Within Islamic banking, governance has been examined mainly at the bank level. Mollah et al. (2017) show that board structure and Shariah supervisory boards influence the risk-taking and performance of Islamic banks. Danlami et al. (2022) combine Capital, Asset, Management, Earnings, and Liquidity (CAMELS) indicators, risk-sharing financing and institutional quality in six OIC countries and find that institutional quality contributes to Islamic bank stability. Evidence on country-level institutions is also emerging. Using 61 Islamic banks in 12 OIC countries, Yunan (2021) shows that corruption affects profitability and soundness simultaneously, rather than only one of these dimensions. Yunan et al. (2024) extend this line of inquiry to political risk and find that government stability and law and order are positively correlated with the health of Islamic banks, whereas corruption and military involvement in politics create uncertainty for the businesses that Islamic banks finance. Mergaliyev et al. (2021) show that the ethical performance of Islamic banks, measured against the higher objectives of Shariah, is associated with governance and country-level characteristics. These studies establish that governance matters for Islamic banks, but they generally model institutional quality as a direct determinant of stability rather than as a moderator of how external shocks are transmitted to asset quality.

A fourth strand concerns what reported asset-quality ratios measure. Peek and Rosengren (2005) show that Japanese banks extended credit to impaired firms to avoid recognising losses, a practice commonly described as evergreening. Bushman and Williams (2015) find that delayed recognition of expected loan losses increases banks' exposure to downturns and their contribution to systemic risk. For Islamic banking, the IFSB compilation guide defines gross NPF as financing past due or impaired before the deduction of specific provisions, but national classification rules and forbearance practices differ (IFSB, 2019). During the pandemic, temporary relaxation of classification rules may have lowered reported NPF independently of underlying credit quality (IFSB, 2021). Institutional quality itself is also measured imperfectly. The Worldwide Governance Indicators (WGI) aggregate perception-based sources into estimates with explicit margins of error, and they are designed primarily for cross-country comparison rather than for detecting small within-country changes

over time (Kaufmann et al., 2011). These measurement features matter for any within-country estimation strategy.

The literature thus establishes three points: growth shocks raise credit risk; Islamic banking resilience during crises is heterogeneous; and institutions shape how regulation affects bank risk. Closely related studies either examine Islamic bank stability during crises without modelling national institutions as a moderator (Elnahass et al., 2021; Ibrahim & Rizvi, 2018) or treat institutions as direct determinants of bank-level stability (Danlami et al., 2022; Yunan et al., 2024). The literature does not, however, provide a systematic cross-country test of whether national institutional quality moderates the transmission of growth shocks to Islamic banking NPF, nor does it examine whether such moderation survives small-cluster inference, alternative estimators and the exclusion of the pandemic years (Bermpei et al., 2018; Cameron & Miller, 2015; Danlami et al., 2022). This study addresses that gap by specifying the moderation explicitly, testing it with aggregate Islamic banking data from the IFSB and reporting the conditions under which the apparent relationship appears or disappears.

The framework combines the macroeconomic view of credit risk with institutional theories of financial contracting. A macroeconomic shock reduces the income of households and firms, but whether lower income turns into recorded non-performance depends on the rules and organisations that govern credit relationships (Barth et al., 2004; Bermpei et al., 2018). Institutions shape three stages of this conversion. Ex ante, they affect the quality of screening and the allocation of financing, and thus the vulnerability of the portfolio when a shock arrives (Khwaja & Mian, 2005). During the shock, they affect the availability and credibility of restructuring, rescheduling and public support (Demirgüç-Kunt et al., 2021). Ex post, they affect whether impaired exposures are recovered, written off or reported accurately (Bushman & Williams, 2015; Djankov et al., 2003).

Two competing mechanisms follow from this sequence. Under the shock-absorber mechanism, stronger institutions reduce the conversion of shocks into losses at the first two stages, because enforceable contracts, credible regulation and clean credit allocation make it easier for borrowers and banks to restructure, enforce and recover (Jappelli et al., 2005; Qian & Strahan, 2007). Under the recognition mechanism, stronger institutions raise the reported share of impaired financing at the third stage, because they limit evergreening and the extensive use of forbearance (Huizinga & Laeven, 2012; Peek & Rosengren, 2005). The two mechanisms predict opposite signs for the same interaction term, so the net effect on observed NPF is an empirical question.

Four institutional dimensions are relevant to this transmission. Government effectiveness captures the quality of public administration and can shape how supervisory policies are implemented during a downturn. Regulatory quality captures the government's capacity to formulate and implement sound policies that permit private-sector development, and it may affect bank conduct

and the consistency of prudential rules. The rule of law captures confidence in contract enforcement, property rights and the courts, and it can influence repayment discipline and the recovery of impaired financing. Control of corruption captures the extent to which public power is exercised for private gain and may influence both the screening of financing applications, for instance by limiting connected lending, and the honesty with which impairment is recognised (Kaufmann et al., 2011). Corruption has been shown to weaken Islamic bank profitability and soundness (Yunan, 2021) and to deter foreign investors who face an opaque regulatory landscape in emerging economies (Yunan & Pacheco-Jaramillo, 2025), both of which can narrow the capacity of borrowers and banks to absorb a downturn. The WGI estimates are broad, perception-based measures of national governance; they are not direct indices of Islamic banking supervision or of Shariah governance (Kaufmann et al., 2011; World Bank, 2026a).

Figure 1 summarises the research model. The growth shortfall is the independent variable, gross NPF is the dependent variable, and institutional quality, measured by the composite index and by its four dimensions, is the moderating variable that may weaken or strengthen the effect of the shock on NPF. Inflation and log real GDP per capita are control variables, while country and year effects absorb time-invariant national characteristics and common global shocks. In the empirical model, the moderation is captured by the interaction between the growth shortfall and institutional quality (Brambor et al., 2006).

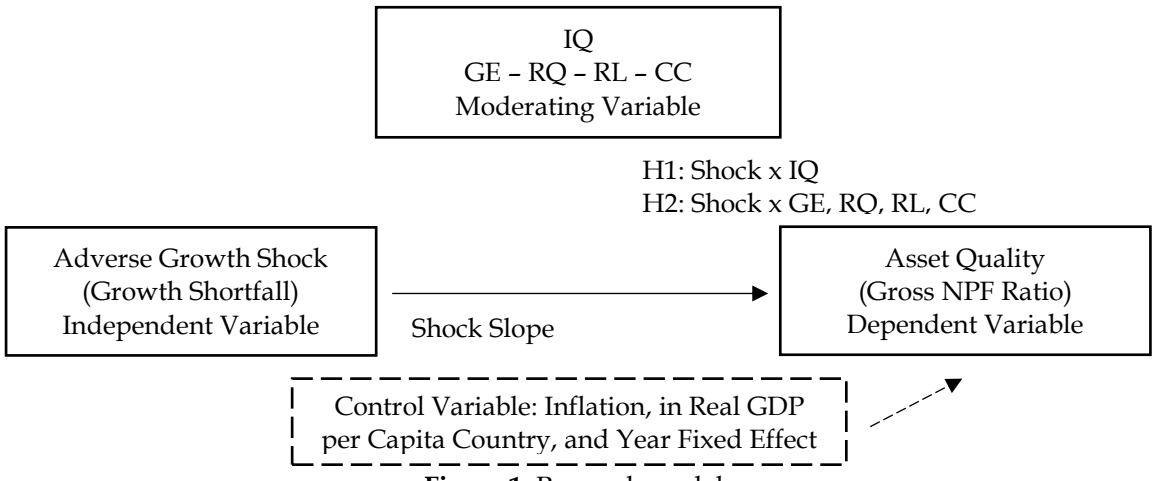

Figure 1: Research model

Hypotheses

Building on this framework, the hypotheses are formulated as two-sided tests. If the shock-absorber mechanism dominates, the NPF response to a growth shortfall should be weaker where institutional quality is higher, which implies a negative interaction coefficient. If the recognition mechanism dominates, better-

governed systems should report impaired financing more promptly during downturns, which implies a positive interaction coefficient (Bushman & Williams, 2015; Peek & Rosengren, 2005). Because the four WGI dimensions capture different features of governance, their moderating roles may also differ (Kaufmann et al., 2011). The hypotheses are therefore stated as follows:

H1: National institutional quality significantly moderates the effect of adverse growth shocks on NPF in Islamic banking systems. A negative interaction indicates a shock-absorber role, whereas a positive interaction indicates a recognition role.
H2: The moderating effect of institutional quality on the relationship between adverse growth shocks and NPF differs across government effectiveness, regulatory quality, rule of law and control of corruption.

The four dimension-specific interactions in H2 are treated as exploratory probes of which institutional features correlate with the observed shock slope. A dimension that crosses a conventional p-value threshold is not treated as a confirmed channel, because the four tests are not adjusted for multiplicity and the dimensions are highly correlated with one another. In the estimation sample, the pairwise correlations between the four WGI dimensions range from 0.91 to 0.95, and the variance inflation factors of the dimension-specific interaction terms range from about 13 to 26 when the four are entered jointly. Separating the dimensions therefore inflates the variance of the estimates and makes individual coefficients difficult to isolate, which is consistent with evidence that the WGI dimensions largely measure a common underlying perception of governance (Langbein & Knack, 2010). For this reason, the dimensions are entered one at a time, and their coefficients are not interpreted as separately identified channels.

3. Research Method

The dependent variable is the fourth-quarter gross NPF ratio for the Islamic Banking category in the IFSB Prudential and Structural Islamic Financial Indicators (PSIFIs), series CP04 (IFSB, 2026); Islamic banking windows are excluded. The ratio divides gross NPF by total outstanding financing before the deduction of specific provisions (IFSB, 2019). The source grid covers 26 jurisdictions from 2013 to 2025 and contains 237 reported NPF records. Real GDP growth, consumer price inflation and real GDP per capita are drawn from the World Development Indicators (WDI) (World Bank, 2026b). Institutional quality (IQ) is the unweighted mean of four WGI estimates, namely government effectiveness, regulatory quality, rule of law and control of corruption, taken from the 2026 WGI release (World Bank, 2026a). Because this release recalculates historical estimates, it is not interchangeable with earlier vintages. In the

regressions, IQ is centred at its estimation-sample mean so that the coefficient on the growth shock measures the shock slope at average institutional quality.

Table 1: Variable operationalisation and sources

Variables	Status	Operational definition	Sources
Gross NPF	Dependent Variable	Gross NPF / total financing outstanding before specific provisions; Q4; %	IFSB PSIFIs, Islamic Banking
CAR (alternative outcome)	Alternative dependent variable	Total regulatory capital / risk-weighted assets, as reported; Q4; %	IFSB PSIFIs, Islamic Banking
Return on Assets (ROA) (alternative outcome)	Alternative dependent variable	Reported ROA for Islamic banking; Q4; %, annualization subject to IFSB and country metadata	IFSB PSIFIs, Islamic Banking
Institutional quality (IQ)	Moderating variable	Arithmetic mean of the four WGI estimates below; centred at the estimation-sample mean in regressions	WGI
Government effectiveness (GE), regulatory quality (RQ), rule of law (RL), control of corruption (CC)	Moderating variables (dimension-specific models)	Four WGI governance dimensions; governance estimate units (approximately −2.5 to 2.5)	WGI
Growth shortfall (Shock)	Independent variable	Max [0, mean real GDP growth over t−3 to t−1 minus growth at t]; percentage points	World Development Indicators
Real GDP growth	Input to the growth shortfall	Annual change in constant-price GDP; %	World Development Indicators
Inflation	Control variable	Annual consumer price inflation; %	World Development Indicators
Real GDP per capita	Control variable	Constant-US$ GDP per capita; natural logarithm in regressions	World Development Indicators

Notes: Data are annual and identified by country (ISO3 code) and year. GDP growth for 2010–2012 is used only to calculate the first growth shortfalls. IFSB data cover full-fledged Islamic banks only; Islamic banking windows are excluded. GE = government effectiveness; RQ = regulatory quality; RL = rule of law; CC = control of corruption.

Sector-level aggregates are used because national institutions vary only across jurisdictions and over time, so bank-level data would add observations without adding independent institutional environments (Moulton, 1990).

Aggregates compiled under a common IFSB methodology are also more comparable across countries than bank financial statements (IFSB, 2019). The cost is that individual banks and business models cannot be distinguished, a limitation discussed in Section 5.

The adverse growth shock is defined as a growth shortfall, the maximum of zero and the difference between the mean real GDP growth rate over the preceding three years and contemporaneous growth, expressed in percentage points. This definition captures deviations below a country's recent growth path rather than low growth as such, and it is non-negative by construction, so that only adverse surprises enter the interaction. Inflation and the natural logarithm of real GDP per capita are the common controls, and all models include year effects. NPF is not interpolated. Of the 237 NPF records, 236 have complete information on all regressors; removing Libya's single complete record, which is a singleton in the fixed-effects model, leaves 235 country-years from 25 jurisdictions. Only 14 jurisdictions provide a completely balanced panel over 2017–2023, and the years 2024 and 2025 each contribute 11 observations to the estimation sample. This attrition limits generalisation and motivates the sample-composition checks reported below. Table 1 summarises the operational definitions and sources, and Table 2 lists the reporting jurisdictions and their coverage.

Table 2: Reporting jurisdictions and coverage of the gross NPF ratio, 2013–2025

Jurisdiction	Code	Group	Years with reported NPF	NPF records	TWFE sample
Bahrain	BHR	GCC	2013–2022, 2024–2025	12	12
Kuwait	KWT	GCC	2013–2023, 2025	12	12
Oman	OMN	GCC	2013–2025	13	13
Qatar	QAT	GCC	2016–2023	8	8
Saudi Arabia	SAU	GCC	2013–2023	11	11
United Arab Emirates	ARE	GCC	2013–2025	13	12
Egypt	EGY	Other MENA	2013–2023	11	11
Iran	IRN	Other MENA	2014–2017	4	4
Iraq	IRQ	Other MENA	2020–2023	4	4
Jordan	JOR	Other MENA	2013–2025	13	13
Lebanon	LBN	Other MENA	2013–2021	9	9
Libya	LBY	Other MENA	2021	1	0
Morocco	MAR	Other MENA	2018–2020	3	3
Palestine	PSE	Other MENA	2016–2025	10	10
Sudan	SDN	Other MENA	2013–2022	10	10
Afghanistan	AFG	South & Central Asia	2022–2023	2	2

Bangladesh	BGD	South & Central Asia	2013–2021	9	9
Kazakhstan	KAZ	South & Central Asia	2015–2018	4	4
Kyrgyz Republic	KGZ	South & Central Asia	2020–2023	4	4
Pakistan	PAK	South & Central Asia	2013–2025	13	13
Brunei Darussalam	BRN	Southeast Asia	2013–2025	13	13
Indonesia	IDN	Southeast Asia	2013–2025	13	13
Malaysia	MYS	Southeast Asia	2013–2025	13	13
Nigeria	NGA	Sub-Saharan Africa	2013–2025	13	13
Turkey	TUR	Europe	2013–2024	12	12
United Kingdom	GBR	Europe	2018–2023, 2025	7	7

Notes: The IFSB uses the term "jurisdiction" because data are reported by supervisory authorities, such as central banks or Indonesia's Financial Services Authority (OJK), and not every reporting authority represents a sovereign state. NPF records are year-end gross NPF ratios for full-fledged Islamic banks (IFSB PSIFIs, series CP04). The TWFE sample counts observations with complete data on all variables. The 2025 record of the United Arab Emirates is not used because the WDI data for that year are not yet available, and Libya's single observation is dropped because it cannot be used in a fixed-effects model. GCC marks the six jurisdictions excluded in the robustness check in Table 5.

Table 3 shows that the mean NPF ratio is 4.779%, above its median of 3.420%, and that the ratio reaches 41.290% in the most distressed observation, indicating a right-skewed distribution. The growth shortfall is zero in 113 of the 235 observations, so roughly half of the sample experiences no adverse growth surprise. Inflation is highly dispersed, its median is 3.324%, while its maximum is 359.093%. Institutional quality is, on average, slightly below the WGI global mean of zero, with the rule of law and control of corruption the weakest dimensions. Reporting coverage is also uneven, the estimation sample contains 22 observations in 2020 but only 11 in each of 2024 and 2025. These distributions motivate the country-influence, pandemic-period and reporting-coverage checks; they do not on their own identify a governance effect.

Table 3: Descriptive statistics for the fixed-effects estimation sample (235 country-years, 25 jurisdictions, 2013–2025)

Variable	N	Mean	SD	Median	Min	Max
Gross NPF (%)	235	4.779	5.328	3.420	0.000	41.290
Growth shortfall (pp)	235	1.626	3.191	0.145	0.000	26.232
Institutional quality	235	−0.119	0.727	−0.041	−1.655	1.616
Government effectiveness	235	−0.003	0.748	0.052	−2.105	1.517
Regulatory quality	235	0.021	0.764	0.106	−1.590	1.715
Rule of law	235	−0.286	0.668	−0.223	−1.897	1.459
Control of corruption	235	−0.210	0.816	−0.276	−1.531	1.810
Inflation (%)	235	10.536	30.593	3.324	−4.645	359.093
Log real GDP per capita	235	8.977	1.240	9.041	5.934	11.117

Notes: Institutional quality is the simple average of the four WGI dimensions; in the regressions it is centred at its sample mean (−0.119). The WGI dimensions are shown on their original scale. The growth shortfall cannot be negative; pp = percentage points.

The analysis proceeds sequentially. The baseline specification is a Two-Way Fixed-Effects (TWFE) model:

$$NPF_{it} = a + \beta\, IQ_{it} + \delta\, Shock_{it} + \gamma' X_{it} + \mu_i + \lambda_t + \varepsilon_{it}$$

where NPF_{it} is the gross NPF ratio of jurisdiction i in year t, X_{it} comprises inflation and log real GDP per capita, μ_i denotes country effects and λ_t denotes year effects. Country effects absorb time-invariant differences in legal tradition, market structure, the share of partnership contracts and national classification conventions, whereas year effects absorb common global shocks such as oil-price movements and the worldwide component of the pandemic. The model is estimated by ordinary least squares with explicit two-way indicators. It estimates a within-country conditional association and does not resolve time-varying confounding.

The main specification adds the interaction between the growth shortfall and centred institutional quality:

$$NPF_{it} = a + \beta_1\, IQ_{it} + \beta_2\, Shock_{it} + \beta_3\, (Shock_{it} \times IQ_{it}) + \gamma' X_{it} + \mu_i + \lambda_t + \varepsilon_{it}$$

The coefficient β_3 measures the change in the conditional shock slope associated with a one-unit increase in IQ, and $\partial NPF/\partial Shock = \beta_2 + \beta_3 IQ$ gives the marginal effect of the shock at any level of institutional quality. Because the significance of an interaction coefficient does not by itself indicate where the marginal effect is distinguishable from zero, marginal shock slopes are computed at the 25th and 75th percentiles of IQ, together with their confidence intervals (Brambor et al., 2006). The four dimension-specific models replace IQ with each WGI dimension in turn.

The dynamic stage adds lagged NPF to the interaction model in Equation (2):

$$NPF_{it} = \rho\, NPF_{i,t-1} + \beta_1\, IQ_{it} + \beta_2\, Shock_{it} + \beta_3\, (Shock_{it} \times IQ_{it}) + \gamma' X_{it} + \lambda_t + \mu_i + \varepsilon_{it}$$

System GMM is used as a complementary estimator for two reasons. First, NPF is persistent, and adding its lag to a fixed-effects regression produces biased estimates when the time dimension is short (Nickell, 1981). Second, institutional quality and the growth shock may be correlated with past shocks to NPF, and internal instruments help to address this possible endogeneity (Arellano & Bover, 1995; Blundell & Bond, 1998). The estimation follows five steps. First, Equation (3) adds lagged NPF to the interaction model. Second, the equation is first-differenced to remove the country effect, and the differenced variables are instrumented with their own lagged levels (Arellano & Bond, 1991); system GMM adds the equation in levels, instrumented with lagged differences, which improves precision when the series are persistent (Blundell & Bond, 1998). Third,

the instrument set is kept small: NPF, the growth shortfall, IQ and Shock × IQ are instrumented with collapsed GMM-style instruments at lags 2–3, inflation and log real GDP per capita are treated as exogenous standard instruments, year indicators are included, and the window is restricted to 2014–2023 so that the number of instruments stays below the number of jurisdictions (Roodman, 2009b). Fourth, the model is estimated in two steps, with standard errors corrected for finite-sample bias (Windmeijer, 2005). Fifth, the estimates are interpreted only if the diagnostics are satisfactory: the AR (1) test should reject and the AR(2) test should not, the Hansen test should not reject the instrument set, and the persistence coefficient ρ should be below one. A Hansen non-rejection does not validate the additional level moments, and the test weakens as the instrument count approaches the number of groups (Roodman, 2009a, 2009b). Because only 24 jurisdictions are eligible, the dynamic models therefore serve as a diagnostic complement rather than as the main evidence.

With 25 country clusters, conventional cluster-robust standard errors may understate uncertainty (Cameron & Miller, 2015). Standard errors in the TWFE models are therefore clustered by country with a CR1 finite-sample adjustment, and t tests use $G - 1$ degrees of freedom, where G is the number of clusters. For the main interaction, a restricted-null, country-level wild cluster bootstrap-t p-value is also computed using 999 Rademacher draws (Cameron et al., 2008). In simple terms, the bootstrap repeatedly re-randomises the residuals at the country level while imposing the hypothesis of no interaction, and it asks how often an interaction as large as the observed one arises by chance; unlike conventional standard errors, it does not rely on approximations that require many clusters. Robustness is assessed by estimating Equation (2) separately for each institutional dimension, excluding the six GCC jurisdictions, excluding 2020 and 2021, ending the sample in 2023, restricting the sample to a balanced panel of consistently reporting jurisdictions, estimating the model in first differences, using one-year lagged IQ and replacing NPF with Return on Assets (ROA) or CAR as alternative outcomes.

Each sensitivity check addresses a specific threat to interpretation. The dimension-specific models examine whether the composite index conceals differences across institutional features that are relevant for Islamic banking. Excluding the six GCC jurisdictions tests whether the result is driven by a group of relatively high-income, oil-exporting economies with large Islamic banking sectors and correlated business cycles. Excluding 2020 and 2021 tests whether the interaction reflects the pandemic, during which relief measures may have affected reported NPF independently of underlying credit quality (IFSB, 2021). Ending the sample in 2023 addresses the thin reporting of the last two years. The balanced-panel estimate addresses the possibility that reporting in later years is related to institutional quality, and the first-difference estimate addresses persistence in the level of NPF. Lagging IQ by one year reduces the possibility that contemporaneous governance perceptions respond to the banking outcome

itself (Bellemare et al., 2017; Reed, 2015). Finally, the alternative outcomes examine whether any moderation extends to profitability and capital, which would be expected if better institutions reduced economic losses rather than only the recorded share of impaired financing (Beck et al., 2013; Čihák & Hesse, 2010).

Formal stationarity tests have limited power in a panel of this length but are reported for transparency. Because the panel is unbalanced, Fisher-type tests that combine country-specific augmented Dickey–Fuller p-values are used instead of the Im–Pesaran–Shin test, which requires a balanced panel (Choi, 2001; Im et al., 2003; Maddala & Wu, 1999). For the 18 jurisdictions with at least eight observations, the tests on the level of NPF are inconclusive, rejecting a unit root in only two of eight variants, whereas a unit root is strongly rejected for the first difference of NPF and for the residuals of the TWFE interaction model ($p < 0.01$ in all variants). As a bounded ratio, NPF cannot follow a unit-root process indefinitely, and the stationarity of the TWFE residuals indicates that the fixed-effects estimates do not reflect a spurious regression (Kao, 1999; Phillips & Moon, 1999). As a further safeguard, Equation (2) is also estimated in first differences.

The interaction coefficient β_3 would have a causal interpretation only under assumptions that the design cannot guarantee, and the estimates are therefore interpreted as conditional associations (Imai & Kim, 2019). Three assumptions are relevant. First, conditional on country effects, year effects and the controls, the growth shortfall must not be correlated with unobserved time-varying determinants of NPF whose influence differs by institutional quality (Balli & Sørensen, 2013; Imai & Kim, 2019). This assumption fails if, for example, better-governed jurisdictions responded to downturns with larger fiscal or prudential relief, because the relief would then load onto the interaction term (Demirgüç-Kunt et al., 2021). Second, institutional quality must not respond to NPF within the year; the lagged-IQ specification reduces, but does not eliminate, this concern (Bellemare et al., 2017; Reed, 2015). Third, measurement error in the WGI estimates and in reported NPF must be unrelated to the shock (Kaufmann et al., 2011). The third assumption is particularly demanding, because NPF is jointly determined by economic losses and by classification and recognition practices, and the aggregate ratio does not allow these two components to be observed separately (Bushman & Williams, 2015; Huizinga & Laeven, 2012). A negative β_3 is therefore compatible with at least three explanations: genuine loss prevention in better-governed systems, more extensive or better-targeted forbearance and relief in those systems during downturns, and differences in the composition of reporting jurisdictions (Demirgüç-Kunt et al., 2021; Peek & Rosengren, 2005). The sensitivity analyses are designed to discriminate among these explanations as far as the data allow, but they cannot rule any of them out conclusively. No coefficient in this study is therefore interpreted as the causal effect of institutions on NPF.

4. Results and Discussion

Table 4 reports the main estimates. In the baseline TWFE model, the IQ coefficient is −6.365, and the growth shortfall carries a positive but imprecise coefficient of 0.191. Within countries, therefore, neither governance nor the growth shortfall is individually associated with NPF at conventional levels once year effects, inflation and income are controlled for. Adding the interaction produces $\beta_3 = -0.219$. The wild cluster bootstrap gives $p = 0.139$. The estimated shock slopes at the 25th and 75th percentiles of IQ are 0.242 and 0.001, respectively, and both marginal-effect confidence intervals include zero. Read literally, a one-percentage-point growth shortfall is associated with an NPF ratio about a quarter of a percentage point higher in a jurisdiction at the lower quartile of institutional quality, and with essentially no change at the upper quartile. The pattern is thus consistent in sign with the shock-absorber view in H1, but it does not provide conventional, small-cluster-robust evidence for it. The small number of clusters also limits statistical power. At a five per cent two-sided level and 80 per cent power, the minimum detectable interaction is approximately 2.8 times its standard error, or about 0.31, which exceeds the point estimate of 0.219 in absolute value. The absence of significance should therefore be read as an absence of conclusive evidence rather than as evidence that institutions play no moderating role.

Two further features of Table 4 deserve comment. First, the negative but imprecise IQ coefficient is typical of fixed-effects estimates of slow-moving governance variables, because WGI scores change little within countries, the within-country variation that identifies β is small relative to the noise in NPF, and the standard error is large. Second, the size of the interaction implies that moving from the lower to the upper quartile of institutional quality, a difference of roughly 1.1 governance units, nearly eliminates the estimated shock slope. In economic terms, at the mean growth shortfall of 1.626 percentage points, the implied difference in NPF between the two quartiles is about 0.4 percentage points, which is modest relative to the sample mean NPF of 4.779%. It corresponds to only about 0.07 standard deviations of NPF and about 0.13 standard deviations of its within-country variation. Even for a growth shortfall one standard deviation above its mean (4.8 percentage points), the implied difference is about 1.2 percentage points, or roughly 0.22 standard deviations of NPF. Even if the point estimate were taken at face value, therefore, the institutional buffer would be economically moderate rather than transformative.

Table 4: Baseline, interaction and dynamic estimates for gross NPF

Estimate	TWFE baseline	TWFE interaction	Dynamic GMM	Interaction GMM
IQ	−6.365 (5.218)	−5.871 (5.132)	−11.162 (29.233)	−0.519 (8.938)
Shock	0.191 (0.148)	0.106 (0.136)	0.868 (2.655)	1.132 (0.970)
Shock × IQ	—	−0.219 (0.112)	—	−0.112 (0.275)
Lagged NPF	—	—	1.026 (0.148)	1.037 (0.151)
Observations / jurisdictions	235 / 25	235 / 25	147 / 24	147 / 24
Instruments	—	—	20	23
AR(1) / AR(2) p-value	—	—	0.372 / 0.857	0.040 / 0.745
Hansen p-value	—	—	0.720	0.774

Notes: Standard errors are in parentheses: country-clustered (CR1) for TWFE and Windmeijer-corrected for system GMM. TWFE models cover 2013–2025 and GMM models 2014–2023. All models include inflation, log real GDP per capita and year effects; country effects enter the TWFE models directly and are removed by transformation in GMM. Because the samples and assumptions differ, GMM and TWFE coefficients should not be compared directly.

The dynamic specification without the interaction yields an estimated persistence coefficient of 1.026 and an insignificant AR (1) test. Because first-differenced errors should display first-order serial correlation when the level errors are serially uncorrelated, the absence of AR (1) correlation, combined with persistence above unity, indicates that this model is not a credible independent confirmation of the baseline. Put simply, a persistence coefficient above one implies that NPF would keep rising after any shock, which is implausible for a bounded ratio, and the missing first-order correlation suggests that the assumptions about the error structure on which the estimator relies are not met. The dynamic interaction model uses 23 instruments for 24 jurisdictions and reports AR (1), AR (2) and Hansen; its interaction coefficient is −0.112, and its estimated persistence is 1.037. With only one instrument fewer than the number of countries, no separate difference-in-Hansen test for the level moments and persistence above unity, the GMM estimate cannot be treated as causal or as evidence of a stable long-run effect. The high Hansen p-values should be read in the light of Roodman's (2009b) warning that instrument proliferation weakens the test, rather than as evidence of instrument validity.

Table 5 shows that the separate regulatory-quality and control-of-corruption interactions are negative, whereas government effectiveness and rule of law are less precise. These four tests are exploratory and vulnerable to multiple comparisons; with four correlated tests, a nominal p-value just below 0.05 carries limited evidential weight. Formally, the Holm (1979) step-down adjustment for the four tests raises the p-values to 0.089 for control of corruption and 0.147 for regulatory quality, and the corresponding wild cluster bootstrap p-values are

0.118 and 0.121. When the four interactions are entered jointly, their variance inflation factors range from about 13 to 26, their signs become mixed and none is individually significant (all $p > 0.20$), which confirms that the data cannot attribute the moderation to a specific dimension. The sensitivity analysis is more informative. Excluding 2020 and 2021 reverses the sign of the aggregate interaction to +0.041, while excluding the six GCC jurisdictions yields −0.186. Ending the sample in 2023 sharpens the estimate to −0.288, indicating that the thin 2024–2025 reporting years do not generate the result. Using one-year lagged governance yields −0.216, close to the contemporaneous estimate. Neither ROA nor CAR shows a comparable interaction. An unrestricted threshold search with 25 country clusters would generate unstable cut-off points and substantial multiple-testing risk, so no threshold effect is claimed. Overall, the aggregate result is closely tied to the pandemic period and to the composition of reporting jurisdictions, and H1 is therefore not supported: the interaction has the negative sign expected under the shock-absorber mechanism, but it is not significant under small-cluster inference and is not stable across samples. H2 receives only partial and exploratory support, because regulatory quality and control of corruption show significant negative interactions only before adjustment for multiple testing.

Two further checks address the unbalanced structure of the panel. First, attrition is not random with respect to governance. The 13 jurisdictions that report NPF in 2024 or 2025 have an average institutional quality of 0.15 over the sample period, compared with −0.67 for the 12 jurisdictions that do not (Welch t-test $p = 0.007$), and their average NPF ratio is lower (3.7% against 7.9%). Second, restricting the estimation to the 14 jurisdictions that report in every year from 2017 to 2023 yields an interaction of −0.077 ($p = 0.391$; wild cluster bootstrap $p = 0.454$), and the first-difference specification yields −0.001 ($p = 0.968$) (Table 5). The negative interaction in the full sample is therefore carried mainly by jurisdictions with shorter or intermittent reporting, which reinforces the conclusion that the apparent institutional buffer depends on sample composition and should not be generalised. Because the first-difference estimator relies only on year-to-year changes in governance, which account for about two per cent of the total variance of IQ, its imprecision also reflects the limited within-country variation.

Table 5: Dimension-specific and sensitivity estimates of the interaction coefficient (TWFE)

TWFE specification	N	G	Interaction β	p-value
Government effectiveness	235	25	−0.127	0.277
Regulatory quality	235	25	−0.247	0.049
Rule of law	235	25	−0.222	0.089
Control of corruption	235	25	−0.248	0.022
Excluding six GCC jurisdictions	167	19	−0.186	0.144
Excluding 2020 and 2021	192	25	0.041	0.806
Sample ending in 2023	213	25	−0.288	0.008
One-year lagged IQ	219	25	−0.216	0.069
Balanced panel, 2017–2023	98	14	−0.077	0.391
First-difference specification	207	25	−0.001	0.968
ROA as outcome	237	24	0.002	0.894
CAR as outcome, excluding Sudan	208	22	−0.018	0.898

Notes: Each row is a separate TWFE interaction model with country-clustered (CR1) standard errors; G = number of country clusters. Dimension-specific coefficients are on the scale of each WGI dimension, and their p-values are not adjusted for multiple testing. The ROA and CAR models use their own samples. Sudan is excluded from the CAR model because its reported CAR of zero is implausible. The balanced panel contains the 14 jurisdictions that report NPF in every year from 2017 to 2023. The first-difference model uses year-to-year changes in all variables for consecutive years only.

The ordering of the dimension-specific coefficients is nevertheless suggestive. Regulatory quality and control of corruption are the dimensions most directly connected to the design of prudential rules and to the integrity of credit decisions, whereas government effectiveness reflects broader administrative capacity and the rule of law reflects judicial enforcement that typically operates with long lags. It is plausible that, over the horizon of a single year, the consistency of regulation and the absence of connected or politically directed financing matter more for the observed NPF response than court-based recovery, which may take several years to be reflected in asset-quality ratios.

In jurisdictions with dual or mixed legal systems, such as Malaysia and the GCC states, recovery can be further delayed when the Shariah compliance of a contract must be referred to a separate authority or reconciled with codified civil law (Hasan & Asutay, 2011). Because these legal arrangements are largely time-invariant, they are absorbed by the country fixed effects, and the WGI rule-of-law estimate does not measure them directly. The relatively strong control-of-corruption interaction is consistent with bank-level evidence that corruption impairs the soundness of Islamic banks (Yunan, 2021; Yunan et al., 2024). The imprecise rule-of-law interaction, by contrast, differs from the finding of Yunan et al. (2024) that law and order is positively associated with Islamic bank health. The difference may reflect the outcome and the measure, that study examines bank-level profitability and soundness using political-risk indicators, whereas the present study examines the response of aggregate NPF to growth shocks using perception-based WGI estimates. Because the four WGI dimensions are strongly correlated across countries and years, however, the data cannot separate

their individual contributions, and the differences between their coefficients are small relative to their standard errors (Langbein & Knack, 2010).

The main finding is that the shock-absorber role of national institutions in Islamic banking is conditional rather than general, consistent with evidence that the stabilising role of institutions depends on the setting in which they operate (Ashraf, 2017; Bermpei et al., 2018). The data are consistent with a weaker NPF response to growth shortfalls in better-governed jurisdictions (Danlami et al., 2022; Yunan, 2021), but this pattern is concentrated in the pandemic years, is sensitive to inference that accounts for the small number of clusters (Cameron et al., 2008; MacKinnon & Webb, 2017) and is not found among consistently reporting jurisdictions and is not corroborated by dynamic estimation. This result speaks to three debates in the literature.

First, it qualifies the conventional finding that growth-shocks are the dominant driver of credit risk. Salas and Saurina (2002), Louzis et al. (2012) and Beck et al. (2015) report a robust negative association between GDP growth and non-performing loans, and Alandejani and Asutay (2017) find macroeconomic sensitivity in GCC Islamic banking. In the present within-country design, the growth shortfall alone does not have a precisely estimated effect on NPF. Two explanations are consistent with the literature. The shortfall measure isolates deviations below each country's recent growth path and is zero in almost half the sample, so it captures fewer but sharper episodes than the level of growth used in earlier studies. In addition, Islamic banking aggregates pool heterogeneous portfolios whose contractual structures, as Abedifar et al. (2013) and Kabir et al. (2015) emphasise, can dampen or delay the recorded response of impairment to income shocks. The weak unconditional shock slope therefore does not contradict the earlier literature; it indicates that the slope is heterogeneous, which is precisely the premise of a moderation analysis.

Second, the negative sign of the interaction is consistent with studies showing that the institutional environment conditions the effect of shocks and regulation on bank soundness (Bermpei et al., 2018; Boudriga et al., 2009). Bermpei et al. (2018) find that institutional quality determines whether supervisory and regulatory tools stabilise banks in emerging and developing economies, and Boudriga et al. (2009) show that the association between supervisory practices and non-performing loans depends on the institutional setting. Danlami et al. (2022) similarly report that institutional quality supports the stability of Islamic banks in six OIC countries. The present results are directionally consistent with these findings, particularly for regulatory quality and control of corruption, which are the dimensions most closely related to prudential rulemaking and to the integrity of credit allocation (Kaufmann et al., 2011; Khwaja & Mian, 2005). Unlike these studies, however, this paper evaluates the moderation with explicit small-cluster inference and period exclusions, and on that basis finds that the evidence is considerably weaker than a conventional significance test would suggest (Cameron & Miller, 2015; MacKinnon & Webb,

2017). The contrast is instructive: bank-level studies with many observations can report precise coefficients even when the effective number of independent institutional environments is small, because governance varies only at the country level (Moulton, 1990). Inference on country-level moderators should therefore be based on the number of countries, not on the number of banks or bank-years (Cameron & Miller, 2015).

Third, the dependence of the result on 2020 and 2021 connects the findings to research on forbearance and loss recognition. The IFSB (2021) observed that relief measures may have flattened reported NPF during the pandemic, and Mansour et al. (2022) show that Islamic banking regulators adopted markedly different combinations of moratoria, restructuring allowances and prudential relief. If jurisdictions with stronger regulatory quality deployed more extensive or better-targeted relief, the negative interaction during the pandemic may reflect policy support and deferred recognition rather than genuinely lower credit losses. This interpretation is consistent with Peek and Rosengren (2005) and Bushman and Williams (2015), who show that forbearance and delayed recognition can mask impairment in the short run while increasing vulnerability later. Nugrohowati and Fakhrunnas (2024) also report that the pandemic affected cross-country Islamic banking performance unevenly. The disappearance of the interaction outside the pandemic years therefore does not show that institutions are irrelevant; it shows that the aggregate data cannot distinguish between loss prevention and policy-induced deferral of loss recognition.

The findings also relate to the long-standing question of whether Islamic banks are inherently more resilient. Čihák and Hesse (2010) and Beck et al. (2013) find that Islamic banks performed comparatively well in some dimensions during the global financial crisis, whereas Bourkhis and Nabi (2013) find no significant difference in soundness. The present evidence suggests that part of this heterogeneity may stem from the environment in which Islamic banks operate rather than from the Islamic business model alone. The GCC exclusion is informative in this respect: removing six relatively high-income and well-governed jurisdictions weakens the interaction only modestly, which implies that the pattern is not driven solely by the Gulf, even though GCC banking systems display distinct NPF dynamics (Alandejani & Asutay, 2017). Yet the loss of precision also indicates that the sample outside the GCC is too small to establish the effect independently. Ashraf (2017) cautions that stronger institutions do not uniformly reduce bank risk, and the absence of any comparable interaction for ROA and CAR in the present analysis is consistent with that caution: institutions may affect what is recorded as impaired without changing profitability or capital buffers in a parallel way.

The results also bear on how governance is conceptualised in Islamic finance. Mollah et al. (2017) and Mergaliyev et al. (2021) show that internal governance and Shariah oversight shape the risk-taking and ethical performance of Islamic banks. National governance indicators, by contrast, are slow-moving

and perception-based and do not directly record the actions of Islamic banking regulators or Shariah boards (Kaufmann et al., 2011). Within-country variation in WGI scores is limited, which reduces the power of fixed-effects designs to detect moderation. The imprecision documented here is thus partly a statement about measurement: it indicates that sector-specific indicators of supervisory practice (Barth et al., 2013), restructuring rules and Shariah governance are needed to identify the channels through which institutions might protect Islamic banking asset quality (Mergaliyev et al., 2021; Mollah et al., 2017).

The dynamic results also carry a methodological lesson for empirical Islamic finance research, in which system GMM is widely applied to short panels. System GMM was developed for panels with many cross-sectional units and few periods (Arellano & Bover, 1995; Blundell & Bond, 1998). When the number of groups is small, the instrument count quickly approaches the number of groups, the Hansen test loses power and coefficient estimates drift towards those of pooled least squares (Roodman, 2009b). The estimated persistence above unity and the insignificant AR (1) test in the present dynamic baseline are symptomatic of these problems. Studies that rely on system GMM with a small number of countries should therefore report the instrument count, collapse instruments and interpret high Hansen p-values cautiously, as recommended by Roodman (2009a). In this study, the fixed-effects results with small-cluster inference provide a more transparent, if less ambitious, basis for inference than dynamic estimates whose identifying assumptions cannot be verified.

Although the evidence does not support a general shock-absorber effect, it has practical implications. Supervisors and standard setters should interpret declines in reported NPF during crisis periods with caution, because such declines may reflect relief measures rather than improved credit quality. Consistent application of the IFSB classification definitions (IFSB, 2019) and disclosure of restructured and moratorium-affected financing would allow analysts to separate prevention from deferral. The exploratory results for regulatory quality and control of corruption, which do not survive adjustment for multiple testing (Holm, 1979), at most indicate that the credibility of prudential rules and the integrity of credit allocation are worth monitoring (Bermpei et al., 2018; Khwaja & Mian, 2005); they do not justify specific institutional interventions or the expectation that governance reforms will by themselves reduce NPF in downturns. From the perspective of maqasid al-Shariah, which emphasises the protection of wealth (hifz al-mal) and fairness in financial dealings, transparent recognition of impaired financing is itself a governance objective, because concealed impairment transfers risk to depositors, investment account holders and the public (Mergaliyev et al., 2021). Finally, the IFSB's effort to harmonise PSIFI reporting is valuable for research, but the uneven coverage of recent years limits cross-country inference and argues for more complete and timely reporting.

The study also has implications for teaching. In Islamic banking and risk-management courses, it illustrates that reported asset quality is the joint product of credit risk and recognition practice, and it provides a worked example of why inference with few country clusters requires small-cluster corrections and why system GMM diagnostics must be checked before dynamic results are reported as robust. The policy recommendations that follow from these implications are summarised in the Conclusion.

5. Conclusion

This study examined whether national institutional quality moderates the association between adverse growth shocks and NPF in Islamic banking systems. Gross NPF was the dependent variable, a growth shortfall measured against each jurisdiction's recent growth path was the independent variable, the composite WGI index and its four dimensions were the moderating variables, and inflation and income were control variables. The analysis used IFSB aggregates for 25 jurisdictions over 2013–2025, estimated two-way fixed-effects interaction models with inference suited to a small number of countries, and used two-step system GMM as a diagnostic complement.

Better-governed jurisdictions tend to show a weaker NPF response to growth shortfalls, but this pattern is imprecise and unstable. It does not survive inference that accounts for the small number of countries, is not confirmed by the dynamic estimates, and disappears when the pandemic years are excluded or when only consistently reporting jurisdictions are retained. Regulatory quality and control of corruption show the clearest negative moderation, although this evidence is exploratory. A universal institutional buffer is therefore not established: the apparent buffer is tied to the COVID-19 period, when forbearance and relief measures were widespread, and to the composition of reporting jurisdictions.

For supervisors, low reported NPF during crises should not be read as evidence of resilience. Standard setters such as the IFSB could ask jurisdictions to report, alongside the gross NPF ratio, the share of restructured and moratorium-affected financing and the classification rules applied during relief periods; national supervisors could publish the timing and scope of forbearance measures; and more complete and timelier PSIFI reporting, especially by jurisdictions with weaker institutions, would reduce selection problems. The study has several limitations. The estimates are conditional associations rather than causal effects, WGI scores are slow-moving and perception-based, missing data are concentrated in later years and in jurisdictions with weaker institutions, the GMM level moments could not be tested separately, and national aggregates conceal differences across banks and contract types. Future research should combine IFSB aggregates with bank-level data in multilevel models, incorporate

jurisdiction-level data on forbearance and restructured financing, and revisit dynamic estimation as longer PSIFI series become available.

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