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Do global uncertainties and financial distress impact Sukuk issuance?

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Abstract

Purpose — This paper examines the impact of major uncertainty indices and global uncertainty on the volume of Sukuk issuance in Türkiye.

Method — The NARDL method is applied to determine the short- and long-term relationships between Türkiye's sukuk issuance and global uncertainty and financial stress indices, capturing both symmetric and asymmetric dimensions.

Findings — Although a symmetric relationship exists between Global Economic Policy Uncertainty (GEPU) and Sukuk issuance, the Financial Stress Index (FSI) has no long-term impact on Sukuk issuance. During periods of global uncertainty, sukuk issuances increase, whereas in conditions of less uncertainty, they fall. There is an inverse relationship between Geopolitical Risk (GPR) and Sukuk issuance. Since all factors affect sukuk issuance in the short run, GEPU has the highest impact. Decreases in the GEPU index positively affect sukuk securities and increase their issuance volumes. Therefore, GPR and GEPU indices have asymmetric effects on sukuk issuances in the short and long term.

Implication — Evidence suggests that sukuk is more resilient to crises than its conventional equivalents. Sukuks are strategically crucial for portfolios and provide sufficient assurance to reduce risk.

Originality — No study has assessed how global financial distress and uncertainty influence Türkiye's sukuk issuance. This study differs from previous studies by focusing on sukuk issuance volumes rather than sukuk yields.

Keywords — Sukuk Issuance, Financial Stress Index, Global Economic Policy Uncertainty, NARDL

Introduction

By their nature and functional structure, financial markets must keep pace with the growth of financial transactions. The world has become nearly a single market due to economic liberalism and integration, leading to a steady increase in global trade and economic activity. This is because globalisation has made national borders obsolete. This circumstance not only promotes the expansion of financial markets and the improvement of social well-being but also causes financial instability and the transfer of adverse consequences across markets. Investors now consider risk and expected return in investment decisions, especially in forming a portfolio. Particularly in portfolio diversification, the low risk of Islamic securities, the reduced portfolio risk and the increase in the daily volume of Islamic securities. This increase in demand for Islamic finance instruments is also reflected in the sukuk issuance volumes. Sukuk issuances are becoming increasingly prominent in the market as a rival to participation banking, especially in developing

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countries like Türkiye. Projects that sukuk issuances would rise by 16.2% cumulatively each year, reaching 1.987 trillion dollars as of 2027 and 2.276 trillion dollars by the end of 2028 (Ratings, 2023). Since the first issuance in 2001, the number of Sukuk issuances has skyrocketed. The growth rate is further accelerated by securities issuance, particularly in the West.

Fluctuations and crises in financial markets are among the main factors that enable the development of Islamic finance. Asutay and Hakim (2018) and Sclip, Dreassi, Miani, and Paltrinieri (2016) state that Islamic finance instruments are strategically crucial for portfolios and provide sufficient assurance to reduce risk, particularly during economic downturns. Evidence indicates that Islamic securities, notably sukuks, are more resilient to crises than their conventional equivalents. It is a fact, nevertheless, that financial crises no longer stay local but spread worldwide. Therefore, a risk emerging anywhere in the world, whether local or global, affects all financial markets and creates spillover effects. In this regard, factors widely recognised as global risk indicators—namely, the financial stress index (FSI), geopolitical risks (GPR), and international economic policy uncertainty (GEPU)—are currently the variables financial markets actively monitor when making decisions. It is critical to monitor these factors closely to identify and implement the necessary actions to improve financial stability and economic resilience.

The present study contributes to the limited literature on the empirical determinants of global uncertainties related to sukuk issuance in Türkiye. However, despite the significance of this topic, there has been relatively little research on it in Türkiye, making it worth examining and generating new empirical evidence that could provide valuable insights for policymakers and investors. The primary purpose of this paper is to investigate how major uncertainty indices affect the size of Sukuk issuance in Türkiye. More specifically, we address the following unanswered questions. Does there exist a relationship between global uncertainty factors and Sukuk issuance? Is dependence symmetric or asymmetric? This study differs from previous studies by focusing on sukuk issuance volumes rather than sukuk yields. This is because investors often seek safe havens over higher yields during periods of global uncertainty. In response to this, Sukuk issuance volumes are expected to increase during uncertain times. Until now, uncertainty indices have overlooked the sukuk market, focusing instead on conventional markets and Islamic stocks. The Sukuk market is expanding annually, so the current study aims to fill this gap. The results of this study are significant for policymakers and sukuk investors as they will provide a clear picture of their investment in Islamic finance.

This study aims to assess the impact of the economic policy uncertainty (EPU) index developed by Baker, Bloom, and Davis (2016), the geopolitical risk (GPR) index introduced by Caldara and Iacoviello (2022), and the FSI index founded by Illing and Liu (2003) on the sukuk issuance of Türkiye. To this end, this section summarises studies that discuss how the three uncertainty indices affect Islamic securities, notably sukuk issuance. Caldara & Iacoviello (2018) introduced the concept of Geopolitical Risk (GPR) to the literature. GPR is a metric that quantifies the uncertainty surrounding wars, military operations, terrorist attacks, and other conflicts. This variable, which mainly causes the deterioration of financial stability, is one of the subjects that researchers have been intensively addressing in recent years. The financial markets are also affected by this process, as the uncertainties that arise when the GPR increases affect the volume of international trade. Consequently, investors' willingness to take risks will wane, resulting in significant losses. Billah and Adnan (2024) examined how geopolitical risks affect the dynamic interconnectedness between sukuk, green sukuk, Islamic stocks, and Islamic green assets. The results demonstrate that Islamic stocks are net volatility receivers, whereas sukuk and green sukuk are short-term volatility spreaders. Long-term results show that the relationship is precisely the reverse. In a distinct study, Billah, Elsayed, and Hadhri (2023) utilised global risk indicators to determine the asymmetric relationship between sukuk and green bonds. The most significant conclusion is that global risk variables do not affect either sukuk or green bonds. In contrast, Balcilar, Bonato, Demirer, and Gupta (2018) investigated the effect of GPRs on Islamic securities and found that GPR had a statistically significant impact on sukuk pricing. To examine the relationship and volatility between global stock markets and sukuks, Sclip et al. (2016) used the dynamic conditional correlations (DCC) model. Results show that sukuk volatility is low throughout the financial crisis. The authors claim that sukuk-containing portfolios offer substantial benefits, particularly during uncertain and crisis-ridden periods. Boukhatem (2022) examined the impact of financial and geopolitical risks on the Saudi Arabian Sukuk market. Financial risk factors encompass exchange rate stability, foreign debt and debt service stability, international liquidity stability, and current account stability. It has been illustrated that the sukuk market is impacted by exchange rate stability, foreign debt stability, and debt service stability.

Recently, many empirical studies have examined the impact of global risk and uncertainty factors on the dynamics of Islamic securities. Notably, GEPU and FSI are among the variables whose effects on Islamic securities and sukuk are investigated in many studies (Ajmi, Hammoudeh, Nguyen, & Sarafrazi, 2014; Kenourgios, Naifar, & Dimitriou, 2016). Naifar, Hammoudeh, and Al-Dohaiman (2016) examined the effects of several risk variables, including FSI and GEPU, on sukuk returns. The findings demonstrate that the sukuk returns of GCC countries are negatively impacted by global risk factors. Additionally, results indicate that GCC sukuk returns are highly susceptible to changes and extremely sensitive to global occurrences. Naifar, Mroua, and Bahloul (2017) explored how local and global uncertainty factors (e.g. VIX and GEPU) affect sukuk and bond returns. The results indicate that sukuk behave differently from bonds in uncertain conditions, which adds them to portfolios and reduces risk. This result suggests that sukuk returns are more resilient to uncertainties. Kasal (2023) used the Bayesian Vector Autoregressive (BVAR) model to examine how FSI affected economic activity and government debt. In conclusion, a rise in FSI will lead to a decline in economic activity, thereby increasing national debt. Ajmi et al. (2014) investigated the interaction between the Dow Jones Islamic Market Index (DJIM) and various financial indexes, including stock, crude oil, and bond indexes, as well as EPU, MOVE, and VIX. The results reveal that all variables, except the DJIM index and VIX, are causally associated. Sukuk performed better than traditional bonds during challenging economic periods, as Billah, Hadhri, Balli, and Sahabuddin (2024) demonstrated, and they recommend that portfolio managers consider this. Similar findings are supported by Billah, Alam, and Hoque (2024), which examines the connection between Islamic markets and green assets during turbulent periods worldwide. The results of Balli, Billah, Balli, and Gregory-Allen (2020), Naifar and Hammoudeh (2016), and Billah, Alam, et al. (2024) indicate that the sukuk market and VIX and GEPU are negatively associated. These findings are also supported by Said and Grassa (2013) and Mirza and Sultana (2020), indicating that it becomes increasingly important, particularly during rising financial uncertainty. In contrast, these results do not align with other studies, such as Billah, Kapar, Hassan, Pezzo, and Rabbani (2024), which includes the financial stress index, and Al-Raeai, Zainol, and Abdul Rahim (2019), which examines political risks.

Apart from these studies, the literature also includes various studies examining sukuk issuances and their performance in uncertain environments. Al-Yahyaee, Mensi, Rehman, Vo, and Kang (2020) find that Islamic stocks outperform conventional stocks during the crisis. The authors proposed that the best way to reduce risk in unstable conditions is to integrate Islamic securities, such as sukuk, into portfolio construction. In a similar vein, Jatmiko, Ebrahim, and Smaoui (2023) probed the relationship between income inequality and Sukuk issuance. The results show a significant positive relationship between income inequality and Sukuk issuance. As a result, Sukuk issuances rise in parallel with increases in income inequality. Godil, Sarwat, Sharif, and Jermsittiparsert (2020) used GEPU and GPR indices to analyse how conventional and Islamic equities behave in bull and bear markets. The impact of oil prices has consistently been shown to cause long-term differences between Islamic and traditional equities during bull markets. Nevertheless, due to GEPU repercussions, Islamic stocks differ from regular stocks during the bear market. As Bouri, Demirer, Gupta, and Marfatia (2019) indicated, GEPU has of greater effect on the volatility of Islamic securities than on their returns. Hasan, Hassan, and Alhomaidi (2023) examined the role of GEPU, GPR, and oil price shocks on sectoral Islamic securities. The results of this study demonstrate that Islamic securities are resilient to GPR and GEPU indices and are effective hedgers. In contrast, they are unable to fend against oil price shocks.

Several empirical studies have examined the effects of geopolitical risk factors, global economic policy uncertainty, and financial stress on equity and bond returns, and, more recently, on the Islamic stock index. This paper contributes to the debate by analysing the impact of global risk and uncertainty on Türkiye's sukuk issuances—an area of the literature that has been largely overlooked. Given previous findings, exploring these factors is well justified. To our knowledge, no study has assessed how global financial distress and uncertainty influence Türkiye's sukuk issuance, a crucial aspect for portfolio strategies optimising risk-return dynamics. This perspective is the paper's originality and provides recommendations to current and potential investors. The principal contribution of this study is its emphasis on sukuk issuance volumes rather than sukuk yields. Determining the extent to which demand for sukuk issuances is influenced by global risk factors will enable the formulation of effective policy measures for countries such as Türkiye, Indonesia, Malaysia, and the Gulf Cooperation Council members that aspire to advance in the field of Islamic finance.

Methods

This study focuses on global uncertainty factors expected to affect Türkiye's sukuk issuance, which has become increasingly popular in Islamic finance. The analyses will be implemented by establishing equation no. 1 below.

$$LnY_{it} = \beta_0 + \beta_1 LnGPR_{it} + \beta_2 LnFSI_{it} + \beta_3 LnGEPU_{it} + \mu_{it} \quad (1)$$

The detrimental effect of financial stress and global uncertainty indices on the Sukuk issuance of Türkiye will be examined using the model developed in Equation 1. To this end, the dependent variable in the equation is expressed as LnY for Sukuk issuance volume. As for the independent variables, Geopolitical Risk is denoted by $LnGPR$, Financial Stress Index by $LnFSI$ and finally, Global Economic Policy Uncertainty by $LnGEPU$. Since the correlation analysis reveals no multicollinearity between the variables, the study used the single model developed for detecting long- and short-term relationships among variables. As indicated in Table 1, sukuk issuance and other variables are gathered from several sources and contain monthly data from 2014 to 2024. In this context, the study's dependent variable, sukuk issuance numbers, is obtained from the database of the Participation Banks Association of Türkiye (TSKB). Sukuk issuance numbers encompass both the private and public sectors, including all sukuk issuances. The Economic Policy Uncertainty website provides data for the Geopolitical Risk Index, while the Federal Reserve Bank supplies data for two independent variables (GEPU and FSI). For data reconciliation, the natural logarithm of all data is computed and analysed. Table 1 lists the variables utilised in the study, including their full names, abbreviations, data sources, data period, and expected signs on the dependent variable.

Table 1. Variables and Explanations

Period	Variables	Expected Sign	Abbreviation	Source
January 2014-September 2024	Sukuk Issuance size of Türkiye	Dependent Variable	LnY	TKBB
	Geopolitical Risk Index	-	LnGPR	EPU
	Financial Stress Index	-	LnFSI	FRED
	Global Policy Uncertainty Index	-	LnGEPU	FRED

Source: TKBB (2025): The Participation Banks Association of Türkiye, EPU: Economic Policy Uncertainty, FRED: Federal Reserve Bank of St. Louis.

The Brock-Dechert-Scheinkman (BDS) test is one of the most essential discriminant tests to perform before determining the long- and short-run relationships between the dependent and independent variables. The BDS test provides necessary evidence on whether the series is linear. Depending on the findings of BDS, linear or nonlinear models can be preferred for cointegration tests between the time series. If this distinction is ignored, there may be situations in which a nonexistent relationship is erroneously treated as if it exists, or a hidden relationship between variables is not revealed. Consequently, linearity analysis is performed to prevent inadequacy in determining the relationship between variables or distrust in the analyses. Thus, Table 2 presents the results of BDS.

Table 2. BDS Test Results

	m=2	m=3	m=4	m=5	m=6
lnY	0.079***	0.179***	0.245***	0.296***	0.328***
lnGPR	0.074***	0.118***	0.146***	0.157***	0.154***
lnFSI	0.025***	0.046***	0.054***	0.055***	0.054***
lnGEPU	0.134***	0.227***	0.288***	0.326***	0.345***

Embedding dimensions = m; *** Significant at 1%.

BDS test results show that the series is not linearly distributed at a 1% significance level. To clarify, the series is not linear. In that instance, models such as the ARDL (Autoregressive Distributed Lag) lose validity and become inefficient at determining the short- and long-term relationships between series. Hence, the NARDL model is the best fit among the alternatives for identifying the relationship in a nonlinear series (Syed, Kamal, & Tripathi, 2021; Göksu, 2024). The NARDL model, just like the ARDL model, must fulfil certain assumptions. The first of these criteria is that none of the variables should be stationary at the second difference *I(2)*. In ARDL and NARDL models, whether the variables are *I(0)* or *I(1)* is essential for the validity and reliability of the analyses. It is preferred that the dependent variable is *I(1)*. The OLS model is established once these assumptions are met, as shown in equation 1. The bounds test (F_{PSS} and t_{BDM}) determines the co-integration relationship in the third stage. Given the presence of a relationship between the variables in the bounds test results, in the fourth stage, the WALD test attempts to identify both short- and long-term asymmetry. The constructed model is verified in the last and fifth steps to see whether it fulfils the assumptions. Moreover, the normality assumption is examined using the Jarque-Bera test, the serial correlation assumption is reviewed using the LM test, and the heteroscedasticity assumptions are analysed using the BPG and ARCH tests. Furthermore, QUSUM tests are used to assess structural breakdowns.

The model initially proposed by Pesaran, Shin, and Smith (2001) and subsequently refined by Shin, Yu, and Greenwood-Nimmo (2014) serves as an illustration; equation 2 was developed to estimate the asymmetric impact of global uncertainty indicators on Türkiye's Sukuk issuance volume.

$$\Delta lnY_{t} = \alpha_0 + \sum_{i=1}^{g=2} \alpha_{1i}\, \Delta lnY_{t-i} + \sum_{a=1}^{h=1} \alpha_{2a}\, \Delta GPR^{+}_{t-i} + \sum_{b=1}^{j=1} \alpha_{3b}\, \Delta GPR^{-}_{t-i} + \sum_{c=1}^{k=0} \alpha_{4c}\, \Delta FSI^{+}_{t-i} + \sum_{d=1}^{l=2} \alpha_{5d}\, \Delta FSI^{-}_{t-i} + \sum_{e=1}^{m=2} \alpha_{6e}\, \Delta lnGEPU^{+}_{t-i} + \sum_{f=1}^{n=1} \alpha_{7f}\, \Delta lnGEPU^{-}_{t-i} + \vartheta lnY_{t-1} + \gamma_1^{+} lnGPR^{+}_{t-1} + \gamma_1^{-} lnGPR^{-}_{t-1} + \gamma_2^{+} lnFSI^{+}_{t-1} + \gamma_2^{-} lnFSI^{-}_{t-1} + \gamma_3^{+} lnGEPU^{+}_{t-1} + \gamma_3^{-} lnGEPU^{-}_{t-1} + \mu_t \quad (2)$$

In this equation, "Δ" shows the primary difference; "μ_t" represents the error term; "g, h, j, k, l, m, n" are the lag orders; "β_0" is the constant; "α_1", "α_2", "α_3", "α_4","α_5", "α_6" "α" are coefficients of the short-run effects; "ϑ", "γ_1", "γ_2", "γ_3" are coefficients of the long-run impacts.

The following hypotheses are examined by using the Wald test to determine long-term asymmetric relationships.

$$"H_0: \frac{\gamma_1^{+}}{-\vartheta} = \frac{\gamma_1^{-}}{-\vartheta};\ H_0 = \frac{\gamma_2^{+}}{-\vartheta} = \frac{\gamma_2^{-}}{-\vartheta};\ H_0 = \frac{\gamma_3^{+}}{-\vartheta} = \frac{\gamma_3^{-}}{-\vartheta}"$$

$$"H_A: \frac{\gamma_1^{+}}{-\vartheta} \neq \frac{\gamma_1^{-}}{-\vartheta};\ H_A = \frac{\gamma_2^{+}}{-\vartheta} \neq \frac{\gamma_2^{-}}{-\vartheta};\ H_A = \frac{\gamma_3^{+}}{-\vartheta} \neq \frac{\gamma_3^{-}}{-\vartheta}"$$

The Wald test indicates that if H_0, it implies the model has long-term asymmetric relationships. Similarly to the long-run, the Wald test is used to evaluate short-run asymmetric relationships.

$$"H_0: \sum_{i=0}^{a} \omega_{1i}^{+} = \sum_{i=0}^{b} \omega_{1i}^{-};\ \sum_{i=0}^{c} \omega_{2i}^{+} = \sum_{i=0}^{d} \omega_{2i}^{-};\ \sum_{i=0}^{e} \omega_{3i}^{+} = \sum_{i=0}^{f} \omega_{3i}^{-}"$$

$$"H_A: \sum_{i=0}^{b} \omega_{1i}^{+} \neq \sum_{i=0}^{c} \omega_{1i}^{-} ; \sum_{i=0}^{d} \omega_{2i}^{+} \neq \sum_{i=0}^{e} \omega_{2i}^{-} ; \sum_{i=0}^{f} \omega_{3i}^{+} \neq \sum_{i=0}^{g} \omega_{3i}^{-} "$$

Depending on the Wald test used, if Ho is rejected and Ha cannot be rejected, there are short-term asymmetric relationships among the model's variables.

Results and Discussion

The descriptive statistics of the study are initially acquired for the sake of analysis. The data set comprises 126 monthly sukuk issuances from January 2014 to September 2024. The low standard deviation, mean, and median are close to one another, suggesting the series may follow a normal distribution. The Jarque-Bera statistic is computed, and the test statistic indicates that the series is normally distributed if the p-value exceeds 0.05. In this context, it is determined that the sukuk issuance numbers and the financial stress index do not exhibit normal distributions.

Table 3. Descriptive Statistics

	Mean	Max.	Min.	Std. Dev.	Skewness	Kurtosis	J-Bera	Prob.	Obs.
lnY	21.787	24.532	16.860	1.671	-0.619	2.473	9.514	0.009	126
lnGPR	4.397	5.120	3.848	0.264	0.193	2.491	2.142	0.343	126
lnFSI	-1.089	1.662	-4.213	0.905	-0.865	4.572	28.678	0.000	126
lnGEPU	5.326	6.080	4.492	0.361	-0.391	2.267	6.025	0.049	126

The high correlation between the series reveals the issue of multicollinearity. This results in relationships that do not exist, or in the inability to recognise or misunderstand relationships that do. Consequently, Spearman's Rank Order is used to conduct the correlation analysis between the variables, and Table 4 presents the findings. These findings indicate slight positive and negative correlations between the dependent and independent variables, and no multicollinearity among the series. This is essential for the general well-being of analysis.

Table 4. Spearman Rank-Order

	lnY	lnGPR	lnFSI	lnGEPU
lnY	1			
lnGPR	0.070	1		
lnFSI	0.237	-0.101	1	
lnGEPU	0.637	-0.052	-0.094	1

As a result of the BDS test, the NARDL model is used after detecting the nonlinear structure. To use this model, some assumptions must be fulfilled. The first and most important assumption is the non-stationarity of any variables at the *I(2)* level. Secondly, the series is expected to exhibit different levels of stationarity, and preferably, the dependent variable is *I(1)*. Thus, unit root tests are performed to detect the stationarity level of each variable. The findings of the unit root test are provided in Table 5, based on the estimation of the constant model and the eighth lag length.

Table 5. Unit Root Test Results

Variables	Level	Prob.	First-difference	Prob.	Decision
lnY	-1.093	0.717	-6.984***	0.000	I(1)
lnGPR	-3.414**	0.012	-	-	I(0)
lnFSI	-7.212***	0.000	-	-	I(0)
lnGEPU	-2.423	0.138	-9.155***	0.000	I(1)

The dependent variable, *lnY*, represents sukuk issuance. In contrast, the independent variable, lnGEPU, reflects uncertainty in global economic policy, which was found to be stationary at the first difference (I(1)) using the ADF unit root test. Test statistics indicate that the results are significant at the 1% level. Meanwhile, the variables *lnFSI* for the financial stress index and *lnGPR* for geopolitical risks are stationary at level *I(0)*.

Table 6. F-Bounds and t-Bounds Test Results

$f\left(lnY \mid lnGPR_t^{+}, lnGPR_t^{-}, lnGEPU_t^{+}, lnGEPU_t^{-}, lnFSI_t \right)$ k:5 m:4				
F PSS: 15.743*** t BDM: -9.647***		10%	5%	1%
F-Statistic	I(0)	2.26	2.62	3.41
	I(1)	3.35	3.79	4.68
t-Statistic	I(0)	-2.57	-2.86	-3.43
	I(1)	-3.86	-4.19	-4.79

Note: k: number of independent variables; m: lag length. (***) Significant at 1%.

After testing of unit roots, the dependent variable is found to be stationary in first differences. Additionally, the results of the BDS test indicate that using non-linear models is now more acceptable. In this sense, the bounds test approach assesses the non-linear cointegration relationship between the variables. The appropriate lag lengths for the series must be determined just before examining the cointegration relationship by the F-bounds and t-bounds tests. According to the Akaike Information Criterion (AIC), the most appropriate lag length for the monthly data set is 4. As shown in Table 6, the FPSS statistic exceeds the upper critical value at the 1% significance level, and the tBDM value is below the lower critical value at the 1% significance level. These results suggest a non-linear cointegration relationship between the variables.

Table 7. Estimation of NARDL Results

A) Long-run results	Coefficient		p-value	t-statistic	
lnY_{t-1}	-1.284***		0.000	-9.647	
$lnGPR_{t-1}^{+}$	-1.128***		0.001	-3.356	
$lnGRP_{t-1}^{-}$	-1.781***		0.000	-4.073	
$lnGEPU_{t-1}^{+}$	0.944		0.007	2.771	
$lnGEPU_{t-1}^{-}$	1.034***		0.030	2.201	
$lnFSI$	0.096		0.411	0.825	
B) Short-run results	Coefficient		p-value	t-statistic	
ECT	-1.284***		0.000	-9.950	
ΔlnY_{t-1}	0.124		0.154	1.437	
$lnGPR^{+}$	1.253**		0.031	2.190	
$lnGPR^{-}$	-2.239***		0.003	-3.096	
$lnGPR_{t-1}^{+}$	0.774		0.208	1.267	
$lnGPR_{t-1}^{-}$	1.535**		0.040	2.086	
$lnGPR_{t-2}^{+}$	1.379**		0.029	2.207	
$lnGPR_{t-2}^{-}$	1.533**		0.037	2.114	
$lnGEPU^{+}$	-0.777		0.229	-1.210	
$lnGEPU^{-}$	3.339***		0.000	4.118	
$\Delta lnFSI$	0.100		0.162	1.409	
$\Delta lnFSI_{t-1}$	-0.133*		0.064	-1.868	
Constant	24.857***		0.000	10.044	
C) Asymmetry tests					
$W_{LR.\,lnGPR}$	7.863**	0.020	$W_{SR.\,lnGR}$	13.535***	0.009
$W_{LR.\,lnGEPU}$	4.672**	0.011	$W_{SR.\,lnGEPU}$	9.343***	0.009
D)Diagnostic tests					
X^2_{SC}	0.967	0.617	$X^2_{HET(BPG)}$	18.475	0.359
$X^2_{NORM(J\text{-}B)}$	0.779	0.678	$X^2_{HET(ARCH)}$	1.854	0.173
CUSUM / CUSUM of Sq.	Stable	Stable	X^2_{FF}	0.008	0.927

Notes: "+" and "-" denote negative and positive partial sums, "X^2_{SC}": Serial correlation; "X^2_{NORM} ": Normality: Jarque-Bera; "X^2_{FF}": Functional form; "$X^2_{HET(BPG)}$ and $X^2_{HET(ARCH)}$ ": Heteroscedasticity; "W_{LR}": Long-run Wald test "W_{SR}": Short-run Wald test. Respectively.

NARDL findings are presented in Panel A of Table 7, which reports long-run estimate results adhering to the determination of the cointegration relationship. All coefficients except the

FSI are statistically significant in the long run. In the long run, a positive shock of 1% positive shock in GPR decreases sukuk issuance by approximately 1.13%. Conversely, a 1% negative shock in GPR increases sukuk issuance by approximately 1.78%. Regarding effect size, the impact of decreases in geopolitical risk on Sukuk issuance is greater than that of increases. This finding might be viewed as suggesting that market participants would be inclined towards sukuk as geopolitical issues decline. In the long run, sukuk issuance rises by around 0.94% for every 1% increase in GEPU and falls by roughly 1.03% for every 1% decrease. Sukuk issuance is considerably more affected by adverse shocks in the global economic policy uncertainty index than positive ones. This finding suggests that market players will shift towards sukuk in response to rising global economic policy uncertainties, and towards other financial instruments and investment tools as international monetary policy uncertainties decrease. In contrast, the financial stress index's long-run coefficient, included linearly in the model, is statistically insignificant. The results of the NARDL model are consistent with the literature, such as Asutay and Hakim (2018); Sclip et al. (2016); Balcılar, Demirer, and Hammoudeh (2015); Bhuiyan, Rahman, Saiti, and Mat Ghani (2018); Hasan, Hassan, Rashid, and Alhenawi (2021); Naifar (2023); Gubareva, Sokolova, Umar, and Vo (2024); Kenourgios et al. (2016), who confirm the findings of the current study, especially for GEPU and GPR.

In Panel B of Table 7, short-run estimation results are presented. In the short run, all variables, except the one-period lagged value of sukuk issuance, are statistically significant at different lag periods. In the short run, reducing global economic policy uncertainty is the most critical factor affecting sukuk issuance. Furthermore, the coefficient of the error correction term is statistically significant and negative. An absolute value of the coefficient greater than one means that there will be an uneven convergence, as stated by Narayan and Smyth (2006). Panel C shows the results of the short- and long-term asymmetry tests. As outlined in the Wald test results, it is concluded that geopolitical risk and global economic policy uncertainty have asymmetric effects in both the short and long run. After considering the diagnostic test findings shown in panel D, the model's functional form structure is appropriate, and there is no problem with heteroscedasticity and autocorrelation. Moreover, the error terms are normally distributed. The CUSUM and CUSUMSQ graphs show that the model parameters are stable and that the short- and long-run coefficients are reliable (Figure 1).

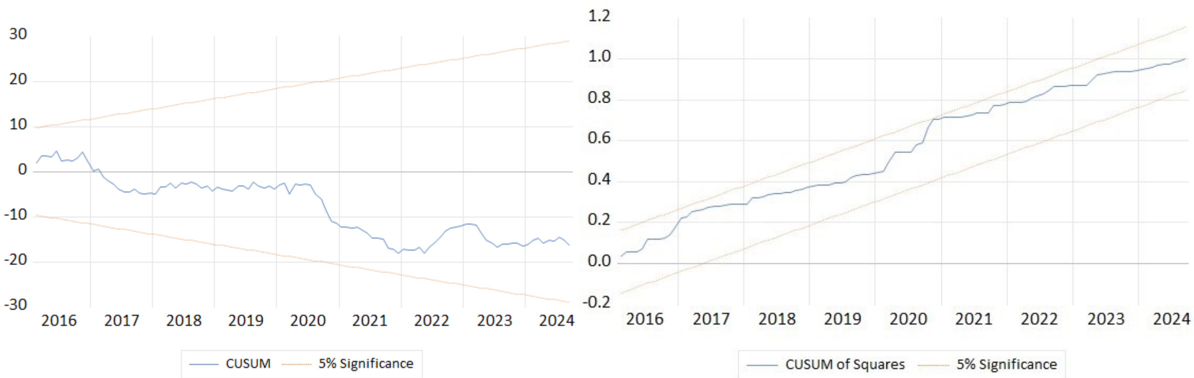

Figure 1: NARDL CUSUM and CUSUM of Squares Graphs

In the last step of the NARDL analysis, Figure 2 below shows the dynamic responses of the dependent variable to one unit of positive and negative shocks in the independent variables, and the formation of the new long-run equilibrium after the shock. Firstly, the left graph in Figure 2 illustrates the response of sukuk issuance to potential shocks arising from geopolitical risks. Specifically, the Sukuk issuance reacts positively to a negative shock of one unit in geopolitical risks. This reaction is much more than the reaction to a positive shock. After about four years, the asymmetric effect of a shock stemming from geopolitical risks on Sukuk issuance ends, and the long-run steady state point is reached.

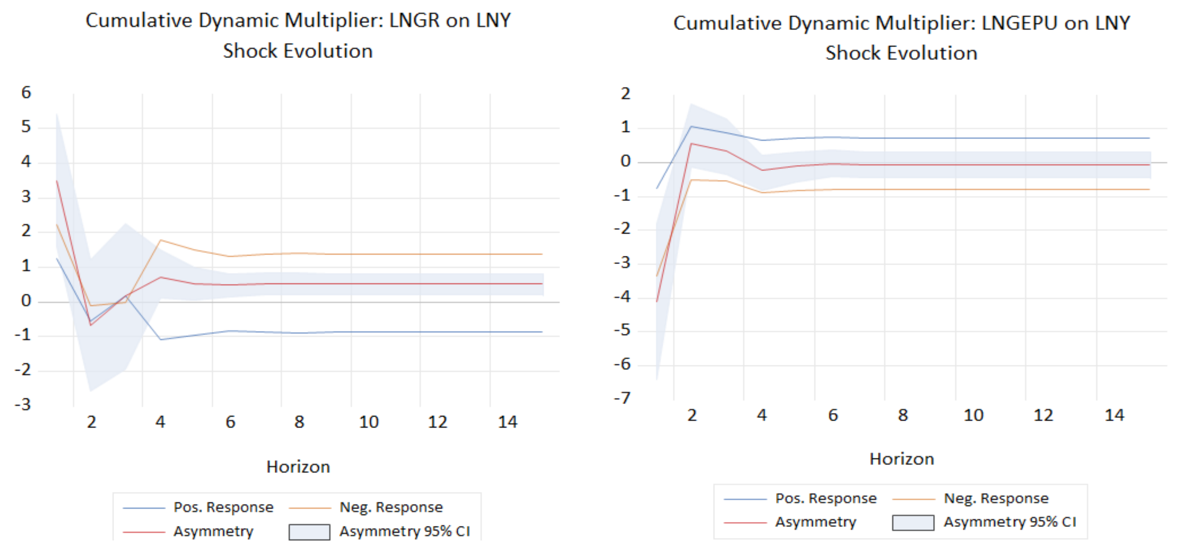

Figure 2: Cumulative Dynamic Shock Evolution

The graph on the right side of Figure 2 shows the response of sukuk issuance to shocks in the global economic policy uncertainty index. The sukuk issuance exhibits a positive response to a 1-unit negative shock caused by the worldwide policy uncertainty index, and this reaction lasted for about 2 years. It then turns negative, and after 4 years, the asymmetric effect of the global policy uncertainty index on Sukuk issuance ends.

Conclusion

This study examined the relationship between the global risk and uncertainty indices FSI, GEPU, and GPR and the sukuk issuance of Türkiye. Monthly totals of all sukuk issuances by Türkiye's governmental and private sectors from January 2014 to September 2024 are examined as a subsection of the study. The NARDL model outputs show short-term and long-term relationships between sukuk issuance and global uncertainty indices. The results demonstrate that the FSI has no long-term impact on sukuk issuances. In contrast, there is a negative relationship between sukuk issuance and GPR. A positive shock of 1% positive shock in GPR reduces sukuk issuance by approximately 1.13%. In contrast, a 1% negative shock in GPR leads to an increase in sukuk issuance of roughly 1.78%. There is a long-run direct relationship between GEPU and Sukuk issuance. A 1% increase in GEPU increases sukuk issuance by approximately 0.94%, while a 1% fall in GEPU decreases sukuk issuance by approximately 1.03%. In general, an assessment of the long-run relationship suggests that lower geopolitical risks increase sukuk issuance. Despite this, increasing GEPU also increases demand for sukuk issuances, as it reinforces the perception of a haven. On the contrary, investors turn to securities other than sukuk when GEPU decreases as their risk appetite increases. In the short run, statistically significant relationships are detected between sukuk issuance and all other variables except the lagged value of Sukuk issuance. The variable with the most critical short-run impact on Sukuk issuance is a decrease in global policy uncertainty. Wald test results also show that GEPU and GPR indices have asymmetric effects in the short and long run. Cumulative Dynamic Shock Evolution reveals that the asymmetric impact of a shock arising from geopolitical risks on Sukuk issuance disappears after 4 years, while the positive response time to a negative shock arising from the global policy uncertainty index is 2 years.

These results provide essential findings for sukuk issuers. Perhaps the most important of these is the decision to determine the timing of Sukuk issuance in response to global uncertainties. The study results suggest that investors tend to switch to safe-haven securities during periods of heightened global uncertainty and financial stress. There will be greater demand for sukuk issuance by issuers, especially during this period of economic and financial turbulence. The fact that companies and governments can access healthier, long-term financing through these issues will also affect future investment opportunities. On the flip side, we do not recommend that companies and governments issue large volumes of sukuk during periods of low global uncertainty and financial turbulence. Due to the high-risk appetite of investors in these periods, there may be a decrease in demand for sukuk issuances. As investors seek alternative instruments for high returns during periods of financial instability, sukuk issuers must offer attractive returns to avoid reduced issuance demand. Accordingly, sukuk-issuing countries, including Türkiye, Indonesia, Malaysia, and the GCC members, as well as other emerging markets, should adopt flexible, countercyclical issuance strategies that account for the asymmetric effects of global uncertainty and geopolitical risk. Specifically, governments and policymakers should strengthen domestic investor bases, diversify sukuk structures toward sustainable and ESG-linked instruments, and maintain transparent policy communication to mitigate the adverse impact of geopolitical shocks. Such measures would enhance market resilience, stabilise issuance volumes, and support the long-term development of Islamic finance across jurisdictions. It is recommended that future studies focus on the volatility spillovers between sukuk issuances and global uncertainty indices. This allows investors to identify significant findings and further support the idea that sukuks are haven assets.

Use of AI tools declaration

The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest

The authors declare no conflicts of interest.

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Quantifying mark-to-market risk in Jamaica's banking sector

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Abstract

Purpose — This study evaluates how well parametric Value-at-Risk (VaR) and Conditional Value-at-Risk (CVaR) models measure market risk from Jamaican banks' sovereign bond exposures.

Method — We calibrate VaR and CVaR models using banks' aggregate portfolio holdings across the entire financial system.

Findings — The parametric VaR model performs reliably, passing standard statistical tests for consistency, independence, and reliability.

Implications — The results suggest that these standard risk measures effectively capture Jamaican banks' market risk exposure to foreign currency-denominated sovereign bonds, which could serve as a helpful tool for regulators to monitor market risk and financial system stability.

Originality — This research applies VaR and CVaR to a novel dataset of Jamaica's entire financial system, demonstrating how regulators can transition from the currently prescribed methods. The findings indicate that these standard risk measures effectively capture risk charges for market risk assessment, as allowed under Basel II, and align with more modern Basel-style frameworks.

Keywords — Value-at-Risk, Conditional VaR, Gaussian distribution, Jamaican bonds, Back testing, Basel II.

Introduction

With the implementation of IFRS 13 in 2013, central banks worldwide face the challenge of maintaining financial sector stability, especially as the country's sovereign bonds—fundamental assets of their banking systems—are prone to sudden, volatile price fluctuations (Bank for International Settlements, 2016). This risk materialised following the twin shocks of the Covid-19 pandemic and the Russia/Ukraine war, resulting in an unexpected surge in global inflation and a significant, synchronised monetary tightening globally. The sharp increases in interest rates worldwide, particularly in US financial markets, led to substantial drops in the prices of fixed-income securities. If not properly hedged, this type of volatility and the significant mark-to-market losses on financial institutions' balance sheets that it causes can also threaten their capital bases, weaken their liquidity, jeopardise their solvency, diminish confidence in the system, and, in the worst cases, trigger full-scale banking crises.

For established banks operating in Jamaica, sovereign Jamaican government bonds are the safest assets in terms of credit risk compared to other Jamaican assets. This creates a challenge. Banks need to hold these assets, but they can sometimes experience significant fluctuations in value. How should central banks evaluate the market risk to the financial system that comes from these bond holdings? Can they estimate the potential range and scale of system-wide mark-to-

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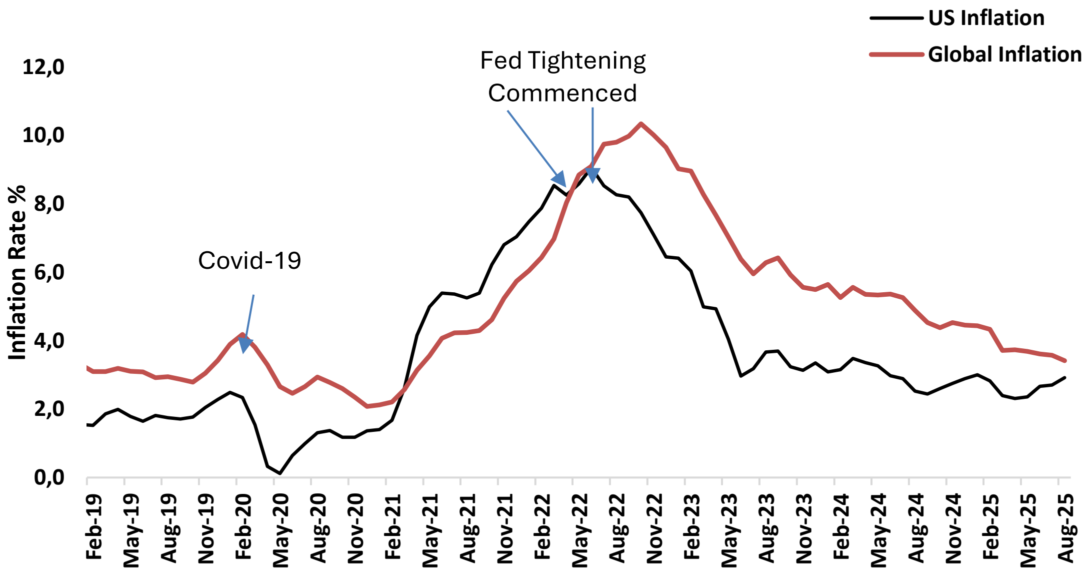
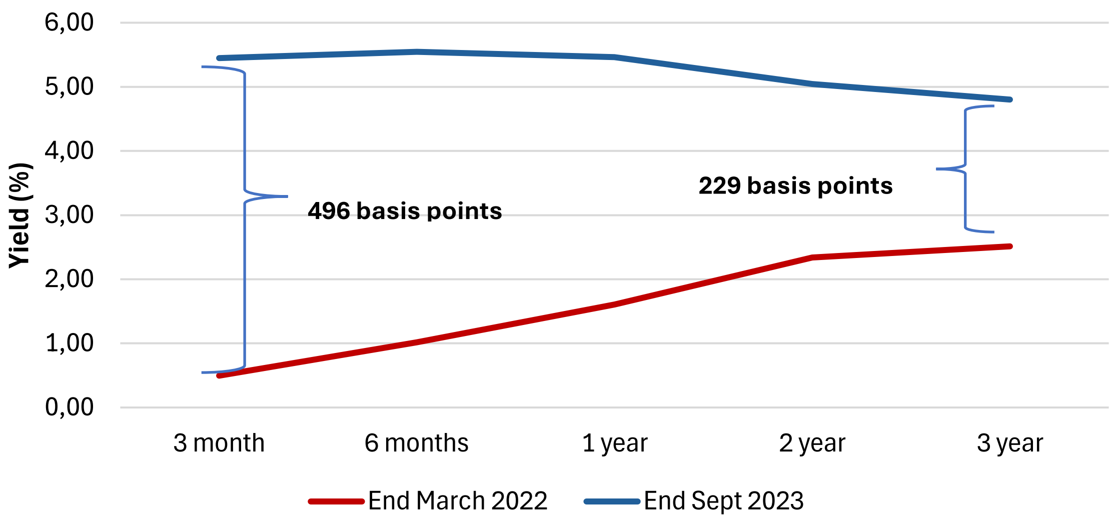
market losses in advance? Are banks sufficiently capitalised to handle this risk? Following the COVID-19 pandemic and the subsequent outbreak of war between Ukraine and Russia, the world experienced a period of high inflation, which prompted central banks around the globe to tighten monetary policy. Annual CPI inflation in the US peaked at 9.1% in June 2022, after gradually rising from 2.5% in January 2020. This pattern closely resembled that of global inflation but showed a slightly sharper increase. The causes of the inflation surge have been well documented (Barnichon et al., 2022; Prokopowicz, 2022).



Source: Bloomberg L.P. (2025).

Figure 1. Global vs US inflation rate (2019-2025)

Although the responses of central banks around the world to the inflation shock were delayed, they were profound. Most notably, the United States Federal Reserve (Fed) initiated an aggressive cycle of raising the target for the federal funds rate. Between March 2022 and July 2023, the Fed raised the federal funds rate by 525 basis points (International Monetary Fund, 2024). This was the fastest rate of interest rate increase since the early 1980s (Justiniano and Barlevy, 2023), and this significant shift affected many countries and regions. In the context of the Fed's policy actions, there was a proportional upward shift at the short end of the US Treasury bond yield curve of six months to three years, as captured in Figure 2.



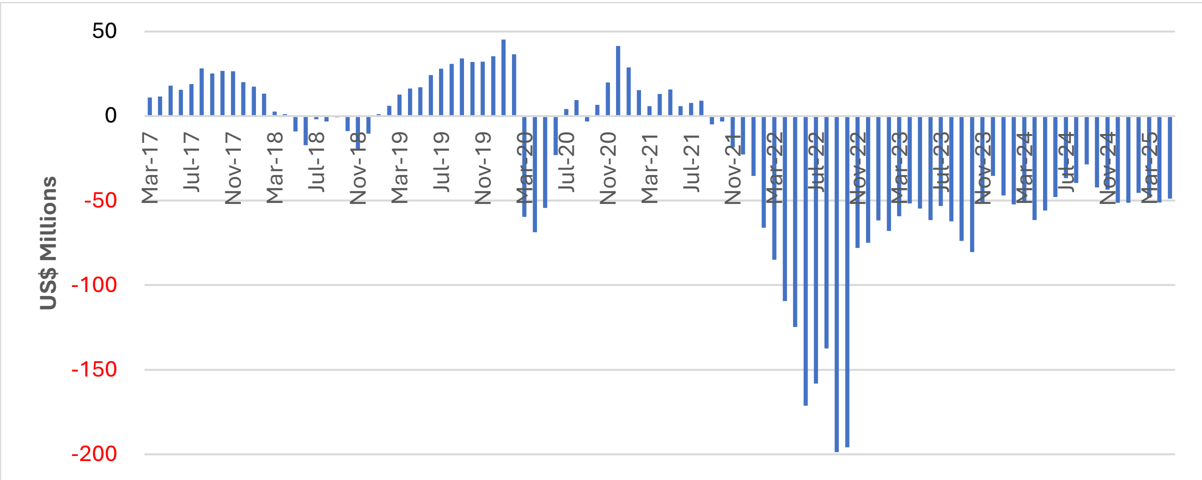
Source: Bloomberg L.P. (2025).

Figure 2. Shift in the US Treasury bill yield curve between March 2022 and September 2023

Before this tightening cycle, the Fed's Monetary Policy Committee began signalling upcoming monetary tightening through its forward guidance, providing a clear and vital signal to portfolio managers. As early as September 2021, the Fed began publicly communicating that future adjustments to the pace of its quantitative easing program were necessary and that it expected progressive interest rate hikes over the next three years (Ihrig & Waller, 2024; Peterson, 2021).

As market interest rates rose from pandemic lows during 2022, financial institutions worldwide recorded significant fair value losses on their balance sheets, the extent of which appeared unanticipated.[1]. For US banks, the value of their fixed-rate assets declined substantially. At the end of 2024, fair value losses, i.e., the value of their available-for-sale portfolios relative to their book values, amounted to US$297 billion, or just over 1.0% of US GDP.

The situation in Jamaica was similar. Jamaican banks hold a significant portion of foreign currency-denominated government bonds as part of their investment activities. In September 2021, six months before the Fed's tightening cycle began, the Jamaican banking system held nearly US$1 billion in USD- and euro-denominated bonds, about 6.8% of Jamaica's gross domestic product. From September 2021 to October 2022, monthly mark-to-market changes ranged from US$62 million in losses to US$13 million in gains, with an average monthly loss just under US$16 million. During this period, cumulative mark-to-market losses increased over the first nine months, reaching a peak of US$198 million in October 2022, equal to 1.0% of GDP or roughly 20.0% of the pre-crisis portfolio value. These banks also experienced mark-to-market losses on their Jamaican-dollar bond holdings, as referenced in the Bank of Jamaica's 2022-2023 report.



Source: Graphs generated by the author using data extracted from the Bank of Jamaica's internal database, which is not publicly available (Access date: July 2025).

Figure 3. Cumulative banking system mark-to-market gains/losses by month (2017-2025)

This level of mark-to-market loss on the safest assets poses a challenge to regulators, raising the question: Could the potential scale of these mark-to-market losses have been quantified in advance? VaR models have been used across various financial markets over the past 30 years for similar purposes, which became a fundamental risk metric (Jorion, 1996). Related papers,

[1] The IMF's Global Financial Stability Report in October 2021 (International Monetary Fund, 2021), for example, highlighted the risk of monetary easing for financial vulnerabilities more than the building market risks associated with monetary policy. Similarly, while US banks publish scenario analyses involving fair value shocks in their annual stress testing disclosures, these scenarios are guided by Supervisory Severely Adverse Scenarios (SSAC) disseminated to the system by the Federal Reserve (Board of Governors of the Federal Reserve System, 2021). The scenario assumptions that were prescribed in early 2021 highlighted the main risks as largely emanating from a protracted global recession (JP Morgan Chase & Co., 2021). The scenario featured economic slowdowns, starting in the first quarter of 2021 among the major developed economies, leading to recessions in the Euro area, the United Kingdom, and Japan (Board of Governors of the Federal Reserve System, 2021). In the context of stable inflation in the US, long-term Treasury rates were projected to rise gradually from 0.30% in the first quarter of 2021 to just 0.90% by the first quarter of 2023 (JP Morgan Chase & Co., 2021).

subsequently, bridged academic and practitioner views and introduced parametric, historical simulation, and Monte Carlo methods of VaR. In a series of papers, the Basel Committee on Banking Supervision enshrined VaR as the global standard for capital adequacy related to market risk, especially under the Internal Models Approach (IMA) (Basel Committee on Banking Supervision, 1996) and (Basel Committee on Banking Supervision, 2004). These papers had a profound influence on bank risk management practices.

Over time, improvements to the VaR framework have been suggested and incorporated into analytical toolkits (Barone-Adesi et al., 1999). In 2001, Danielsson et al. critiqued the VaR methodology, arguing that it can underestimate risk due to model risk, procyclicality, and the failure to capture systemic risk (Danielsson et al., 2001). This paper prompted the work by Acerbi, Carlo, and Tasche (2002), which provided formal proofs that the expected shortfall approach (or Conditional VaR, or CVaR for short) satisfies all coherence axioms and showed how to compute it under various distributional assumptions (Acerbi et al., 2002). In 2016, the Basel Committee on Banking Supervision officially replaced VAR with Expected Shortfall in regulatory capital calculations for market risk under the Fundamental Review of the Trading Book (Bank for International Settlements, 2016).

In the wake of the 2008 financial crisis, stress testing has become a core risk management and regulatory tool in global finance. It has now been formally adopted by the BIS, the IMF, the ECB, the US Federal Reserve, the BoE (with its Annual Stress Tests program), and the EBA (EU-wide stress tests). The first formal application of stress testing emerged when the Fed, in 2009, published its Supervisory Capital Assessment Program, a large-scale, top-down test applied to U.S. banks. Its introduction was novel at the time, as it used a standard set of macroeconomic scenarios and a common forward-looking conceptual framework. Other papers after that one extended the stress testing framework to include more formal transmission of macroeconomic shocks to the banking system (Jobst & Ong, 2016) and to systemically important institutions (Henry & Kok, 2013).

In research on Jamaican bonds, a VaR methodology was developed to identify uncorrelated market risk factors on which the VaR model would be based (Tracey, 2009). This approach eliminated the need to compute covariance matrices, simplifying the calculation of VaR for the portfolios involved. Aside from this paper and White (2009), there is little, if any, published work applying VaR methodologies to Jamaican bonds. This represents a gap that offers an opportunity for further research, including the work described here.

This paper employs parametric Value at Risk (VaR) and Conditional Value at Risk (CVaR) models to estimate the system-wide market risk inherent in the foreign currency-denominated Jamaican sovereign bond portfolio held by the Jamaican banking system prior to the tightening of global financial conditions in early 2022. It also back-tests the predictive power of the VaR models. Its main contribution to the literature is to demonstrate the ability of benchmark models to quantify market risk in the Jamaican setting reasonably. The paper suggests that Gaussian VaR and CVaR models can provide a customised view of market risk and, with further research, offer regulators the option to update the current fixed risk-weighting system to better reflect market dynamics.

Methods

Value-at-risk (VaR) models estimate the minimum potential loss in value over a fixed period that the owner of a financial asset should experience with a specified probability, assuming the asset is held over that period. It is crucial to bear in mind that VaR provides only an estimate, and VaR calculations say nothing about the distribution of potential losses when losses exceed the VaR estimate.

There are three types of VaR models: 1) parametric VaR where financial asset returns or changes in the financial assets' risk factors are assumed to follow a particular probability distribution function, 2) historical VaR, where the VaR model is estimated entirely from the actual historical distribution of actual past asset prices or risk factor behavior and does not assume that any particular parametric probability distribution can describe these asset prices or risk factors, and 3) simulation VaR where the model is computed by simulating asset price or risk factor behavior. The simulation is then run many times, and the distribution of simulated losses is used to estimate VaR.

Model Framework

We now set out, from first principles, the theoretical model framework used for the application of parametric multivariate Gaussian VaR and CVaR to our dataset. We begin with the assumption that the price of a bond, B, that matures at the time T is a function of its yield to maturity y. That is,

$$B \equiv B(y;T)\ y \geq 0, T > 0$$

We further assume that the bond pays an annual coupon C and at maturity pays the principal P. Then, the price of the bond can be calculated as

$$B(y;T) = \sum_{t=1}^{T} \frac{C}{(1+y)^t} + \frac{P}{(1+y)^T} \tag{1}$$

The Taylor series expansion of $B(y;T)$ gives

$$B(y + \Delta y) = B(y) + \frac{B'(y)\Delta y}{1!} + \frac{B"(y)(\Delta y)^2}{2!} + \cdots + \frac{B^n(y)(\Delta y)^n}{n!} + \cdots \tag{2}$$

Ignoring the terms in $(\nabla y)^n$, for $n \geq 2$, given that these terms will be small for small ∇y, we obtain a first-order Taylor expansion that approximates how $B(y;T)$ changes for small changes in y:

$$B(y + \nabla y) \approx B(y) + B'(y)\Delta y + \cdots \tag{3}$$

Where $B'(y)$ is the first derivative of B with respect to y and is given by

$$B'(y) = \sum_{t=1}^{T} -\frac{tC}{(1+y)^{t+1}} - \frac{TP}{(1+y)^{T+1}} \tag{4}$$

We can rearrange Equation (3) and divide both sides by $B(y)$ to yield the following first-order approximation:

$$\frac{B(y+\Delta y) - B(y)}{B(y)} \approx \frac{B'(y)}{B(y)}\Delta y \tag{5}$$

That is, we have an expression for the decimal form of the percentage change in bond price given slight changes in yield. The quantity $\frac{B'(y)}{B(y)}$ is known in the fixed income literature as the negative modified duration, D^*, of the bond B. That is,

$$D^* = -\frac{B'(y)}{B(y)} \tag{6}$$

Incorporating the definition of D^*, in Equation (5), allows us to express the following first-order approximations:

$$\frac{\Delta B(y)}{B(y)} \approx D^*\Delta y \tag{7}$$

$$\Delta B(y) \approx D^* B(y)\Delta y \tag{8}$$

as stated in (Jorion, 1996) where $\Delta B(y)$ is the change in bond price due to a slight change in yield, i.e. $\Delta B(y) = B(y + \Delta y) - B(y)$. By multiplying the result in Equation (7) by 100%, we obtain a first-order approximation for the percentage change in the price of a bond due to small changes in yield, expressed in terms of the modified duration of cap D to the asterisk operator. Similarly, Equation (8) gives us a first-order approximation for the dollar change in the price of a bond, ΔB *, due to small changes in yield cap delta y, where cap delta y* is expressed as a decimal. If, as is the practice case, Δy is expressed in units of basis points, and since one basis point = 0.0001, Equation (8), which provides the dollar change in the bond price, becomes:

$$\Delta B(y) \approx \left[\frac{-D^* B(y)}{10{,}000}\right].\Delta y \tag{9}$$

The expression in square brackets is often referred to as the dollar value of a basis point (DV01) and represents the change in a bond's price for a one-basis-point change in its yield.

Estimating VaR for a single bond

Daily changes in Jamaican bond yields are influenced by many, small random events e.g. 1) perceptions of the creditworthiness of the issuer, 2) global events such as decisions of the Federal Reserve, oil price changes, geopolitical events 3) macro news such as inflation, unemployment and economic growth data and 4) events within local Jamaican and global financial markets relating to sentiment and liquidity among other factors. Each of these factors and their component sub-factors contributes a small part to the overall daily change in bond yield. These factors can be thought of as independent or weakly dependent random variables, and the daily change in bond yields is therefore a summation of these many small factors. The Central Limit Theorem informs us that when a variable is the sum of many independent random variables or weakly dependent factors, its distribution approximates the normal distribution.

This provides some basis for us to assume that actual daily changes in bond yields, ϵ, expressed in basis points (i.e., one basis point = 1/100th of a percentage point), are generally distributed with a mean μ_ϵ and variance σ_ϵ^2. That is,

$$\epsilon \sim N(\mu_\epsilon, \sigma_\epsilon^2) \tag{10}$$

We are therefore interested in leveraging this Gaussian assumption of daily bond yield changes to characterise daily changes in bond prices ΔB given changes in yield ϵ. Given the linear relationship between ΔB and small changes in yield in Equation (9), coupled with the Gaussian assumption in (10), we can deduce the variance (Var) of ΔB which we denote as σ_B^2:

$$\sigma_B^2 = Var[\Delta B(y)] \approx Var\left[\frac{-D^*B(y)}{10{,}000} \, . \epsilon\right] = \left[\frac{D^*B(y)}{10{,}000}\right]^2 \sigma_\epsilon^2 \tag{11}$$

The standard deviation of bond price changes is given by

$$\sigma_B = \left[\frac{D^*B(y)}{10{,}000}\right] \sigma_\epsilon \tag{12}$$

Providing a linear relationship between the standard deviation of daily bond price changes and the standard deviation of daily yield changes. Similarly, the expected daily bond price change, μ_B , is given by:

$$\mu_B = E[\Delta B(y)] \approx E\left[\frac{-D^*B(y)}{10{,}000} \, . \epsilon\right] = \left[\frac{-D^*B(y)}{10{,}000}\right] \mu_\epsilon \tag{13}$$

Therefore, from our assumption in (10), it follows that daily bond price changes would also be normally distributed with a mean μ_B and standard deviation σ_B. That is,

$$\Delta B \sim N(\mu_B, \sigma_B^2) \tag{14}$$

We further leverage the properties of the normal distribution, which imply that daily changes in the price of bonds ΔB should observe the probabilistic relationship

$$P(\Delta B \geq \mu_B + Z_{1-\alpha}\sigma_B) = \alpha \tag{15}$$

where $0 < \alpha < 1$ and $Z_{1-\alpha}$ is the $(1-\alpha)$ quantile of the standard normal distribution, and we refer to the quantity $\mu_B + Z_{1-\alpha}\sigma_B$ as the $VaR_{\alpha.100\%}$ estimate with respect to the bond.
From Equation (12), therefore, $VaR_{\alpha.100\%}$ is given by

$$VaR_{\alpha.100\%} = \left[\frac{-D^*.B(y)}{10{,}000}\right]\mu_\epsilon + Z_{1-\alpha}\left[\frac{D^*.B(y)}{10{,}000}\right]\sigma_\epsilon \tag{16}$$

Equation (16) gives us the daily Value at Risk at $\alpha.100\%$ confidence level. To obtain the corresponding Value at Risk for monthly, quarterly, semi-annual, and annual intervals, we use a property of the normal distribution, which states that if ϵ_i are independent and identically normally distributed random variables where $\epsilon_i \sim N(\mu, \sigma^2)$ for $i = 1,2,\cdots,T$ then $\sum_{i=1}^{T}\epsilon_i \sim N(\mu, \sigma^2 T)$. We thereby scale the daily Value at Risk by $\sqrt{T}$ where T is now the number of trading days in the interval, we assume that there are, on average, 21 trading days in a month.

Estimating VaR for a portfolio of bonds

Now, assume we have a portfolio of n bonds $B_1, B_2, \cdots, B_n$ where each bond is a function of its yield y_i and matures at the time T_i. That is,

$B_i \equiv B_i(y_i; T_i)\ y_i \geq 0, T_i > 0$ for i = 1, 2, $\cdots, n$

Furthermore, assume each bond B_i pays the annual coupon C_i and at maturity pays the principal P_i. Then, from (1), the price of each bond B_i for $i = 1,2,\cdots,n$ can be calculated as

$$B_i(y_i; T_i) = \sum_{t=1}^{T_i} \frac{C_i}{(1+y)^t} + \frac{P_i}{(1+y)^{T_i}} \tag{17}$$

We further assume that k_i units of each bond B_i are held in the portfolio such that the value of the portfolio V is given by:

$$V = k_1B_1(y_1) + k_2B_2(y_2) + \cdots + k_nB_n(y_n) \tag{18}$$

where $V: \mathbb{R}^n \rightarrow \mathbb{R}$ is a multivariate function that takes the n-tuple $(y_1, y_2, \cdots, y_n)$ as input and returns the scalar value of the portfolio. We define

$$w_i = \frac{k_i B_i(y_i)}{V} \tag{19}$$

for $i = 1, 2, \cdots, n$ where each w_i represents the weight of the holdings of the bond B_i in the portfolio V and the matrix $\boldsymbol{W}$ is given by the diagonal matrix of the w_i:

$$\boldsymbol{W} = \begin{bmatrix} w_1 & & \\ & \ddots & \\ & & w_n \end{bmatrix} \tag{20}$$

Now, the first-order Taylor series expansion of $V(y_1, y_2, \cdots, y_n)$ yields:

$$V(y_1 + \Delta y_1, y_2 + \Delta y_2, \cdots, y_n + \Delta y_n) = V(y_1, y_2, \cdots, y_n) + \sum_{i=1}^{n} \frac{\partial V}{\partial y_i} \Delta y_i + \cdots \tag{21}$$

Where we ignore higher-order terms in $\Delta y_i^k, k > 1$, $\Delta y_i y_j$, and beyond. A rearrangement of Equation (21) gives us the first-order approximation of the change in V, denoted ΔV, that results from a change in yields $y_1, y_2, \cdots, y_n$ by the small amounts $\Delta y_1, \Delta y_2, \cdots, \Delta y_n$

$$\Delta V \approx \sum_{i=1}^{n} \frac{\partial V}{\partial y_i} \Delta y_i \tag{22}$$

From Equations (18), (6), and (19):

$$\frac{\partial V}{\partial y_i} = k_i B'(y_i) = -k_i D_i^* B(y_i) = -w_i D_i^* V \tag{23}$$

where D_i^* is the modified duration of the ith bond. So, again, if the changes in yield Δy_i are all measured in basis points, as opposed to decimals, ΔV will be given by:

$$\Delta V \approx \frac{V}{10{,}000} \sum_{i=1}^{n} -w_i D_i^* \Delta y_i \tag{24}$$

Like the previous section, we assume that actual changes in the column vector of bond yields.

$$\boldsymbol{y} = \begin{bmatrix} y_1 \\ \vdots \\ y_n \end{bmatrix} \tag{25}$$

are given by

$$\boldsymbol{\epsilon} = \begin{bmatrix} \epsilon_1 \\ \vdots \\ \epsilon_n \end{bmatrix} \tag{26}$$

which are characterised by a multivariate normal distribution of the mean $\boldsymbol{\mu_\epsilon}$ with a variance-covariance matrix $\boldsymbol{\Sigma}$

$$\boldsymbol{\mu_\epsilon} = \begin{bmatrix} \mu_{\epsilon_1} \\ \vdots \\ \mu_{\epsilon_n} \end{bmatrix} \tag{27}$$

and

$$\boldsymbol{\Sigma} = \begin{bmatrix} \sigma_{11}^2 \sigma_{12} \cdots & \sigma_{1n} \\ \sigma_{21} \sigma_{22}^2 & \vdots \\ \vdots \quad \vdots \quad \ddots & \\ \sigma_{n1} \cdots & \sigma_{nn}^2 \end{bmatrix} \tag{28}$$

where σ_{ii}^2 represents the variance of the daily yield changes ϵ_i and σ_{ij} represents the covariance between daily yield changes ϵ_i and ϵ_j. Using equation (24) with ϵ_i substituted for Δy_i , we can now derive the expected value μ_V, variance σ_V^2, and standard deviations σ_V of daily portfolio changes as follows:

$$\mu_V = E[\Delta V] = \frac{V}{10{,}000} \sum_{i=1}^{n} -w_i D_i^* \mu_{\epsilon_i} \tag{29}$$

$$\sigma_V^2 = Var[\Delta V] = \left[\frac{V}{10{,}000}\right]^2 \left(\boldsymbol{D^*}^T \boldsymbol{W}\right)\boldsymbol{\Sigma}(\boldsymbol{W}^T \boldsymbol{D^*}) \tag{30}$$

where $\boldsymbol{D^*}$ is the column vector of modified durations, D_i^*, of each bond. We can further deduce that, due to the properties of the normal distribution:

$$\Delta V \sim N(\mu_V, \sigma_V^2) \tag{31}$$

which allows us to express the daily Value at Risk for the change in portfolio value, ΔV, at $\alpha.100\%$ confidence, where $0 < \alpha < 1$, as

$$VaR_{\alpha.100\%} = \frac{V}{10{,}000}\left[-\left(\boldsymbol{D^*}^T \boldsymbol{W}\boldsymbol{\mu_\epsilon}\right) + Z_{1-\alpha}\sqrt{\left(\boldsymbol{D^*}^T \boldsymbol{W}\right)\boldsymbol{\Sigma}(\boldsymbol{W}^T \boldsymbol{D^*})}\right] \tag{32}$$

Estimating CVaR for a portfolio of bonds

Value-at-risk models suffer from the inadequacy that, while they quantify a loss threshold below which losses should be confined with specific probability, they do not inform us of the expected value of losses should this threshold be breached (Rockafellar & Uryasev, 2000; Rockafellar & Uryasev, 2002). If a tail event occurs, we can lose more than the threshold itself; however, VaR does not provide guidance on the potential extent of this loss. Conditional Value at Risk (CVaR), which measures the average loss conditional on the loss threshold being breached, has been proposed as a solution to the VaR insufficiency. CVaR measures the magnitude of potential losses when the VaR loss threshold is breached (Rockafellar & Uryasev, 2000; Rockafellar & Uryasev, 2002).

We are interested in calculating the Conditional Value-at-Risk for the daily change in the value of our portfolio of bonds, that is $\Delta V \sim N(\mu_V, \sigma_V^2)$. We express the Conditional Value-at-Risk of the daily change in our portfolio's value, at the $\alpha.100\%$ confidence level, as:

$$CVaR_{\alpha.100\%} = E[\Delta V | \Delta V \geq VaR_{\alpha.100\%}] \tag{33}$$

The general closed-form solution for Conditional Value at Risk at the $\alpha.100\%$ confidence level under Gaussian assumptions is derived in (Khokhlov, 2016) given by:

$$CVaR_{\alpha.100\%} = \mu_V - \frac{e^{\frac{-Z_{(1-\alpha)}^2}{2}}}{\alpha\sqrt{2\pi}} \sigma_V \tag{34}$$

where $0 < \alpha < 1$ and $Z_{1-\alpha}$ is the $(1-\alpha)$ quantile of the standard normal distribution.

Data and Calculations

We used daily market price data from 28 July 2015 to 31 August 2021 for all of seven US$-denominated Jamaican sovereign bonds that were outstanding during this period. This represents the most extended period for which daily data are available for all seven bonds, since two of the bonds were issued at the beginning of this period, i.e., on 28 July 2015. The seven bonds had/have maturity dates of: 15 January 2022, 17 October 2025, 28 April 2028, 28 February 2036, 15 March 2039, and 28 July 2045. For each of the bonds, over the six years for which we gathered price data, we calculated daily bond yields, the daily changes in bond yields, the standard deviation of these daily changes, as well as the modified duration of each bond at September 2021, December 2021, and March 2022. We also calculated the variance-covariance matrix relating to the seven daily yield change data series.

From bank filings, we then calculated the amounts in which these sovereign dollar bonds were held, in aggregate, by the banking system as at September 2021, December 2021, and March 2022, as a proportion of the banking system's holdings of these bonds. This provided the entries for our diagonal matrix **W**, in equation (20), for each of these dates. One of the assumptions underlying the VaR methodology is that holdings remain constant throughout the forecast period. As would be expected, banks sometimes buy and sell these bonds in response to liquidity and other considerations. Therefore, a particular institution's holdings of these bonds will change over time. However, outside of a bond maturity, removing a bond entirely from the system, the deposit-taking financial system's aggregate holdings of US dollar-denominated sovereign Jamaican dollar bonds did not change significantly over the period.

We then employed Equation (32) to compute the daily VaR and the monthly VaR (by scaling the daily VaR as appropriate) at the 95% and 99% confidence levels, respectively, as of September 2021, December 2021, and March 2022. Using equation (34), we calculated the daily and monthly Conditional Value at Risk at the respective 95% and 99% confidence levels for September 2021, December 2021, and March 2022. All these VaR metrics were calculated using bond price data through August 31, 2021. This allows us to ascertain the risk of loss that could have been quantified in advance of the mark-to-market losses that eventually ensued.

The final element of our framework involved the generation of daily portfolio values. We have monthly changes in the value of the banking systems' portfolio of US dollar-denominated sovereign bonds as reported by the banks and aggregated by the central bank. However, this limits us to just 45 monthly observations in the out-of-sample period from October 2021 to May 2025. While banks do not report daily portfolio values, we can construct a close approximation of these values as we have the weights in which the bonds were held on three dates, we know the daily prices of the bonds, and we know the size of the overall portfolio at the beginning of the out-of-sample period. Therefore, to complement our analysis and mitigate concerns about small-sample bias, we construct daily portfolio values by assuming that the March 2022 weights in Table 2 hold through to the end of the period and use the historical bond price data to simulate the daily value of the financial system's portfolio of foreign currency denominated Jamaican sovereign bonds.

Backtesting

Backtesting is an essential step in validating the accuracy and reliability of VaR models. The three standard tests employed in the literature, and which are briefly reviewed below, are the Kupiec Test, Christoffersen's Markov Test, and Christoffersen's Conditional Convergence Test. The Kupiec Test, also called the Proportion of Failures Test or the Unconditional Coverage Test, is one of the most frequently used VaR backtesting techniques. It evaluates whether the observed frequency of VaR breaches is consistent with the model's expected frequency and uses a likelihood ratio to do so (Kupiec, 1995).

The Kupiec Test statistic is given by:

$$LR_{uc} = -2\ln\left[\frac{(1-p)^{N-x}p^{x}}{\left(1-\frac{x}{N}\right)^{N-x}\left(\frac{x}{N}\right)^{x}}\right] \tag{35}$$

where N is the number of observations, x is the number of exceptions where the loss exceeds the VaR estimate, and p is the VaR confidence level (e.g., 0.95). The test follows the $\chi2(1)$ distribution under the null hypothesis that the observed exceptions are consistent with the model.

The Christoffersen's Markov Test examines whether VaR exceptions are independently distributed using a Markov framework. The test checks whether analysing sequences of exception clusters and evaluates whether an exception occurs in the current period, given what happened last period (Christoffersen, 1998). The Christoffersen's Markov test defines the following:

$$\pi_{01} = \frac{n_{01}}{n_{00}+n_{01}}, \pi_{11} = \frac{n_{11}}{n_{10}+n_{11}}, \pi = \frac{n_{01}+n_{11}}{n_{00}+n_{01}+n_{10}+n_{11}} \tag{36}$$

where 1) n_{00} is the number of times that a month, where the observed mark-to-market loss is less than the VaR threshold (i.e., no VaR exception), is followed by a similar month 2) n_{01} is the number of times that a month without a VaR exception is followed by a month with a VaR exception, 3) n_{10} is the number of times that a month with a VaR exception is followed by a month without a VaR exception, and 4) n_{11}. This refers to the number of consecutive months with a VaR exception.

Then the Likelihood under Markov is computed as

$$L_1 = (1-\pi_{01})^{n_{00}} \left(\pi_{01}^{n_{01}}\right)(1-\pi_{11})^{n_{10}} \left(\pi_{11}^{n_{11}}\right) \tag{37}$$

The Likelihood under independence is computed as:

$$L_0 = (1-\pi)^{(n_{00}+n_{10})}\left(\pi^{(n_{01}+n_{11})}\right) \tag{38}$$

The Christoffersen's Markov Test combines these likelihoods to compute:

$$LR_{ind} = -2\ln\left(\frac{L_0}{L_1}\right) \tag{39}$$

which follows the $\chi^2(1)$ distribution under the null hypothesis that the VaR exceptions are independent over time.

The Christoffersen's Conditional Convergence Test is a joint test, both that the observed VaR exceptions are consistent with model expectations and that VaR exceptions are independent (Christofferson, 1998). It combines the previous test statistics LR_{uc} and LR_{ind} as follows:

$$LR_{cc} = LR_{uc} + LR_{ind} \tag{40}$$

where $LR_{cc} \sim\chi^2(2)$ under the null hypothesis that the observed VaR exceptions are both model consistent and independent.

Results and Discussion

The primary objective of this paper is to use historical price data prior to September 2021 to assess the efficacy of multivariate Gaussian VaR and CVaR models in estimating the risk of loss due of the severe monetary tightening that followed in 2022/23. In this regard, this objective differs from that of other papers, such as Omari (2017), Swami (2016), Abad et al. (2013), Rossignolo, Duygun, & Shaban (2012), and Vlaar (2000), where the relative efficacy of competing VaR techniques is compared. It also differs from papers such as those by Obadović et al. (2016), which assess the accuracy of a particular method. However, the relative performance of a variety of VaR models was applied to emerging financial markets, and their performance was evaluated over the period of the global financial crisis (Miletic & Miletic, 2015; Žiković & Aktan, 2009). The similarity in intent of these papers with ours lies in the assessment of the efficacy of a VaR methodology during a period of market distress. We have been motivated by the Jamaican experience and seek to answer the specific question of whether the heavy system-wide mark-to-market losses could have been anticipated using the Gaussian VaR/CVaR methodology, with the system-wide data available up to the end of August 2021.

The descriptive statistics of changes in daily bond yields, over the period 28 July 2015 to 31 August 2021, are shown in Table 1, and the corresponding frequency distribution is shown in

Figure 4. Of the 1,590 values of daily yield changes for each bond, 98% of these daily changes, i.e., everything between the 1st and 99th percentiles, lie well within -20 and 20 basis points. However, the daily yield changes for each bond are characterised by a few extreme movements, tens of standard deviations away from the mean, as indicated by the minimum and maximum daily yield changes for each bond. The extreme kurtosis and excess skew are clear signs of the non-normality of these daily bond yield changes. In Lau (2012), descriptive statistics are provided on sovereign bonds issued by 11 different European governments over the period 1999 - 2012. All these bond return series exhibit extreme kurtosis. Greek and Portuguese bonds exhibit kurtosis levels similar to those of Jamaican bonds. In addition, Omran & Semnkova (2019) examine returns on Russian government and corporate bond indices, finding excess kurtosis like the levels reported here, with conclusions of non-normality.

Table 1. Descriptive statistics of daily bond yield change by bond maturity (basis points)

Maturity	Jan'22	Jul '25	Oct '25	Apr '28	Feb '36	Mar'39	Jul '45
1st percentile	−7.87	−12.24	−17.63	−14.41	−14.19	−12.11	−12.28
5th percentile	−0.51	−5.67	−6.43	−7.08	−5.24	−5.73	−5.50
95th percentile	0.00	4.36	5.24	6.12	4.37	5.42	4.87
99th percentile	10.25	11.34	16.99	12.45	10.59	11.92	10.64
Minimum	−43.16	−34.06	−59.94	−108.11	−273.15	−36.59	−35.29
Maximum	24.02	86.57	296.37	98.93	508.40	88.92	76.81
Mean	−0.08	−0.15	−0.16	−0.18	−0.14	−0.14	−0.19
Std. Deviation	3.04	5.02	9.12	6.93	18.34	5.65	5.12
Skew	−3.51	6.24	21.42	1.14	11.30	5.64	5.56
Kurtosis	69.95	97.30	705.43	91.63	450.45	84.93	81.9

This aligns with the findings of Gabriel (2014) and Gabriel & Lau (2012), who found that European bond returns also exhibit excess skewness and kurtosis that are incompatible with the normal distribution. In (Bauer & Chernov, 2024), non-normality and skewness in US Treasury bond yields are well documented and explained in terms of risk premia. Much earlier, it was found that the distribution of daily changes in short-term interest rates in the United Kingdom also exhibited non-normal characteristics with high peaks and fat tails (Robinson & Taylor, 1993). Regarding Jamaican asset prices, White (2009) found that the distributions of the asset classes held by banks were non-normal.

Table 2. The proportion of US$-denominated sovereign Jamaican bonds held by the Jamaican deposit-taking institution

Bond	Sep-21	Dec-21	Mar-22
15-Jan-22	5.3%	5.5%	0.0%
9-Jul-25	13.1%	11.0%	11.7%
17-Oct-25	0.0%	1.0%	0.0%
28-Apr-28	30.8%	29.6%	31.2%
28-Feb-36	1.0%	1.6%	1.8%
15-Mar-39	18.8%	21.0%	22.2%
28-Jul-45	31.2%	30.3%	33.1%
Total	100.0%	100.0%	100.0%

Despite these shortcomings, the Gaussian assumption is attractive in its simplicity. Its invariance property makes it useful as a descriptor of risk factor behaviour when constructing Value-at-Risk models, allowing for analytic solutions for any holding period and parameters that can be easily estimated (Jorion, 1996; Vlaar, 2000). Furthermore, these classes of models can serve as a benchmark for the application of other Value-at-Risk techniques to understudied markets such as Jamaica's. Furthermore, Andersson (2021) shows that although individual bond returns may be non-Gaussian, when bonds are aggregated into bond portfolios, the normal distribution is adequate for modelling of portfolio returns for risk management purposes. So, the normal distribution continues to be used to model daily changes in bond yields in the context of building Value-at-Risk models (Pratiwi, 2024; Obadović et al., 2016).

The proportion in which the various bonds were held in Jamaica's banking system on the respective reporting dates is shown in Table 2. The composition of the banking system's aggregate bond portfolio did not change materially across these dates. The modest observed changes appeared to be related to the maturation of the Jan 2022 bond by March 2022.

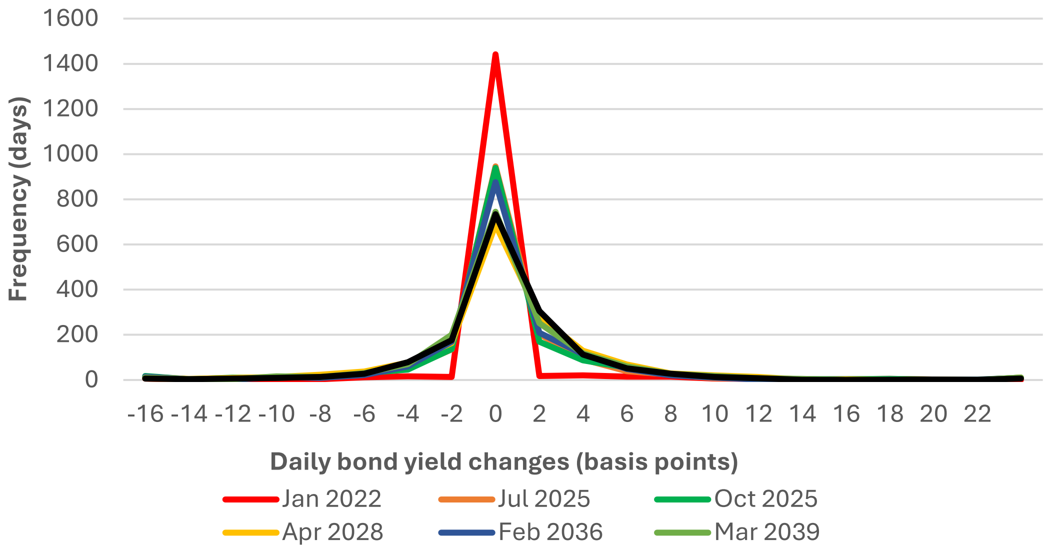


Figure 4. Frequency distribution of daily bond yield changes by bond

The system-wide VaR results at these reporting dates are shown in Table 3. In September 2021, with respect to the banking system's foreign currency-denominated Jamaican sovereign bond portfolio size of just under US$1 billion, the daily VaR was US$5.3 million, and the monthly VaR was US$24.4 million, both at the 95% confidence level. The corresponding VaRs at the 99% confidence level were US$7.5 million and US$34.4 million.

Table 3. Value at Risk (VaR) for the Jamaican banking system's sovereign bond portfolio (US$ '000)

Conf. Level	September 2021		December 2021		March 2022	
	Daily	Monthly	Daily	Monthly	Daily	Monthly
95%	(5,325)	(24,403)	(5,485)	(25,135)	(5,240)	(24,015)
99%	(7,511)	(34,418)	(7,735)	(35,447)	(7,391)	(33,868)

The system-wide CVaR results at the three reporting dates of September 2021, December 2021, and March 2022 are shown in Table 4. In September 2021, the banking system's foreign currency-denominated Jamaican sovereign bond portfolio was characterised by a CVaR of US$6.6 million and a monthly CVaR of US$30.2 million, both at the 95% confidence level. The corresponding CVaRs at the 99% confidence level were US$7.5 million and US$34.4 million.

Table 4. Conditional Value at Risk (CVaR) for the Jamaican banking system's sovereign bond portfolio (US$ '000)

Conf. Level	September 2021		December 2021		March 2022	
	Daily	Monthly	Daily	Monthly	Daily	Monthly
95%	(6,586)	(30,180)	(6,781)	(31,076)	(6,479)	(24,015)
99%	(7,511)	(34,418)	(7,735)	(35,447)	(7,391)	(33,868)

Figure 5 shows the observed monthly mark-to-market losses of the banking system from August 2021 to May 2025, along with a VaR estimate at the 95% confidence level, calibrated using data up to August 2021. This yields a Kupiec test statistic of $LR_{uc} = 1.176$ which supports the null hypothesis that the model is consistent with the observed outcomes. More generally, the VaR model passes all the backtesting procedures for consistency and independence, as shown in Table 5.

This is similar to the results in (Obadovic et al., 2016), where the parametric multivariate Gaussian VaR model applied to a Serbian government bond portfolio is shown to be effective at the 95% level using the Kupiec test. However, this contrasts with results in (Campbell & Smith, 2022), where VaR estimates produced by Australian banks are rejected by the standard backtesting procedures used here. Similarly, Omari (2017) finds that Gaussian VaR significantly underestimates risks on the Indian foreign exchange market. Also, in (Žiković & Aktan, 2009), all VaR models examined fail both the Kupiec and Independence tests when applied to Turkish and Croatian equity returns during the period of the Global Financial Crisis.

Table 5. Monthly VaR model backtesting results (95% confidence level)

Test	Test Statistic	Degrees of Freedom	Critical Value (5%)	p-value	Reject Null Hypothesis?
Unconditional Coverage	1.176	1	3.84	0.278	No
Independence	0.782	1	3.84	0.377	No
Conditional Coverage	1.957	2	5.99	0.376	No

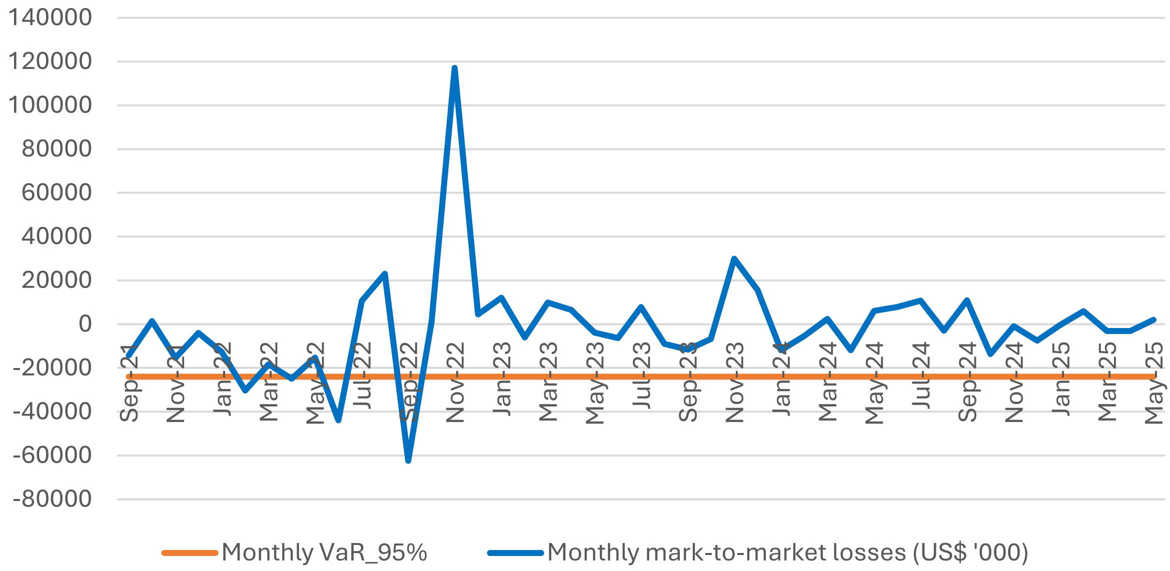


Figure 5. Monthly mark-to-market losses compared with Monthly $VaR_{95\%}$

The drawback of the above backtesting analysis is that it is applied to a financial dataset of only 45 monthly values covering a period of just under 4 years. By contrast, in (Obadovic et al., 2016), backtesting is applied to a bond portfolio with 1,262 daily values over five years, as well as to annual periods with approximately 250 daily values. To mitigate concerns of small-sample bias, as outlined in the Data and Calculations sub-section of this paper, we generate daily values of the aggregate banking system's foreign currency-denominated Jamaican sovereign bond portfolio over the period September 2021 to June 2025. This provides us with 1,000 daily portfolio values from which we calculate daily changes in the value of the portfolio.

Table 6 presents descriptive statistics on daily changes in portfolio value. Furthermore, Table 7 provides the frequency distribution of the daily changes in portfolio value. The distribution shows some asymmetry and is further characterised by fat tails with a notable concentration of daily portfolio changes near the centre. These observations are also evidence of non-normality. However, what both Table 6 and Table 7 show, interestingly, is that, while excess kurtosis rules out the normal distribution as a flawless theoretical descriptor, the empirical percentiles of the distribution of daily portfolio changes nevertheless mimic those of a standard normal distribution. Specifically, the 1st and 99th percentiles are approximately three standard deviations from the mean, while the 5th and 95th percentiles are approximately 1.6 standard deviations from the mean. Similarly, in (Li et al., 2025), descriptive statistics, inclusive of percentiles, are aggregated for thousands of emerging market bond returns, and it is shown that the 1st and 99th percentiles of these bond return distributions lie approximately three standard deviations from the mean bond return.

Table 6. Descriptive statistics of the daily portfolio changes

Statistic	US$ ('000)
Minimum	(23,268)
Maximum	16,174
Mean	(153)
Standard Deviation	3,336
1st percentile	(10,490)
5th percentile	(5,319)
95th percentile	4,652
99th percentile	9,534
Skew	-0.56
Kurtosis	6.43

Table 7. Frequency distribution of daily portfolio changes with corresponding Z Scores

Value Range (US$ '000)	Frequency	Cumulative Frequency	Midpoint	Z-Score
(10,000) and below	12	12	−10,000	−2.95
(9,000) to (10,000)	4	16	−9,500	−2.80
(8,000) to (9,000)	3	19	−8,500	−2.50
(7,000) to (8,000)	10	29	−7,500	−2.20
(6,000) to (7,000)	9	38	−6,500	−1.91
(5,000) to (6,000)	18	56	−5,500	−1.61
(4,000) to (5,000)	31	87	−4,500	−1.32
(3,000) to (4,000)	44	131	−3,500	−1.02
(2,000) to (3,000)	71	202	−2,500	−0.72
(1,000) to (2,000)	98	300	−1,500	−0.42
less than zero to (1,000)	263	563	0	0.05
0 to 1,000	140	703	500	0.20
1,000 to 2,000	112	815	1,500	0.50
2,000 to 3,000	76	891	2,500	0.80
3,000 to 4,000	41	932	3,500	1.10
4,000 to 5,000	24	956	4,500	1.39
5,000 to 6,000	15	971	5,500	1.69
6,000 to 7,000	9	980	6,500	1.98
7,000 to 8,000	5	985	7,500	2.28
8,000 to 9,000	3	988	8,500	2.58
9,000 to 10,000	3	991	9,500	2.87
10,000 and above	9	1,000	—	

Figure 6 shows the daily changes in the value of the aggregate banking system's Jamaican foreign currency-denominated sovereign bond portfolio over the period September 2021 to June 2025. It also shows the daily $VaR_{95\%}$ as calculated in Table 3. During the 1,000 days in this out-of-sample period, the $VaR_{95\%}$ threshold is breached 53 times. We apply the Kupiec Unconditional Coverage Test, yielding a test statistic of 0.18, which indicates that this exception rate is consistent with model expectations at the 95% confidence level.

Of these exceptions, there are six occurrences of breaches of the $VaR_{95\%}$ threshold on two consecutive days.[2] We apply Markov's Independence Test, which yields a test statistic of 3.15, indicating that these exceptions are not statistically clustered and, on the contrary, appear independent over time. Table 8 shows these back-testing results. These results mirror some of the findings of Watson & Rampersad (2011), who found that Kupiec Tests revealed that parametric VaR was most effective in assessing risk in the Barbadian, Eastern Caribbean, Jamaican, and Trinidadian equity markets. The results here also compare with (Fruzzetti et al., 2023), where Gaussian VaR passes backtesting procedures for consistency in an application of Gaussian VaR to Italy's foreign exchange reserves.

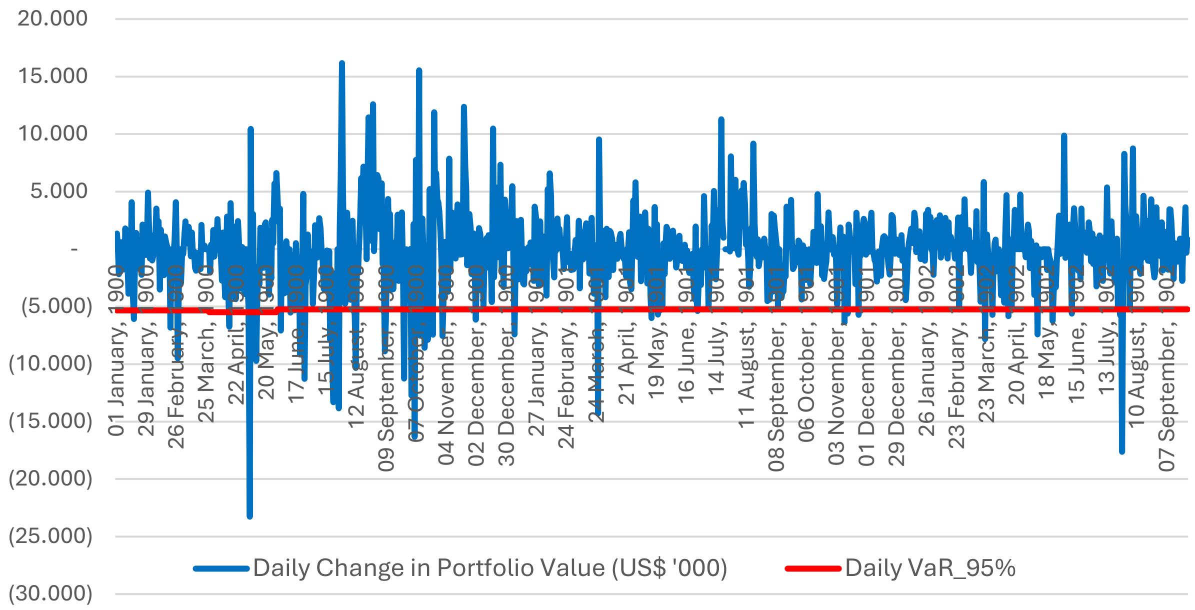


Figure 6. Daily portfolio changes compared with daily $VaR_{95\%}$

Table 8. Daily VaR model backtesting results (95% confidence level)

Test	Test Statistic	Degrees of Freedom	Critical Value (5%)	p-value	Reject Null Hypothesis?
Unconditional Coverage	0.180	1	3.84	0.6717	No
Independence	3.151	1	3.84	0.0759	No
Conditional Coverage	3.331	2	5.99	0.1892	No

Finally, in the case of comparing the banking system's monthly mark-to-market gains/losses on its foreign currency-denominated Jamaican sovereign bond portfolio, we observed that these losses exceeded the monthly $VaR_{95\%}$ threshold 4 times in this 45-month window. The average of these four losses that breached the monthly $VaR_{95\%}$ threshold is US$40.4 million, or approximately 4% of portfolio value, compared to a monthly CVaR estimate of US$30.2 million, or approximately 3% of portfolio value, at the 95% confidence level. Moreover, comparing the daily fluctuations in the value of the banking system's foreign-currency-denominated Jamaican sovereign bond portfolio, we observed that daily losses exceeded the $VaR_{95\%}$ threshold 53 times over the 1,000-day window. The average of these 53 losses that breached the daily $VaR_{95\%}$ threshold was US$8.5 million, or approximately 0.85% of portfolio value, compared to the daily CVaR estimate of US$6.6 million, or approximately 0.66% of portfolio value, at the 95% confidence level. Therefore, CVaR better captures extreme losses than VaR, both monthly and daily. CVaR also better describes extreme losses than VaR as conditional variance models of asset returns were used to generate VaR and CVaR estimates for the S&P 500, DAX, and Hang Seng stock indices (Du and Escanciano, 2016).

[2] These are difficult to see in Figure 6 as breaches, say seven days apart, appear close together in the figure.

Conclusion

This paper uses parametric VaR and CVaR models to measure the market risk from Jamaican banks' sovereign bond exposures before the surge in global interest rate volatility in 2022. The results, including backtesting, show that these standard risk measures effectively captured the Jamaican banking system's market risk exposure to Jamaican foreign currency-denominated sovereign bonds. Specifically, the calibrated parametric VaR and CVaR, based on data prior to the global interest rate increases, estimated much of the risk of mark-to-market losses that later impacted the Jamaican banking system's holdings of foreign currency-denominated Jamaican sovereign bonds. These findings also offer a helpful foundation for regulators, including those in Jamaica, to gradually adopt more current Basel frameworks and assist financial institutions in using these models to monitor market risk and determine their capital needs in response to these risks.

In the publication, the Basel Committee provides and describes two frameworks (Standardised Approach (SA) and Internal Models Approach (IMA)) for evaluating market risk and computing banks' capital requirements (Basel Committee on Banking Supervision, 2019). The capital requirement under Basel II's SA is the interest rate risk in the trading book (IRRTB), which is the sum of issuer-related risk (specific risk) and general market risk, applicable to all qualifying domestic and foreign currency-denominated positions. Under this approach, portfolio positions are grouped into maturity buckets involving a combination of specific risk weights depending on issuer quality, and general risk weights for yield curve risk are applied.

However, many central banks, including the Bank of Jamaica, still use this somewhat outdated Standardised Measurement Method (SMM) to determine capital requirements for banks' market risk. In this method, banks apply the regulator-defined risk weights to calculate their market risk-based capital needs. Although it is a more straightforward approach, it is less sensitive to risk when setting capital charges. The findings suggest that, on a systemic level, standard VaR and CVaR techniques can reasonably estimate the risk of Jamaican sovereign bonds. Therefore, these results offer a valuable example of how Jamaican regulators could enhance capital provisioning for the financial system.

Author contributions

Kishan Clarke: Methodology, Software, Validation, Formal analysis, Writing-original draft, Writing-review & editing.
Robert Stennett: Conceptualisation, Methodology, Validation, Writing-review & editing.

Use of AI tools declaration

The authors declare they have not used Artificial Intelligence (AI) tools in the creation of this article.

Conflict of interest

All authors declare no conflicts of interest in this paper.

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Non-cash food assistance and household food security: Evidence from remote Indonesia

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Abstract

Purpose — This study aims to evaluate the impact of Indonesia's Non-Cash Food Assistance Program (BPNT) on household consumption and food security in Sabu Raijua, a remote region in Indonesia with limited food access.

Methods — The Propensity Score Matching (PSM) with kernel techniques is employed to estimate the BPNT program's effects on household expenditure, caloric intake, and food insecurity using data from 536 households in Sabu Raijua, East Nusa Tenggara, Indonesia.

Findings — The results indicate that the BPNT program had a limited effect on household spending, nutrition, and food security. Beneficiary households spent slightly more on food and non-food items, showed minor improvements in nutritional intake, and were less likely to face food shortages, though concerns about food adequacy persisted.

Implications — The findings suggest that while BPNT helps alleviate food insecurity, further improvements in program implementation are needed to enhance its overall impact on household welfare.

Originality — This research provides critical insights into the effectiveness of BPNT in a remote region with unique socioeconomic challenges that have not been subject to empirical study, highlighting the challenges and opportunities for improving non-cash food assistance programs in similar contexts.

Keywords — BPNT, food expenditure, nutrient intake, food security

Introduction

Food insecurity remains a pressing challenge in many parts of the world, especially in regions that face environmental and socio-economic vulnerabilities (FAO, 2020). In Indonesia, a significant portion of the population struggles with poverty and malnutrition, making food assistance programs an essential component of the government's efforts to alleviate hunger and improve nutritional outcomes (Banerjee et al., 2023; Rammohan & Tohari, 2024). One such initiative is the *Bantuan Pangan Non-Tunai* (BPNT), or Non-Cash Food Assistance program, introduced by the Indonesian government in 2017. This program aims to enhance food security by providing low-income households with electronic vouchers to purchase essential food items from authorised vendors, replacing the previous *Raskin* (Rice for Poor Households) program, which distributed rice directly to beneficiaries (TNP2K, 2018).

The BPNT program is a critical component of Indonesia's social protection framework, designed to empower beneficiaries by offering flexible food choices and reducing logistical challenges associated with direct food distribution. The vouchers, worth IDR 200,000

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(approximately US$13) per month, can be redeemed for staple foods such as rice and eggs to meet the basic nutritional needs of low-income households (Government of Indonesia, 2017). This shift from in-kind distribution to electronic vouchers reflects a broader trend in social assistance programs that favours efficiency, recipient autonomy, and improved targeting to the most vulnerable populations (O'Farrell & Warokka, 2018; Timmer et al., 2017). These reforms mark a progressive step in Indonesia's fight against poverty and hunger.

While BPNT has been implemented nationwide, its effectiveness in geographically isolated regions, such as Sabu Raijua Regency in East Nusa Tenggara Province, remains uncertain. This region faces significant barriers such as poor infrastructure, high transportation costs, and limited market access, leading to food insecurity (SGP Indonesia, 2023). The National Food Agency classifies Sabu Raijua as vulnerable regarding food security, highlighting the need for targeted interventions (National Food Agency, 2021). While food assistance programs can mitigate these challenges, their success depends not only on food availability but also on vendor accessibility and distribution efficiency (Banerjee et al., 2019; Banerjee et al., 2024). Therefore, assessing the impact of BPNT in Sabu Raijua provides critical insights into the program's ability to meet its objectives in areas of food insecurity.

Food security encompasses the availability, access, and utilisation of food that meets the dietary needs and preferences of individuals (FAO, 2021). Achieving food security in regions like Sabu Raijua is particularly challenging due to high poverty levels, environmental stressors, and poor infrastructure. Even when food is available, households often struggle with the economic means to purchase it, leading to insufficient nutrition and adverse health outcomes (Skoufias et al., 2012). The BPNT program seeks to alleviate these issues by ensuring that vulnerable households have consistent access to essential food items, thereby improving overall food security and well-being. However, its impact on household food consumption and nutrition remains underexplored, particularly in remote regions where market constraints and logistical challenges may influence program effectiveness.

An ongoing debate in the literature concerns the relative effectiveness of cash versus in-kind transfers in improving food security and nutrition. The argument for providing food directly is often based on the idea that subsidising staple foods reduces food costs for vulnerable households, allowing them to spend their money on more nutritious, higher-value foods. However, evidence from programs in other countries suggests that in-kind transfers often fail to significantly improve nutrition outcomes, as they typically focus on basic grains that do not contribute to dietary diversity (Jensen & Miller, 2011; Kochar, 2005; Pingali et al., 2017; Rai et al., 2015). Conversely, some studies highlight that both food and cash have significantly different impacts on total food expenditures (Alderman et al., 2017; Banerjee et al., 2024; Cunha, 2014; Gentilini, 2016; Hidrobo et al., 2014). These findings suggest that the form of assistance provided can have significantly different impacts on food expenditure and nutritional outcomes, raising essential questions about the effectiveness of BPNT.

Although previous research has examined the general impacts of the BPNT program, there is a distinct lack of empirical evidence on its performance in challenging environments such as Sabu Raijua. This study evaluates the impact of the BPNT program on household food expenditure, caloric intake, and food security in Sabu Raijua Regency. Using Propensity Score Matching (PSM), it compares outcomes for BPNT recipients and non-recipients while controlling for household characteristics. Focusing on a region with unique socioeconomic challenges, this research aims to enhance the understanding of social protection programs in areas facing food insecurity. The findings will provide policymakers with valuable insights and contribute to broader discussions on using non-cash food assistance to combat food insecurity in resource-constrained settings.

Methods

Data Source

To assess the effectiveness of the BPNT program, this research analyses data from 536 households in Sabu Raijua, East Nusa Tenggara, Indonesia. The sample includes 265 households participating in the program (treatment group) and 271 households not participating (control group). Sabu Raijua, a geographically isolated archipelago in East Nusa Tenggara, faces severe food security challenges. Its remote location, about eight hours by sea southwest of Kupang City, combined with a drought-prone, semi-arid climate, contributes to its vulnerability (National Food Agency, 2021; SGP Indonesia, 2023). The region's food insecurity is underscored by a 33.9% stunting prevalence rate, as reported by the Indonesian Ministry of Health's Nutrition Status Study (SSGI) (Ministry of Health, 2021). These factors make Sabu Raijua an ideal case study for food security challenges.

Table 1. Research Variables

Variable	Definition
	Outcome Variable
Food expenditure	Monthly per capita expenditure on food
Non-food expenditure	Monthly per capita expenditure on non-food items
Total expenditure	Monthly per capita expenditure on food and non-food
Caloric intake	Daily caloric intake per capita
Protein intake	Daily protein intake per capita
Fat intake	Daily fat intake per capita
Carbohydrate intake	Daily carbohydrate intake per capita
Insufficiency concern	0: Worried about food sufficiency in the past year 1: Not worried
Food Intake Instability	0: Ate less due to food shortage in the past year 1: Did not eat less
Food ran out	0: Ran out of food in the past year 1: Did not run out of food
	Treatment variable
BPNT recipient	0: Non-recipient 1: Program recipient
	Control variables
Area residency	0: Urban 1: Rural
Household size	Number of household members
Sex of household head	0: Female 1: Male
Age of household head	Age of household head in years
Education level of household head	0: Not completed primary school 1: Elementary school 2: Middle school 3: High school 4: Diploma/bachelor/postgraduate
KKS ownership	0: No Kartu Keluarga Sejahtera (KKS) card 1: Has KKS card
Employment status	0: Unemployed 1: Employed
House ownership	0: Does not own a house 1: Owns a house
Roof type	0: Other materials 1: Roof tiles or zinc material
Wall type	0: Other materials 1: Concrete or brick
Floor type	0: Other materials 1: Marble, ceramics or cement
Water-sewage distance	Distance from water access to sewage

Source: BPS (2021)

The data for this analysis is drawn from the National Socio-Economic Survey conducted in March 2021 by Statistics Indonesia (BPS). The survey included detailed information on household demographics, socio-economic characteristics, and food security indicators. Key variables include household food and non-food expenditure, caloric and nutritional intake, and food insecurity indicators such as concern about food insufficiency, reduced food consumption, and household reports of running out of food. Covariates such as area of residence, household size, sex, age, education level, employment status, house ownership, and housing quality were also included to control for factors that might influence program participation and food security outcomes.

Analytical Strategy

Analysing at the household level enables a deeper understanding of how social assistance programs like BPNT affect food security, capturing the influence of various socio-economic factors on program uptake and outcomes among vulnerable populations. This household-level approach allows for a more detailed examination of the BPNT program's effectiveness in meeting beneficiaries' food security needs in a geographically isolated, economically constrained environment.

To estimate the causal effect of BPNT on the outcome variables, this study employs PSM with kernel matching. PSM is widely used in observational studies to reduce selection bias by creating a balanced sample of treated and control units based on observable characteristics (Rosenbaum & Rubin, 1983). Since BPNT participation is not randomly assigned, differences between recipients and non-recipients could introduce bias. PSM helps approximate a randomised experiment by ensuring comparability between groups, improving the validity of causal inferences (Caliendo & Kopeinig, 2008).

While PSM does not directly enforce the parallel trends assumption, it enhances its plausibility by minimising systematic differences between groups. By balancing observable characteristics, PSM reduces the likelihood that pre-existing disparities drive differences in outcomes. This mechanism strengthens the assumption that both groups would have followed similar trends in the absence of BPNT participation, increasing the credibility of the causal estimates.

Kernel matching was chosen over other PSM methods for its efficiency and ability to use all available control observations. Unlike nearest-neighbour or calliper matching, which may exclude relevant controls, kernel matching assigns higher weights to closer matches while assigning lower weights to others. This mechanism reduces variance and improves precision, making it particularly suitable for studies with larger samples, where a greater pool of control observations enhances estimation accuracy. Widely used in policy evaluations, kernel matching provides robust, balanced estimates of BPNT's impact on household food security and expenditure.

The matching procedure in this study comprises several stages. First, estimating the propensity score using a probit regression model. The propensity score is the predicted probability of receiving BPNT, where the regression takes the form:

$$Probability\ (bpnt = 1|X_i) = \beta_0 + \beta_i X_{ij} + \varepsilon_j \tag{1}$$

Where X_i represents the household-level covariate variables.

Second, matching observations using the kernel technique. This method utilises a weighted average of control households to match BPNT recipients with similar non-recipients, reducing variance by incorporating multiple comparisons rather than relying on the nearest neighbours. The kernel weight is calculated as:

$$\omega(i,i)_{KM} = \frac{K(\frac{P_j - P_i}{a_n})}{\sum_{k\in C}(\frac{P_j - P_i}{a_n})} \tag{2}$$

Here, Pi and Pj represent the propensity scores of the treated and control individuals, respectively, K(·) is the kernel function, and an is the bandwidth parameter.

Third, assessing covariate balance using paired t-tests. A well-balanced dataset ensures that differences in outcomes between BPNT recipients and non-recipients are attributable to the program rather than to underlying differences in household characteristics. The t-test equation is:

$$t = \frac{\bar{X}_T - \bar{X}_C}{\sqrt{\frac{{S_T}^2}{N_T} + \frac{{S_C}^2}{N_C}}} \tag{3}$$

Where:

- $\bar{X}_T$ and $\bar{X}_C$ are the means of the covariates for the treated and control groups, respectively.
- ${S_T}^2$ and ${S_C}^2$ are the variances of the covariates for the treated and control groups.

Fourth, calculating the Average Treatment Effect on the Treated (ATT) as a measurement of the intervention effect on the household outcome. It compares outcomes between BPNT recipients and their matched non-recipients, calculated as:

$$ATT = E\ (Y_1 - Y_0\ |bpnt = 1) \quad (4)$$

Where Y_1 is the observed outcome for BPNT recipients, and Y_0 is the counterfactual outcome for non-recipients.

Finally, validating the result with Rosenbaum Sensitivity Analysis. This tests how sensitive the estimated treatment effects are to potential unobserved confounders.

$$\Gamma = \frac{P(bpnt_i=1\ |X_i,U_i)}{P(bpnt_i=1\ |X_i)} \quad (5)$$

The sensitivity of the results at Γ=1 (no bias) and Γ=3 (moderate bias) evaluates whether the estimated ATT remains robust under different levels of potential hidden bias. If the ATT remains significant even with Γ=3, we can conclude that the results are robust to unobserved confounders.

Results and Discussion

This section presents the findings from the PSM analysis, examining the impact of the BPNT program on household expenditure, nutritional intake, and food security in Sabu Raijua Regency. The results highlight positive trends across all three dimensions, though many observed effects are not statistically significant. This section aims to provide a detailed yet accessible interpretation of these findings, offering insight into the potential benefits of the BPNT program while acknowledging its limitations.

The impact of the BPNT program on household expenditure is positive but modest, with slight increases observed in food, non-food, and total expenditure among BPNT recipients compared to non-recipients, as shown in Table 2. Food expenditure, representing monthly per capita food spending, increased by 3.7 percentage points. The average log-transformed food expenditure for BPNT recipients was 13.11, compared to 13.08 for non-recipients, with a t-statistic of 0.64, indicating that the increase, while positive, is not statistically significant. This result suggests that BPNT helps households allocate slightly more resources toward food, but the assistance provided may not be substantial enough to cause a significant shift in food spending.

Non-food expenditure, which includes spending on essential household items such as education, healthcare, and transportation, also increased of 8.5 percentage points. However, like food expenditure, this difference is not statistically significant. This result suggests that while BPNT might help households redirect some of their income toward non-food items, the assistance is too limited to change non-food expenditure patterns substantially.

Total expenditure, combining food and non-food, increased by 4.8 percentage points among BPNT recipients. Similarly, the increase is not statistically significant. This increase may suggest that BPNT helps stabilise household spending, thereby preventing consumption declines. Maintaining stable expenditure levels is a valuable outcome in regions like Sabu Raijua, where households are vulnerable to food insecurity and fluctuations in food availability.

The consistency provided by BPNT may help households maintain their standard of living, even if it does not lead to dramatic improvements in spending.

Table 2. Estimated Treatment Effects on Expenditure

Outcomes	Sample	Treated	Controls	Difference	S.E.	t-stat
ln (food expenditure)	Unmatched	13.111	13.297	-0.186	0.040	-4.630
	Matched	13.113	13.077	0.037	0.058	0.640
ln (non-food expenditure)	Unmatched	12.305	12.609	-0.304	0.071	-4.270
	Matched	12.309	12.224	0.086	0.102	0.840
ln (total expenditure)	Unmatched	13.507	13.743	-0.236	0.047	-5.000
	Matched	13.510	13.462	0.048	0.068	0.710

Source: Authors' calculation

Nutritional intake is a critical measure of the quality of household diets and directly influences health outcomes, particularly in regions where food security is a concern. The BPNT program, providing non-cash food assistance, aims to improve the nutritional status of recipient households by ensuring they have access to adequate food. This analysis examines the program's effects on caloric, protein, fat, and carbohydrate intake and reveals small positive changes, none of which are statistically significant.

Table 3. Estimated Treatment Effects on Nutritional Intake

Outcomes	Sample	Treated	Controls	Difference	S.E.	t-stat
ln(caloric)	Unmatched	7.633	7.744	-0.112	0.028	-4.020
	Matched	7.635	7.612	0.024	0.040	0.590
ln(protein)	Unmatched	4.042	4.183	-0.141	0.031	-4.510
	Matched	4.044	4.038	0.006	0.045	0.130
ln(fat)	Unmatched	3.823	3.985	-0.163	0.043	-3.780
	Matched	3.824	3.765	0.060	0.061	0.980
ln(carbohydrate)	Unmatched	5.782	5.851	-0.069	0.028	-2.500
	Matched	5.784	5.767	0.018	0.039	0.450

Source: Authors' calculation

The estimation results for nutritional intake are presented in Table 3. There was a modest 2.4 percentage-point increase in caloric intake among BPNT recipients, suggesting that the program helps households maintain or slightly increase their intake. However, this change was not large enough to be statistically significant, likely due to the limited value of assistance and other contextual factors affecting food access. Similarly, protein intake changed little, rising of only 0.6 percentage points. This suggests that while the BPNT program might help stabilise protein intake, it does not significantly increase access to protein-rich foods, which are often more expensive and less accessible in remote areas like Sabu Raijua. The limited impact on protein consumption may be due to households, despite receiving assistance, continuing to face barriers to purchasing higher-quality food items.

On the other hand, the fat intake showed a larger increase of 6.0 percentage points for BPNT recipients. Although this increase is not statistically significant, the BPNT program may contribute to a modest improvement in dietary quality, particularly in fat consumption. Fats are an important energy source and essential nutrients, and even a slight increase in fat intake could have positive implications for household nutrition. Lastly, carbohydrate intake increased of 1.8 percentage points, reflecting the program's minor effect on carbohydrate consumption. This is likely because staple foods like rice and cassava already form a substantial part of the diet in the region, and BPNT assistance may not significantly alter these existing consumption patterns.

Food security is a central objective of the BPNT program, as it aims to ensure that households have consistent access to sufficient and nutritious food. The analysis of food security indicators provides insight into how BPNT affects households' experiences of food scarcity and insecurity. The three indicators analysed—insufficiency concern, food intake instability, and food running out—offer a comprehensive view of food security among recipient households.

Table 4. Estimated Treatment Effects on Food Security

Outcome	Sample	Treated	Controls	Difference	S.E.	t-stat
Insufficiency concern	Unmatched	0.691	0.819	-0.129	0.037	-3.500
	Matched	0.695	0.749	-0.054	0.048	-1.130
Food intake instability	Unmatched	0.936	0.945	-0.009	0.021	-0.430
	Matched	0.935	0.901	0.034	0.027	1.250
Food ran out	Unmatched	0.981	0.974	0.007	0.013	0.540
	Matched	0.981	0.950	0.031	0.018	1.670

Source: Authors' calculation

The estimation results for food security are presented in Table 4. For the insufficiency concern indicator, which measures whether households were worried about having enough food in the past year, BPNT recipients reported slightly more concern than non-recipients. Among BPNT recipients, 69.5% reported not being concerned about food insufficiency, compared to 74.9% of non-recipients, resulting in a difference of -5.4 percentage points. Although this result is not statistically significant, it suggests that BPNT households may still face significant food security concerns despite receiving assistance. This condition could be due to broader economic factors in Sabu Raijua, such as seasonal food shortages, high prices, and limited market access, which may contribute to ongoing concerns about food sufficiency.

In contrast, the food intake instability indicator, which measures whether households had to eat less due to food shortages, showed a more favourable result. Among BPNT recipients, 93.5% reported not eating less due to food shortages, compared with 90.1% of non-recipients, a difference of 3.4 percentage points. While this result is not statistically significant, BPNT may help some households avoid reducing their food consumption during times of scarcity. This is an important finding, as reducing food intake can seriously affect household health and nutrition, particularly in vulnerable populations. Although modest, the result indicates that BPNT may play a protective role in maintaining food intake during economic or environmental stress periods.

The food ran out indicator, which measures whether households experienced running out of food in the past year, showed the most desirable outcome. Among BPNT recipients, 98.1% reported they were not running out of food, compared to 95.0% of non-recipients, resulting in a difference of 3.1 percentage points. This result approaches statistical significance at the 10% level, suggesting that BPNT may provide a crucial buffer against extreme food insecurity. Running out of food is one of the most severe forms of food insecurity, and the fact that BPNT recipients were less likely to experience this outcome indicates that the program is making a tangible difference in the lives of vulnerable households.

To ensure the reliability of the findings, this study conducted paired t-tests before and after matching, and the Rosenbaum bounds test to assess the sensitivity of our results to unobserved variables. These methods were chosen to mitigate potential biases in the analysis and strengthen the validity of the causal inferences regarding the BPNT program's impact.

Table 5. The Balancing Check

Variable	Sample	Mean		Bias	Bias reduction	t-test	
		Treated	Control	(%)	(%)	t	p>|t|
Area residency	Unmatched	0.030	0.114	-32.90	97.20	-3.80	0.000
	Matched	0.031	0.033	-0.90		-0.16	0.876
Household size	Unmatched	4.849	3.712	54.10	90.70	6.27	0.000
	Matched	4.786	4.680	5.10		0.57	0.567
Sex of household head	Unmatched	0.819	0.804	3.70	82.90	0.43	0.670
	Matched	0.817	0.819	-0.60		-0.07	0.942
Age of household head	Unmatched	50.687	47.554	20.20	52.10	2.34	0.020
	Matched	50.538	52.040	-9.70		-1.16	0.245
Education of the household head	Unmatched	0.834	1.384	-48.60	87.80	-5.61	0.000
	Matched	0.836	0.769	5.90		0.85	0.398
KKS ownership	Unmatched	0.525	0.140	89.20	99.10	10.34	0.000
	Matched	0.519	0.523	-0.80		-0.08	0.934
Employment status	Unmatched	0.940	0.886	19.20	96.60	2.22	0.027
	Matched	0.939	0.941	-0.70		-0.09	0.929
House ownership	Unmatched	0.947	0.900	17.70	63.80	2.04	0.042
	Matched	0.946	0.964	-6.40		-0.93	0.351
Roof type	Unmatched	0.528	0.598	-14.00	92.50	-1.62	0.105
	Matched	0.527	0.532	-1.00		-0.12	0.905
Wall type	Unmatched	0.426	0.469	-8.50	63.10	-0.98	0.327
	Matched	0.424	0.408	3.10		0.36	0.718
Floor type	Unmatched	0.460	0.502	-8.30	58.10	-0.96	0.338
	Matched	0.458	0.475	-3.50		-0.40	0.691
Water-sewage distance	Unmatched	0.902	0.804	27.70	59.10	3.21	0.001
	Matched	0.9008	0.8610	11.30		1.41	0.160

Source: Authors' calculation

First, this study performed paired t-tests to evaluate the balance of covariates between the treatment and control groups before and after matching, as shown in Table 5. Before matching, there were significant differences between the groups in key covariates, such as household size, employment status, and education level, that could bias estimates of the program's impact. However, after matching, the paired t-test results showed no statistically significant differences between the two groups across these observable characteristics. This result demonstrates that PSM successfully reduced bias introduced by differences in observed covariates, yielding a more comparable sample for analysis. By improving the balance between treated and untreated households, the method enhances the reliability of our estimates by focusing on households that are similar in observed characteristics, thus reducing selection bias.

Furthermore, this study applied the Rosenbaum bounds test to evaluate the sensitivity of our results to potential hidden biases. For key outcome variables such as food expenditure per capita, non-food expenditure per capita, total expenditure per capita, caloric intake, protein intake, fat intake, and carbohydrate intake, the Rosenbaum bounds analysis showed that unobserved confounders did not significantly affect the estimated treatment effects. Specifically, the bounds analysis indicated that our findings were robust to hidden biases within a reasonable range (Gamma = 1 to Gamma = 3), as there were no significant deviations in the results for most outcomes, including household food security measures such as insufficiency concern, instances of eating less due to food shortages, and food running out. These findings suggest that the observed treatment effects of the BPNT program are stable and not easily confounded by unobserved variables.

Table 6. The Rosenbaum Bounds Results

Gamma	sig+	sig-	t-hat+	t-hat-	CI+	CI-
ln(food expenditure)						
1	0	0	13.19	13.19	13.15	13.23
3	0	0	12.97	13.42	12.93	13.47
ln(non-food expenditure)						
1	0	0	12.48	12.48	12.41	12.56
3	0	0	12.05	12.87	11.97	12.94
ln(total expenditure)						
1	0	0	13.62	13.62	13.57	13.67
3	0	0	13.35	13.89	13.30	13.95
ln(caloric)						
1	0	0	7.68	7.68	7.65	7.71
3	0	0	7.52	7.85	7.49	7.88
ln(protein)						
1	0	0	4.10	4.10	4.07	4.13
3	0	0	3.93	4.28	3.89	4.32
ln(fat)						
1	0	0	3.90	3.90	3.85	3.94
3	0	0	3.66	4.14	3.61	4.20
ln(carbohydrate)						
1	0	0	5.82	5.82	5.79	5.84
3	0	0	5.66	5.97	5.63	6.01
Insufficiency concern						
1	0	0	1.00	1.00	1.00	1.00
3	0	0	0.50	1.00	0.50	1.00
Food intake instability						
1	0	0	1.00	1.00	1.00	1.00
3	0	0	1.00	1.00	1.00	1.00
Food ran out						
1	0	0	1.00	1.00	1.00	1.00
3	0	0	1.00	1.00	1.00	1.00

Source: Authors' calculation
*gamma: log odds of differential assignment due to unobserved factors
sig+ : upper bound significance level
sig- : lower bound significance level

t-hat+ : upper bound Hodges-Lehmann point estimate
t-hat- : lower bound Hodges-Lehmann point estimate
CI+ : upper bound confidence interval (a= 0.95)
CI- : lower bound confidence interval (a= 0.95)

The combination of PSM and the Rosenbaum bounds test provides a comprehensive approach to reducing bias and validating the results, as presented in Table 6. PSM addressed observable bias by matching similar households, thereby controlling for confounding variables that could distort the treatment effects. Meanwhile, the Rosenbaum bounds test extended the robustness check by examining the influence of unobserved bias. Together, these methods significantly reduce the risk of bias in the estimates, reinforcing the causal interpretation of the BPNT program's impact on household consumption and food security outcomes.

The BPNT program in Indonesia has demonstrated the potential to positively impact household outcomes, particularly in stabilising food expenditure, enhancing nutritional intake, and improving food security. These outcomes align with findings from conditional cash transfer programs in other countries, which have similarly contributed to poverty alleviation and improved household consumption patterns (Banerjee et al., 2023; Habimana et al., 2021; Nazareno & de Castro Galvao, 2023; Parker & Todd, 2017; Torres et al., 2025). The BNPT program provides non-cash food assistance to low-income households through electronic vouchers redeemable for essential food items. While this initiative has led to modest increases in both food and non-food expenditures and improvements in food security indicators, the observed effects in this study were not statistically significant. Several factors may explain both the program's positive contributions and the lack of significant results, with important implications for the future design and implementation of food assistance programs.

One key factor contributing to the BPNT program's positive effects on household welfare is its ability to provide direct financial relief. By offering food vouchers, the program effectively increases the purchasing power of low-income households, allowing them to access essential staples that might otherwise be unaffordable due to economic constraints (Egger et al., 2022; Filmer et al., 2023; Skoufias et al., 2013). The small positive changes in both food expenditure and total expenditure observed in this study suggest that the program enables households to maintain consistent consumption levels even during financial difficulties.

Moreover, the BPNT program plays a crucial role in stabilising food security outcomes, particularly in preventing households from running out of food. This aligns with findings from Hidrobo et al. (2014), Gadenne et al. (2024) and McIntosh and Zeitlin (2024), which highlight the effectiveness of food assistance programs in reducing extreme food insecurity. BPNT appears to serve as a safety net against the most severe forms of food deprivation, ensuring that vulnerable households do not experience shortages. This function is especially critical in remote regions such as Sabu Raijua, where geographic and market constraints limit access to food.

Despite these positive trends, the lack of statistical significance in most results can be attributed to several external factors, particularly food prices and market access, which directly influence the effectiveness of non-cash food assistance programs like BPNT. In remote regions like Sabu Raijua, where food supply chains are weak and transportation costs are high, food prices tend to be significantly elevated (Barrett et al., 2022; Nava et al., 2023). This reduces the purchasing power of BPNT vouchers, making it difficult for recipient households to obtain sufficient quantities and varieties of food items. Previous studies have shown that when market inefficiencies persist, well-intended aid interventions may fail to produce meaningful improvements in household welfare (Cunha et al., 2019; Hirvonen & Hoddinott, 2021). The price elasticity of demand for food in such areas means that a relatively small price increase can substantially erode the value of assistance, leaving households still vulnerable to food insecurity.

Limited market access further weakens the effectiveness of BPNT in Sabu Raijua. Many BPNT households rely on local vendors, but poor infrastructure, weak market integration, and supply shortages mean that even with vouchers, food choices remain limited (Wang et al., 2022). The success of food assistance programs depends not only on household purchasing power but also on the availability and affordability of essential food items (Hidrobo et al., 2014; McIntosh &

Zeitlin, 2024). When vendors face logistical difficulties replenishing stock, food supplies may be irregular, further reducing the intended benefits of BPNT. This issue is exacerbated when transport costs are passed on to consumers, effectively increasing the cost of basic food staples beyond what the voucher system can accommodate.

The findings of this study have significant implications for social policy in Indonesia and in other developing nations that implement non-cash food assistance programs. While the BPNT program shows potential to stabilise household food security, its modest, statistically insignificant impact suggests that further policy improvements are necessary to maximise its effectiveness. To achieve this, policymakers should consider adjusting the value of food vouchers, improving market access, and considering complementary interventions. Strengthening these aspects would enable BPNT to better support households in achieving sustainable food security and overall well-being.

Firstly, the limited impact of BPNT on food expenditure and nutritional intake suggests that the value of food vouchers may not be sufficient to generate meaningful improvements in dietary diversity and overall food security. To ensure beneficiaries can access a more nutritionally diverse food basket, policymakers should adjust voucher values to account for inflation and regional price variations. Aligning food assistance with local economic conditions has proven effective in other contexts, as seen in global best practices where targeted food programs have been adapted to specific economic realities. Studies have shown that food assistance programs are more effective when they account for local market fluctuations and purchasing power, ensuring that beneficiaries receive adequate support to meet their nutritional needs (Cunha et al., 2019; Hirvonen & Hoddinott, 2021).

Secondly, geographic disparities in BPNT's effectiveness further highlight the need for localised policy adaptations. In isolated areas, exemplified by Sabu Raijua, elevated food transport costs and restricted market access substantially undermine household food security. To address this issue, policymakers should explore complementary interventions, such as transportation subsidies for food vendors, investments in local food production, and improved distribution logistics. These measures would strengthen market access and enhance the efficiency of food assistance delivery. A relevant example is Ethiopia's Productive Safety Net Programme (PSNP), which successfully supplemented cash transfers with local infrastructure investments and packages of agricultural support to improve food accessibility in hard-to-reach areas (Gilligan et al., 2009; Mustafa et al., 2023). This demonstrates the potential of such integrated approaches to improve outcomes for vulnerable populations.

Furthermore, the findings indicate that while BPNT recipients experience slight improvements in food security, concerns about food insufficiency persist. This suggests that non-cash food assistance alone may not be sufficient to address broader issues of poverty and economic vulnerability in an isolated region. Therefore, achieving sustainable food security requires a wider policy approach. Food security policies must be embedded within comprehensive economic and social frameworks that prioritise adequate income for all and control the costs of essential goods and services (Penne & Goedemé, 2021). Integrating BPNT with other social protection measures, such as employment programs, microfinance initiatives, and nutrition education, could amplify its impact. This integrated approach would allow BPNT to transcend short-term food assistance and contribute to long-term poverty reduction and increased household resilience.

Incorporating these targeted improvements would enhance BPNT's ability to combat food insecurity and support long-term well-being for vulnerable populations. By refining its design and implementation, Indonesia can create a more effective and resilient social protection system that better serves the needs of its people. A more responsive and adaptive BPNT program would not only strengthen household food security but also provide a model for other developing nations seeking to optimise their food assistance programs.

Conclusion

This study assessed the impact of Indonesia's Non-Cash Food Assistance Program (BPNT) on household consumption and food security in Sabu Raijua, a remote region with significant economic and logistical challenges. The findings suggest that while the BPNT program has a positive effect on household food and non-food expenditure, nutritional intake, and food security, most of the observed changes are modest and not statistically significant. The program showed potential to stabilise household consumption, particularly by preventing food shortages, but its overall impact remains limited.

Beyond these empirical findings, this study contributes to both scholarly and policy discussions on non-cash food assistance and household food security. In the academic domain, this research adds empirical insights into non-cash food assistance in an emerging economy, particularly in contexts where both economic constraints and supply-side limitations drive food insecurity. By employing PSM to control selection bias, this study also contributes methodological insights for future evaluations of social assistance programs in developing economies. From a policy perspective, the results underscore the need to refine the design and implementation of programs like BPNT, especially in regions with limited market access and high food prices. Insights from this study can inform policy adjustments to improve effectiveness.

The findings of this study underscore the need for strategic policy enhancements to improve the effectiveness of the BPNT program in ensuring food security for vulnerable households. Given the modest and statistically insignificant impact observed, policymakers should consider adjusting food voucher values to reflect inflation and regional price variations, ensuring that beneficiaries can afford a nutritionally adequate diet. Additionally, addressing geographic disparities through targeted interventions, such as transportation subsidies, investments in local food production, and improved distribution networks, would enhance access to affordable food in remote areas. Beyond food assistance, integrating BPNT with broader social protection initiatives —including employment programs, microfinance, and nutrition education —could provide a more comprehensive approach to poverty reduction. These recommendations would strengthen the program's ability to support sustainable food security and overall well-being.

However, this study has some limitations. The cross-sectional nature of the data limits the ability to capture the long-term effects of the BPNT program, as the analysis provides only a snapshot of its impact at a specific point in time. Future studies could address this limitation by employing longitudinal data to examine the program's effects over time and conducting comparative analyses to account for geographic and economic variations. Additionally, the use of PSM in this analysis, while helpful in addressing selection bias, has limitations. PSM assumes that all relevant confounding variables are observed and accounted for, which may not be held in practice. Future studies could use alternative methods, such as Randomised Controlled Trials or Instrumental Variables, to address the issue of unobserved confounders.

In brief, the BPNT program provides a valuable framework for supporting food security in vulnerable regions. However, its current limitations indicate a need for strategic improvements. By addressing these challenges and enhancing program implementation, there is potential for significantly improving the lives of households in this underprivileged region. Continued research will be essential to evaluate its long-term effectiveness further and ensure that aid reaches those who need it most.

Author contributions

Rizki Tri Anggara: Conceptualisation, Writing–original draft, Formal analysis, Methodology, Writing–review.
Elsya Gumayanti Alfahma: Writing–review & editing, Formal analysis, Visualisation, Validation.

Use of AI tools declaration

The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest

The authors declare no conflicts of interest.

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Monetary policy and sectoral stock market in Malaysia

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Abstract

Purpose — This paper aims to examine the extent to which monetary policy shocks (domestic and international) will affect the movement of the sectoral stock index in Malaysia.

Methods — The monetary policy shocks are identified using a structural vector autoregressive (SVAR) model to examine the propagation of both monetary policy shocks (domestic and international) on sectoral stock prices.

Findings — The main results show that foreign monetary shocks significantly affect four sectoral stock indices: industrial and services, plantation, telecommunications, and utilities. In contrast, domestic monetary shocks impact three sectoral indices: industrial and services, technology, and utilities. However, domestic monetary policy shocks have a more dominant effect on the sectoral stock market in terms of magnitude.

Implication — The analysis results provide policymakers, particularly Bank Negara Malaysia (BNM), with valuable insights into which sectors are most sensitive to monetary policy fluctuations. Additionally, the results are beneficial for investors, as the analysis can help them manage their assets more effectively by identifying which sectoral stock indices are most affected by both domestic and international monetary policy shocks, and by guiding them to make more accurate investment decisions.

Originality — First, it focuses specifically on sectoral indices, examining all 13 in Malaysia through the lens of theory, with particular emphasis on impulse-response analysis, which explores the cumulative effects of both domestic and foreign monetary policy shocks on these indices. Secondly, the study employs a lagged analysis using the SVAR model, providing a theoretical framework for comparison with other relevant studies.

Keywords — Stock market, Malaysia, SVAR model, monetary policy shocks

Introduction

Stock markets are a fundamental component of the global financial system, serving as a barometer for economic health and providing investment opportunities. Monetary policy affects financial markets through several mechanisms, with changes in interest rates being one of the most direct and widely studied channels. For instance, Yusof and Majid (2007) and Kahler (2008) found that interest rates influence financial markets, contributing to stock market volatility. A tightening or loosening of monetary policy can alter investors' expectations of future economic conditions, thereby impacting asset prices, including stocks. In an increasingly globalised world, the effects of foreign monetary policy, particularly from major economies such as the United States, the

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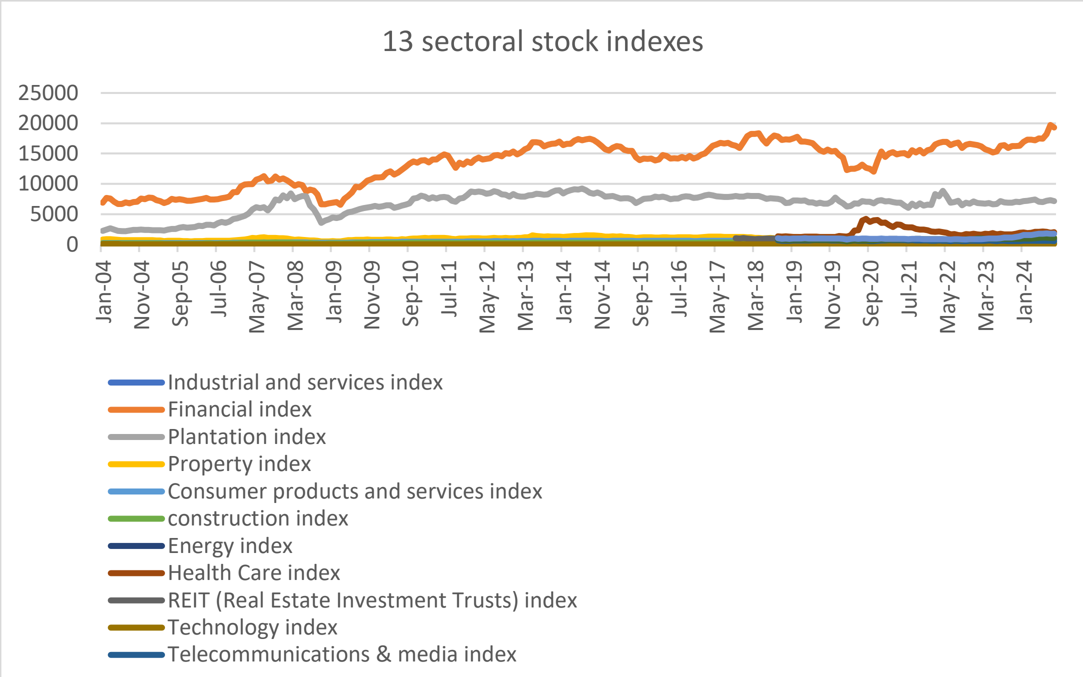
Eurozone, and China, can also spill over into domestic markets, influencing stock prices across sectors, as examined by Ha (2021) and Chiang (2021).

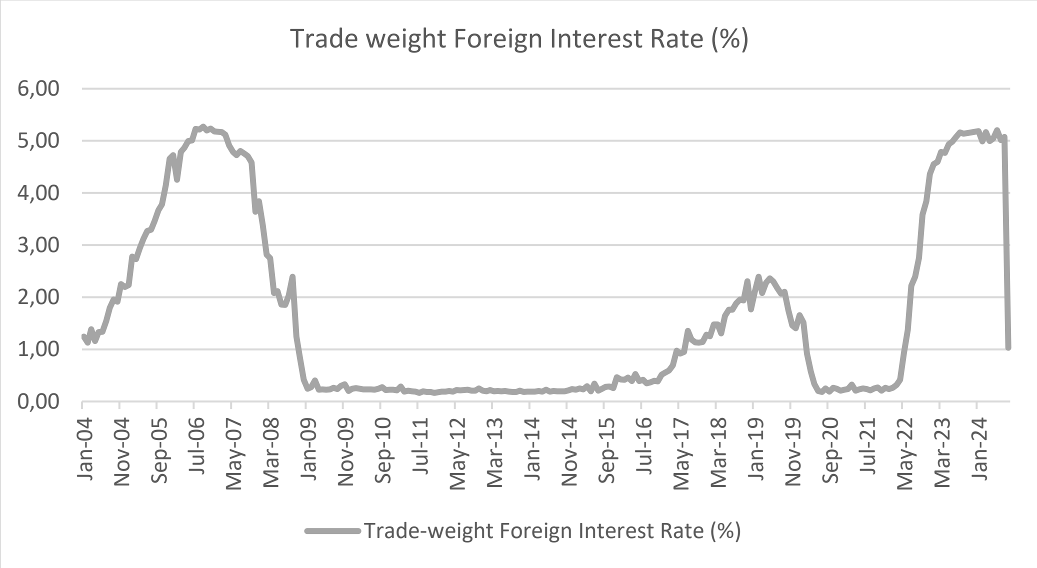


Source: Trading View, 2024.

Figure 1. Substitute Sectoral Index

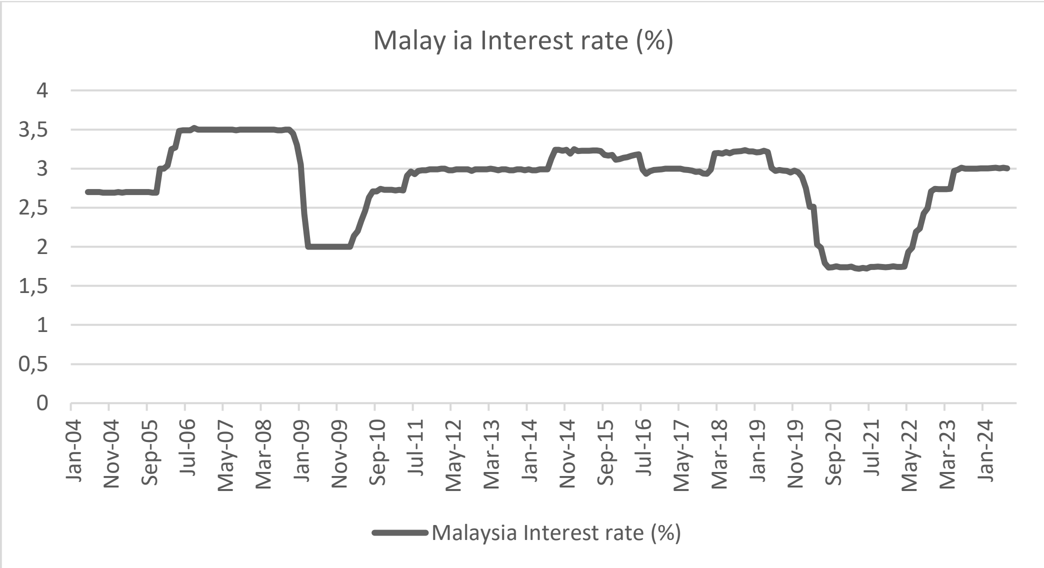
A key issue in financial research is how monetary policy (both domestic and foreign) affects different sectors of the economy. While there is a substantial body of work examining the relationship between overall monetary policy and stock market performance, the sectoral implications of monetary policy shocks have received relatively less attention, especially in Malaysia. Understanding the impacts of monetary policy on aggregate stock markets and sectoral stock indices differently. Most studies focus on aggregate market indices, overlooking the differentiated responses of specific sectors to monetary policy changes. For instance, Yusof and Majid (2007) and Bhatti, Ziaei, and Raheman (2015) examined stock returns in the KLCI composite market and analysed the effects of various monetary policy measures. Stock market sectors, such as banking, industrials, and technology, often exhibit distinct responses to changes in monetary policy because of their differing exposure to interest rates, liquidity, and global demand. For example, the plantation sector shows varied reactions to interest rate changes, with Kadir and Tunggal (2015) noting the effect but without specifying sensitivity levels. The financial industry, however, is expected to be more sensitive to monetary policy, as changes in interest rates directly affect banking operations and liquidity (Law and Ibrahim, 2014). Conversely, sectors like healthcare tend to be less affected, as demand for medical services remains relatively stable regardless of interest rate fluctuations, according to theory. As shown in Figure 1, the banking (financial) sector tends to respond similarly to changes in domestic interest rates, as illustrated in Figure 2, due to its dependence on lending activities and the spread between deposit and loan rates.

Foreign monetary policies, particularly those originating from major central banks like the US Federal Reserve (in combination with other countries as trade-weight according to Figure 2), have been the focus of past studies, such as Tchereni and Mpini (2020). As a small open economy, Malaysia is highly susceptible to changes in global financial conditions. When the Federal Reserve adjusts its monetary policy —whether through interest rate hikes, rate cuts, or quantitative easing —the effects are not confined to the US economy. It typically leads to higher asset returns in those countries, prompting investors to reallocate capital from emerging markets like Malaysia to seek better yields. This capital outflow results in the selling of Malaysian assets, including stocks, which in turn puts downward pressure on the stock market. As foreign investors sell Malaysian assets, the exchange rate between the Malaysian ringgit (MYR) and the US dollar (USD) may weaken, leading to a depreciation of the ringgit. A weaker ringgit increases the cost of imports and can negatively affect companies reliant on foreign goods and services, creating additional selling pressure on stocks.



Source: St. Louis Fred database, 2024.

Figure 2. Trade-weight foreign interest rate (%)



Source: Trading Economics, 2024.

Figure 3. Malaysia interest rate (%)

Furthermore, domestic monetary policy is primarily determined by Bank Negara Malaysia (BNM); some past studies found some impacts on the stock market by influencing economic conditions through tools like the Overnight Policy Rate (OPR) by Yakob, Tzeng, and Jr. (2014) and Hadi, Yap, and Zainudin (2019). Adjustments in the OPR directly influence short-term interest rates, borrowing costs, and credit availability, which in turn drive investor expectations and market movements. For example, when BNM raised the OPR in 2022 following the COVID-19 crisis, as shown in Figure 3, the increased borrowing costs reduced consumer and business spending, leading to declines in stock prices, particularly in interest-sensitive sectors. Conversely, a rate cut reduces financing costs, stimulates consumption and investment, and can drive stock prices higher, particularly in industries that rely on affordable credit, such as banking, real estate, and construction.

Therefore, BNM's monetary policy changes trigger sector-specific responses that directly influence the direction and volatility of stock market performance.

Given this background, this study contributes to two dimensions. First, policymakers, particularly central bankers, should examine several sectoral stock indices to determine which ones are most affected by monetary policy, specifically interest rate changes. For investors constructing more sector-sensitive investment portfolios, this study helps policymakers understand the broader impacts of their decisions on specific sectors. It enriches the academic literature on monetary policy transmission mechanisms in Malaysia. Secondly, by focusing on Malaysia, this research provides valuable insights into how external and internal monetary policy shocks influence sector-specific market movements in developing economies, advancing the work of Zaidi, Abdul Karim, and Azman-Saini (2017) and Karim and Karim (2016), it demonstrates that sectors more closely aligned with monetary policy tools, particularly those whose performance trends correlate with interest rate movements, exhibit stronger and more predictable responses to policy shocks. In contrast, sectors less dependent on monetary policy appear to be driven more by intrinsic or non-policy-related factors. In short, the novelty of this study lies in its application of a robust methodology to a relatively unexplored area of research, with far-reaching implications for theory, policy, and practice. Additionally, previous studies on Malaysia, such as those by Bhatti et al. (2015) and Law and Ibrahim (2014), have mainly focused on the impact of monetary policy on broad stock market indices or macroeconomic aggregates, such as GDP and inflation, rather than on sector-specific impacts. This study not only contributes by examining sector-specific implications of monetary policy but also opens several potential research avenues that could advance understanding of how monetary policy affects different sectors, thereby improving both theoretical and practical knowledge, building on the past study by Zuo (2025).

The remainder of the paper is organised as follows. Section 2 reviews the theoretical background of monetary policy and the previous literature on the effects of domestic and foreign economic policy on sectoral stock markets. Section 3 explained the data and methodology, including a small-open-economy SVAR model. Section 4 presents the main empirical results and discussion; Section 5 concludes.

Monetary Policy on the Stock Market

The impact of monetary policy on stock prices can be explained by a model that connects present value and a company's financial condition, as highlighted by researchers such as Karim and Zaidi (2015) and Sova and Lukianenko (2020), as shown below:

$$SP_t = E_t\left[\sum_{j=1}^{K}\left(\frac{1}{1+R}\right)^J D_{t+j}\right] + E_t\left[\left(\frac{1}{1+R}^K\right) SP_{t+k}\right] \tag{1}$$

Where SP_t is the financial stock price at time t, E_t is the conditional expectation operator value based on the market information at time t, D_{t+j} is the present value of expected future dividends, R is the rate of return, and K is the investor's time horizon for an asset holder. The transversality condition implies that as the horizon K increases, the second term in the right-hand side vanishes to zero (by predicting no rational stock price bubbles) or can be written as:

$$\lim_{K\to\infty} E_t\left[\left(\frac{1}{1+R}\right)^K SP_{t+k}\right] = 0 \tag{2}$$

Therefore, a similar equation for the company's financial situation (discounted cash flow) is:

$$SP_t = E_t\left[\sum_{j=1}^{\infty}\left(\frac{1}{1+R}\right)^j D_{t+j}\right] \tag{3}$$

In equation (3), the relationship between monetary policy and sectoral stock returns varies depending on the type of stock. Monetary policy can directly affect stock prices by influencing companies' finances, which in turn can affect their budgets and lead to either an increase or a decrease in stock prices. Additionally, monetary policy has an indirect effect on stock prices, which is shaped by a company's strategic planning and decisions. When the central bank adjusts monetary policy, companies assess how these changes affect their financial situation, which can influence stock prices, particularly in sectors like industrials and technology. This effect is often linked to changes in cash flow, as higher cash flow can lead to increased dividends in the future, thereby boosting stock prices and returns.

Transmission Mechanism of Monetary Policy Via the Asset Price Channel

Monetary policy has influenced the stock market, often leading to fluctuations in stock prices. This occurs through various channels, such as changes in interest rates, which affect the cost of borrowing and investment decisions. The transmission mechanism in this study concerns the effects of the stock market on investment, as described by Mishkin (2007). A higher stock price will bring high output; the following schematic can explain the transmission mechanism of monetary policy:

$$M \uparrow => SP_s \uparrow => Y \uparrow \tag{4}$$

Where M ↑ represents expansionary of monetary policy, bringing a rise in stock prices (SP_s ↑), which brings a rise in output (Y ↑).

Stock Market Response to Monetary Policy

Numerous studies have examined the impact of domestic monetary shocks on asset prices. For instance, in China, research has focused on sectoral asset prices, analysing several key sectors and their responses to various economic shocks and macroeconomic conditions, including monetary policy (Cai, Zhang, Han, and Liang, 2022). These studies suggest that different asset sectors respond to shocks in ways that depend on the specific economic environment. In Malaysia, Bhatti et al. (2015) examined the combined effects of monetary and fiscal policies on asset prices, finding that while international variables significantly influenced both policies, domestic variables had a lesser impact on asset prices. There is a stable relationship between monetary policy and sectoral output, with monetary policy exerting significant effects on both aggregate and sectoral outputs in terms of magnitude and direction (Ugwuanyi, Ezenekwe, & Kalu, 2021). In the United States, Paul (2020) discovered that house prices (property index) were less responsive to domestic monetary policy shocks when modelled using the SVAR approach, with this relationship varying over time. Similarly, an analysis by Suhaibu, Harvey, and Amidu (2017) indicated that domestic monetary policy influences stock markets through the interest rate channel in 12 African countries. In Nigeria, Nwakoby and Udoka (2016) found that domestic monetary policy shocks accounted for a significant portion of the impact on the stock market. In Indonesia, Handoyo, Jusoh, and Zaidi (2015) observed a positive short-term relationship between domestic monetary policy shocks announced by Bank Indonesia and the trend in the stock market. In Turkey, Baykara (2021) highlighted the critical role of domestic monetary policy on the stock market, particularly within the banking and financial sectors, where interest rates play a pivotal role in transmitting monetary policy effects to asset prices. Zuo (2025) demonstrates the direct impact of domestic monetary policy on the Malaysian stock market using a GARCH model. Moreover, Siang and Rayappan (2023) highlight that Malaysia's interest rate exerts a significant long-term influence on the performance of the domestic stock market.

The stock market is influenced not only by domestic variables but also by foreign variables. For example, US monetary policy has shown a significant and positive relationship with stock markets in Australia, the United Kingdom, and Canada, with specific periods of strong correlation. In contrast, New Zealand and South Korea exhibit an "S" shape, with both positive and negative relationships at different times (Ha, 2021). Studies have found that US monetary policy can affect global stock markets, although its influence is weaker on Latin American and Asian stock markets (Chiang, 2021). US monetary policy has a notably strong impact on European stock indices and similarly affects stock markets in Latin America and Asia (Kishor & Marfatia, 2013). The tightening of US monetary policy has been shown to boost the Irish stock market index (Bredin, Gavin, & O'Reilly, 2005). Additionally, UK monetary policy has been found to impact the German stock market, particularly the industrial sector (Bredin, Hyde, Nitzsche, & O'Reilly, 2009). US monetary policy also has a significant immediate effect on the Chinese financial market, surpassing its impact on European countries and Japan (Yang, Chen, & Mo, 2023). Monetary policy adjustments by the United States and Germany have been shown to influence the Moroccan stock market, driven by the economic policies of these countries (Belcaid & El Ghini, 2019). In Bangladesh, US and European monetary policies have a weaker but significant impact on major export countries, while major import countries like China and India experience a more substantial effect (Uddin, Hoque, & Ali, 2020). However, the announcement of US monetary policy had an unexpectedly negligible effect on the Indian stock market, despite India's large open economy and its economic ties with the US (Prabu A, Bhattacharyya, & Ray, 2016).

The Identification of Monetary Policy Using the SVAR Model

The monetary shock can affect all parts, including asset prices, which will also be studied in this paper. There have been many studies on the SVAR model to examine how the impact of monetary policy affects all parts of the economy, and vice versa. For example, in the United States, Wolf (2020) studied the real effect of monetary policy shocks on the US economy. The data the author found appear consistent with the theory and past studies, indicating that monetary policy shocks have a contemporaneous effect on the variable, and vice versa. Bacchiocchi, Castelnuovo, and Fanelli (2018) concluded that their analysis results using non-recursive SVAR were similar to those of the previous study, which employed Cholesky-type recursive SVAR in the United States via the impulse response function.

Several studies have focused on the SVAR model in Malaysia, including works by Karim and Karim (2016), Zaidi et al. (2017), and Zaidi and Fisher (2010), all of which provide detailed explanations of the SVAR approach. Prior to these studies, various research papers employed the VAR model (Ibrahim, 2005; Tang, 2006). However, the studies mentioned above demonstrated that the SVAR model is better suited for analysing monetary policy and macroeconomic variables in Malaysia, yielding more accurate results than those obtained with models used in other countries. For instance, Nizamani, Abdul Karim, Zaidi, and Khalid (2016) applied the SVAR model in Pakistan, Adrangi, Baade, and Raffiee (2019) in the United Kingdom, and Anwar and Nguyen (2018) in Vietnam, all of which also highlighted the superiority of the SVAR model in terms of imposing restrictions and aligning with theoretical expectations.

Monetary Policy and Stock Return from A Malaysian Perspective

IBOR, the Interbank Overnight Rate, has been utilised in Malaysia since 2004, following an announcement by Bank Negara Malaysia. Previously known by another name, it is the focus of this paper, which examines IBOR's usage since 2004. IBOR in Malaysia was divided into two categories, of which Malaysia was the Islamic country: conventional interbank overnight rate (CIBOR) and Islamic interbank overnight rate (IIBOR). CIBOR was the interest rate that did not comply with the rules of Shariah. At the same time, IIBOR was the interest rate that complied with the rules of shariah, which prohibited collecting interest improperly, such as casinos and beer profits, as well as receiving a reasonable return from the investor and lending to others via a proper channel (Ramasamy & Zangeneh, 2013).

As a result of the past study, the reflection of Islamic finance, including IIBOR, was to assist conventional finance in carrying out operations in the money market and to examine its influence on macroeconomic variables (Mat Sari, Mirakhor, & Mohd Subky, 2017). Although the Islamic finance system has the IIBOR and other related instruments related to Islamic monetary policy, the conventional rate, such as the CIBOR, also directly influences the Islamic financial system, according to the analysis (Othman & Masih, 2014). There is a high correlation between CIBOR and IIBOR. Based on the analysis, the Islamic banking system is expected to face the same risk as the conventional banking system, except for Bank Negara Malaysia's implementation of a special monetary policy tailored to the Islamic banking system (Bacha, 2008). The Islamic interest rate was found to be in long-run equilibrium with the conventional interest rate, according to the analysis. This is particularly evident in Middle Eastern countries such as Saudi Arabia, the UAE, and Bahrain, which are considered among the more influential Islamic nations, especially in terms of their economies (Nechi & Smaoui, 2019).

According to explanations from past studies on conventional bank systems, including CIBOR, and Islamic banking systems, including IIBOR, IIBOR appears to be a tool that functions well in Islamic finance systems. However, their roles remain less competitive than those of the CIBOR in conventional finance systems. Therefore, this research paper uses CIBOR as the variable for the Malaysian interest rate, also known as the domestic interest rate (DINR).

While some studies in Malaysia, such as Bhatti et al. (2015), indicate that short-term interest rates have little effect on stock prices. Yusof and Majid (2007) highlight a clear impact of interest rates on the stock market. Research by Kahler (2008) also concluded a negative relationship between interest rates and the stock market in Malaysia. Specifically, Law and Ibrahim (2014) found that domestic interest rates have a more pronounced effect on sectors such as the output and consumer sectors in Malaysia. Additionally, according to Yakob et al. (2014), when interest rates decline, the stock market tends to benefit.

Given this backdrop, this study fills the literature gaps in the following ways. First, it focuses specifically on sectoral indices, examining all 13 in Malaysia through the lens of theory, with particular emphasis on impulse responses. This approach analyses the accumulated effects of both domestic and foreign monetary policy shocks on these sectoral indices. Secondly, the study employs a lagged analysis using the SVAR model, which provides a theoretical framework for comparison with other relevant studies. This methodology provides a comprehensive understanding of how domestic and foreign monetary policies impact sectoral stock indices over time.

Methods

Data and Description of Variables

For this research, the methodology combines variables as studied by Karim and Karim (2016) and Zaidi et al. (2017). In the exogenous block, oil prices are included, as they can influence global demand and supply, a relationship previously explored by Karim and Karim (2016), Zaidi, Abdul Karim, & Zaidon (2021), and Ali, Zaman, and Islam (2018). Foreign GDP representing the top five trading partners of Malaysia is incorporated using a trade-weighted calculation, as done by Zaidi et al. (2017) and Zaidi and Fisher (2010). Similarly, foreign interest rates for Malaysia's top five trading partners are calculated using trade weights, following the same methodology. The top five trading partners, based on the most recent data, are Singapore, Mainland China, the United States, Hong Kong, and Japan. Using data from these five countries is more appropriate for the SVAR model, as it allows for more comprehensive and moderate predictions. Relying on data from just one country would limit the model's ability to produce robust, accurate results. The second block consisted of Malaysia's GDP, Malaysia's inflation rate (%), Malaysia's interest rate (the conventional interbank overnight rate, CIBOR), the exchange rate (ER), and Malaysia's sectoral stock index, with 13 types of indices substituted as well.

The study utilised monthly data from January 2004 to September 2024. The domestic interest rate was converted from daily to monthly using EViews software, with data available from May 2004 to July 2024. Malaysian GDP and foreign GDP were also converted from annual to monthly data using EViews. The Malaysian GDP data spanned from January 2004 to January 2023 due to technical issues with frequency conversion. The sectoral stock index data, including the energy, healthcare, telecommunications, media, transportation, logistics, and utilities indices, was available from September 2018 to September 2024. The REIT index data covered October 2017 to September 2024, while the other data followed the exact timeframes as previously mentioned. Oil price data were sourced from the Macrotrends website; foreign output, foreign interest rates, and domestic output were obtained from the St. Louis Fed database; and domestic inflation and domestic interest rate data were sourced from Trading Economics. Exchange rate and Malaysia's sectoral stock index data were sourced from Trading View. The study focused on eight key variables: oil price (OP), foreign output (FO) representing the GDP of Malaysia's top five trading partners (in billion dollars), foreign interest rate (FIR) as trade weights of the top 5 countries, domestic output (DO) as Malaysia's GDP (in billion dollars), domestic inflation rate (DIF) representing Malaysia's inflation rate as a percentage, domestic interest rate (DINR) as the average conventional interbank overnight rate (CIBOR), exchange rate

(ER) between the USD and Malaysian Ringgit, and the Industrial & Services Index (ISI), which is the sectoral stock prices and substitutes with various sector indices in one model (financial, plantation, property, consumer & services, construction, technology, energy, healthcare, real estate investment trust, telecommunications and media, transportation and logistics, and utilities indices).

There will be three dummies, which are the financial crisis of 2008-2009 (Ali & Hatta, 2013; Bahaludin, Abdullah, Lam, & Lam, 2019; Lee, 2011), the COVID-19 pandemic from 2020 to 2023, June Hassan (2024), and currency pegging between the USD and the Malaysian ringgit from January 2004 until May 2007 (Pourkalbassi, Bahiraie, Hamzah, & Lee, 2011; Sidek & Yusoff, 2009). This research will consider exogenous variables to standardise and justify the impact of monetary shocks on sectoral stock indices, assigning values of 1 or 0 only during the crisis period, using the SVAR model. The long form of the variables will be the table below:

Table 1. Variables description

Variable	Notation	Description
Oil Price	OP	Oil Price in US dollars (USD) per barrel based on Karim and Karim (2016)
Foreign Output	FO	Malaysia's Top 5 trading partners trade-weight GDP (USD) based on Zaidi and Fisher (2010)
Foreign Interest Rate	FIR	Malaysia's Top 5 trading partners' trade-weight interest rate (%) based on Zaidi and Fisher (2010)
Domestic Output	DO	Malaysia GDP (million dollars)
Domestic Inflation Rate	DIF	Malaysia Interest Rate (%)
Domestic Interest Rate	DINR	Conventional interbank overnight rate (CIBOR) (Average)
Exchange rate	EX	Exchange rate between USD and Ringgit Malaysia
Industrial & Services index	ISI	Malaysia's industrial and services stock market performance
Financial index	FI	Malaysia's financial stock market performance
Plantation index	PLTI	Malaysia plantation stock market performance
Property index	PRPI	Malaysia's property stock market performance
Consumer & Services index	CSI	Malaysia's consumer and services stock market performance
Construction index	CSTI	Malaysia construction stock market performance
Technology index	TECHI	Malaysia's technology stock market performance
Energy index	ENI	Malaysia's energy stock market performance
Health Care Index	HCI	Malaysia's health care stock market performance
Real Estate Interest Trust (REIT) index	REITI	Malaysia real estate interest trust stock market performance
Telecommunication and Media Index	TMI	Malaysia's telecommunications and media stock market performance
Transportation and Logistics Index	TLI	Malaysia's transportation and logistics stock market performance
Utilities index	UTI	Malaysia's utilities stock market performance

Source: Author's work

All data were logarithmic except for foreign interest rate, domestic inflation rate, and domestic interest rate. The model below shows that there will be a study of about 1 model, which is substituted on the ISI with other sectoral stock indices as follows:

Table 2. Model variables

Variables	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1	Model 1
1	OP	OP	OP	OP	OP	OP	OP	OP	OP	OP	OP	OP	OP
2	FO	FO	FO	FO	FO	FO	FO	FO	FO	FO	FO	FO	FO
3	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR	FIR
4	DY	DY	DY	DY	DY	DY	DY	DY	DY	DY	DY	DY	DY
5	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF	DIF
6	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR	DINR
7	ER	ER	ER	ER	ER	ER	ER	ER	ER	ER	ER	ER	ER
8	ISI	FI	PLTI	PRPI	CSI	CSTI	TECHI	ENI	HCI	REITI	TMI	TLI	UTI

Source: Author's work

Structure VAR Modelling

According to Karim and Karim (2016) and Zaidi (2011), the structural VAR equation is:

$$A_0\, Y_t = C + (\Gamma_1 L + \Gamma_2 L^2 + \ldots + \Gamma_k L^k) Y_t + \varepsilon_t \tag{5}$$

Where A_0 is an inverse square matrix of the contemporaneous structural parameters, Y_t is a (8 x 1) matrix of economic variables or $[\Delta LOP \quad \Delta LFO \quad \Delta FFR \quad \Delta LDO \quad \Delta INF \quad \Delta INT \quad \Delta ER \quad \Delta ISI]'$, C is a vector of deterministic variables, Γ (L) is the k-th order of the matrix in the lag operator, L, and ε_t is the structural shocks which fulfil the conditions that E(ε_t)=0, E($\varepsilon_t \varepsilon_s'$)= I for all the t and s and E($\varepsilon_t \varepsilon_s'$)= otherwise. (/) means there are different asset price indices in the system based on different models.

Pre-multiplying equation (1) with A_0^{-1} acts as the reduced form of the VAR equation according to Zaidi (2011):

$$Y_t = A_0^{-1} C + A_0^{-1} (\Gamma_1 L + \Gamma_2 L^2 + \ldots + \Gamma_k L^k)\, Y_t + A_0^{-1} \varepsilon_t \tag{6}$$

Where the $e_t = A_0^{-1} \varepsilon_t$ is the reduced form from VAR, which satisfies the conditions that E($e_t) = 0$, E($e_t e_s$')=$\sum_e$. $\sum_e$ is the (nxn) symmetric, positive-definite matrix, which can also be estimated. Alternatively, they have another way to show the SVAR model in function form as below:

$$Y_t = f(Y_{1,t}, Y_{2,t}) \tag{7}$$

$$Y_{1,t} = f(OP_t, FO_t\, , FIR_t) \tag{8}$$

$$Y_{2,t} = f(DO_t, DIF_t, DIR_t, ER_t, ISI_t) \tag{9}$$

$Y_{1,t}$ represents the foreign block, and $Y_{2,t}$ represents the domestic block. Additionally, ISI_t is replaced by sectoral index variables in the SVAR model.

Identification Scheme

According to the SVAR order condition, the system is suitable for both identified and over-identified cases, requiring K(K-1)/2 = 8 (7)/2 = 28 zero restrictions on the contemporaneous matrix A_0. Since the contemporaneous matrix A_0 has 35 zero restrictions; the model was over-identified. In matrix form, it is shown in Equation (10).

Since Malaysia was a small-open country, the foreign variables did not respond contemporaneously or with a lag to the domestic variables as well (Karim & Karim, 2016). More specifically, the oil price was the structural disturbance. A straightforward way to understand this is that oil prices act as demand and supply shocks on other variables or exogenous variables. Malaysia's top 5 trading partners, including the United States, can affect oil prices. The United States was the world's largest economy, and oil was its most significant import.

$$\begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 & 0 & 0 & 0 \\ a_{41} & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ a_{51} & 0 & 0 & a_{54} & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & a_{64} & a_{65} & 1 & 0 & 0 \\ a_{71} & a_{72} & a_{73} & a_{74} & a_{75} & a_{76} & 1 & 0 \\ a_{81} & a_{82} & a_{83} & a_{84} & a_{85} & a_{86} & a_{87} & 1 \end{bmatrix} \begin{bmatrix} \mu_t^{\Delta LOP} \\ \mu_t^{\Delta LFO} \\ \mu_t^{\Delta FIR} \\ \mu_t^{\Delta LDO} \\ \mu_t^{\Delta INF} \\ \mu_t^{\Delta INT} \\ \mu_t^{\Delta LER} \\ \mu_t^{\Delta LISI} \end{bmatrix} = \begin{bmatrix} \varepsilon_t^{\Delta LOP} \\ \varepsilon_t^{\Delta LFO} \\ \varepsilon_t^{\Delta FIR} \\ \varepsilon_t^{\Delta LDO} \\ \varepsilon_t^{\Delta INF} \\ \varepsilon_t^{\Delta INT} \\ \varepsilon_t^{\Delta LER} \\ \varepsilon_t^{\Delta LISI} \end{bmatrix} \tag{10}$$

The foreign output does not have an immediate effect on domestic output, domestic inflation rate, and domestic interest rate, as domestic output and domestic inflation are slow-moving; regarding the monthly data, domestic output is a slow-moving variable, while, when measured annually or quarterly, it will be affected by foreign output and foreign interest rate. Foreign interest rates are also affected by dynamics like those affecting foreign output, as mentioned above. For exchange rate and industrial & services indices, substitute other sectoral stock indices that react contemporaneously to the foreign output shock and foreign interest rate shock, indicating that policymakers must ensure low and stable inflation, adjusting immediately when necessary (Zaidi et al., 2017). Regarding oil prices, no effect was observed on domestic interest rates, as Bank Negara Malaysia (BNM) does not base its interest rate decisions on oil price fluctuations, as confirmed in BNM's 2024 announcement. The exchange rate affects the interest rate contemporaneously, with studies providing evidence to resolve the "exchange rate puzzle" (Zaidi et al., 2017). The exchange rate also impacts the ISI index and other sectoral stock indices, reflecting their fast-moving nature. Lastly, the Industrial & Services Index (ISI) and other substitute indices are assumed to influence both foreign and domestic variables, as these indices are fast-moving variables in the system.

As the foreign block is assumed to be the block exogenous, the zero restrictions on the lag values of the domestic variables are represented in equation (11) of the lag coefficient. The vector C contains the intercept and dummy variables. The oil price, foreign output, and foreign interest rate equations are only functions of the lag of oil price, lag of foreign output, and lag of foreign interest rate, and all the domestic variables prefer to lag foreign and domestic variables as well, based on the study from Zaidi et al. (2017).

$$\begin{bmatrix} \Delta LOP_t \\ \Delta LFO_t \\ \Delta FIR_t \\ \Delta LDO_t \\ \Delta INF_t \\ \Delta INT_t \\ \Delta LER_t \\ \Delta LISI_t \end{bmatrix} = \begin{bmatrix} \delta_{11} & \delta_{12} & \delta_{13} \\ \delta_{21} & \delta_{22} & \delta_{23} \\ \delta_{31} & \delta_{32} & \delta_{33} \\ \delta_{41} & \delta_{42} & \delta_{43} \\ \delta_{51} & \delta_{52} & \delta_{53} \\ \delta_{61} & \delta_{62} & \delta_{63} \\ \delta_{71} & \delta_{72} & \delta_{73} \\ \delta_{81} & \delta_{82} & \delta_{83} \end{bmatrix} \begin{bmatrix} dfc \\ dcv \\ dpg \end{bmatrix} + \begin{bmatrix} a_{11} & a_{12} & a_{13} & 0 & 0 & 0 & 0 & 0 \\ a_{21} & a_{22} & a_{23} & 0 & 0 & 0 & 0 & 0 \\ a_{31} & a_{32} & a_{33} & 0 & 0 & 0 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & a_{44} & a_{45} & a_{46} & a_{47} & a_{48} \\ a_{51} & a_{52} & a_{53} & a_{54} & a_{55} & a_{56} & a_{57} & a_{58} \\ a_{61} & a_{62} & a_{63} & a_{64} & a_{65} & a_{66} & a_{67} & a_{68} \\ a_{71} & a_{72} & a_{73} & a_{74} & a_{75} & a_{76} & a_{77} & a_{78} \\ a_{81} & a_{82} & a_{83} & a_{84} & a_{85} & a_{86} & a_{87} & a_{88} \end{bmatrix} \begin{bmatrix} \Delta LOP_{t-i} \\ \Delta LFO_{t-i} \\ \Delta FIR_{t-i} \\ \Delta LDO_{t-i} \\ \Delta INF_{t-i} \\ \Delta INT_{t-i} \\ \Delta LER_{t-i} \\ \Delta LISI_{t-i} \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \\ e_{3t} \\ e_{4t} \\ e_{5t} \\ e_{6t} \\ e_{7t} \\ e_{8t} \end{bmatrix} \quad (11)$$

Equation (11) indicates that domestic variables do not affect all the foreign variables as initially assumed. Specifically, Malaysia, as a small open economy, lacks the power to influence the global economy. Although the lag identification was approximately 2, and the econometric tools provided analysis, it appears this study was based on monthly data. Using 8–12 lags in the analysis can help ensure an accurate estimate of the monetary policy shock for the sectoral index in Malaysia. Still, it seems this study used six sectoral stock indices that lacked data, so it is considered to have two lags.

Results and Discussion

The optimal lag for the model was determined using Akaike's Information Criterion (AIC) and Schwarz Criterion (SC). Table 3 shows that SC selected one lag, while AIC chose two lags for model 1 based on the sectoral stock indices. The study employs two lags for further analysis, as this approach is deemed more suitable and accurate according to the AIC criterion. The stability test showed that all eigenvalues were less than 1, confirming the stability of the SVAR model.

Table 3. Results of the lag length test on Model 1

Number of Lags	AIC	SC
0	-27.341	-26.841
1	-31.254	-29.754*
2	-31.396*	-28.896
3	-31.317	-27.817
4	-31.203	-26.703
5	-31.156	-25.655

*Represents optimal lag length
Source: Author's work

Stationarity Test

Table 4 presents the results of the unit root test for the variables. According to the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, all variables were stationary at the first difference, or I(1), except for DIF ($p<0.01$), OP and ENI ($p<0.05$), and PLTI, DINR, and TMI ($p<0.10$), which were stationary at the level form, or I(0). The Foreign Output (FO) variable appeared to be I(2), but for model selection, it was decided to substitute between the ISI index and the sectoral index. These stationarity tests follow methodologies used by Zaidi et al. (2021), Ali et al. (2018), and Karim and Karim (2016).

Furthermore, the unit root test (KPSS) indicates that when the test is significant, the data are non-stationary, and when it is not, the data are stationary. Based on the results as shown in Table 5, most of the variables are stationary, except for OP (oil price), FO (foreign output), DO (domestic output), CSI (consumer and services index), TECHI (technology index), and UTI (utilities index).

Table 4. Result of the unit root tests

Augmented Dickey-Fuller				Phillips-Perron			
Level-Form		1st Difference		Level-Form		1st Difference	
OP	-3.164**	DOP	-13.805***	OP	-2.904**	DOP	-13.834***
FO	1.323	DFO	-2.309	FO	2.269	DFO	-2.382
FIR	-2.126	DFIR	-4.332***	FIR	-1.909	DFIR	-10.851***
DO	-2.298	DDO	-3.164**	DO	-3.146**	DDO	-3.348**
DIF	-4.861***	DDIF	-8.069***	DIF	-3.689***	DDIF	-10.103***
DINR	-2.682*	DDINR	-6.437***	DINR	-2.274	DDINR	-10.219***
EX	-1.155	DEX	-14.865***	EX	-1.212	DEX	-14.864***
ISI	-1.333	DISI	-13.712***	ISI	-1.431	DISI	-13.709***
FI	-1.597	DFI	-14.336***	FI	-1.678	DFI	-14.398***
PLTI	-2.808*	DPLTI	-14.377***	PLTI	-2.766*	DPLTI	-14.456***
PRPI	-1.573	DPRPI	-14.072***	PRPI	-1.954	DPRPI	-14.292***
CSI	-2.012	DCSI	-16.303***	CSI	-2.027	DCSI	-16.303***
CSTI	-1.831	DCSI	-14.632***	CSTI	-2.199	DCSI	-14.734***
TECHI	-1.057	DTECH	-13.712***	TECHI	-1.246	DTECH	-13.737***
ENI	-3.063**	DENI	-8.631***	ENI	-3.089**	DENI	-8.755***
HCI	-1.446	DHCI	-7.494***	HCI	-1.763	DHCI	-7.729***
REITI	-2.054	DREITI	-9.941***	REITI	-1.959	DREITI	-10.322***
TMI	-2.893*	DTMI	-7.822***	TMI	-2.968**	DTMI	-9.587***
TLI	-1.304	DTLI	-8.819***	TLI	-1.166	DTLI	-9.393***
UTI	-1.060	DUTI	-8.428***	UTI	-1.117	DUTI	-8.425***

Notes: *, ** and *** represent significance at 10%,5% and 1% levels, respectively.
Source: Author's work

Table 5. Result of the unit root tests (KPSS test)

Kwiatkowski–Phillips–Schmidt–Shin test (KPSS test)			
Level-Form		1st Difference	
OP	0.657**	DOP	0.103***
FO	1.892***	DFO	0.504**
FIR	0.299	DFIR	0.075
DO	1.726***	DDO	0.532**
DIF	0.186	DDIF	0.023
DINR	0.320	DDINR	0.057
EX	1.171***	DEX	0.152
ISI	1.810***	DISI	0.035
FI	1.503***	DFI	0.085
PLTI	0.995***	DPLTI	0.316
PRPI	0.373*	DPRPI	0.067
CSI	1.633***	DCSI	0.451*
CSTI	0.316	DCSI	0.068
TECHI	0.971***	DTECH	0.456*
ENI	0.327	DENI	0.076
HCI	0.207	DHCI	0.117
REITI	1.026***	DREITI	0.243
TMI	0.536**	DTMI	0.047
TLI	0.857***	DTLI	0.230
UTI	0.472**	DUTI	0.715**

Notes: *, ** and *** represent significance at 10%,5% and 1% levels, respectively.
Source: Author's work

To ensure consistency in the unit root testing of the variables, the HEGY test is used, as it is appropriate for monthly data and helps determine whether differencing is required. Based on the results shown in Table 6, most variables need to be differenced, except for FIR (Foreign Interest Rate), CSI (Consumer and Services Index), and TLI (Telecommunication Index). Therefore, differencing is necessary for the remaining variables. Based on the four types of differencing unit root tests, all variables needed to be differenced, as this stabilised the data and made it easier to observe and interpret.

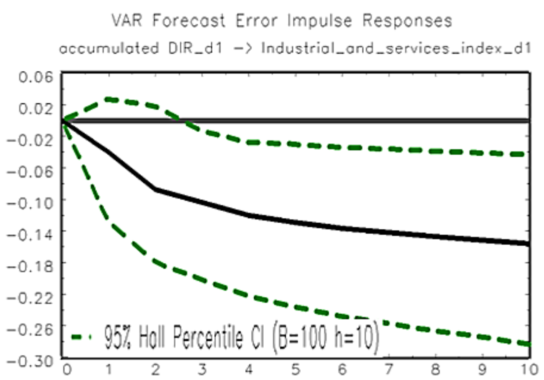
Table 6. Result of the seasonal unit root tests (HEGY test)

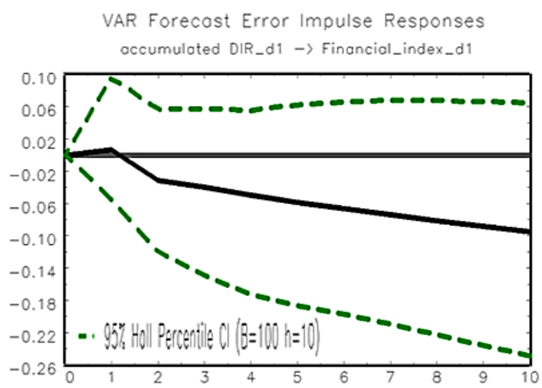
HEGY Test	
	π
OP	-0.133
FO	-1.609
FIR	-2.868***
DO	-1.495
DIF	-1.228
DINR	-0.370
EX	-0.090
ISI	-1.098
FI	-1.075
PLTI	-1.002
PRPI	-0.225
CSI	-1.632*
CSTI	-0.270
TECHI	-0.034
ENI	-0.237
HCI	-1.005
REITI	-0.426
TMI	-1.285
TLI	-3.494***
UTI	-0.786

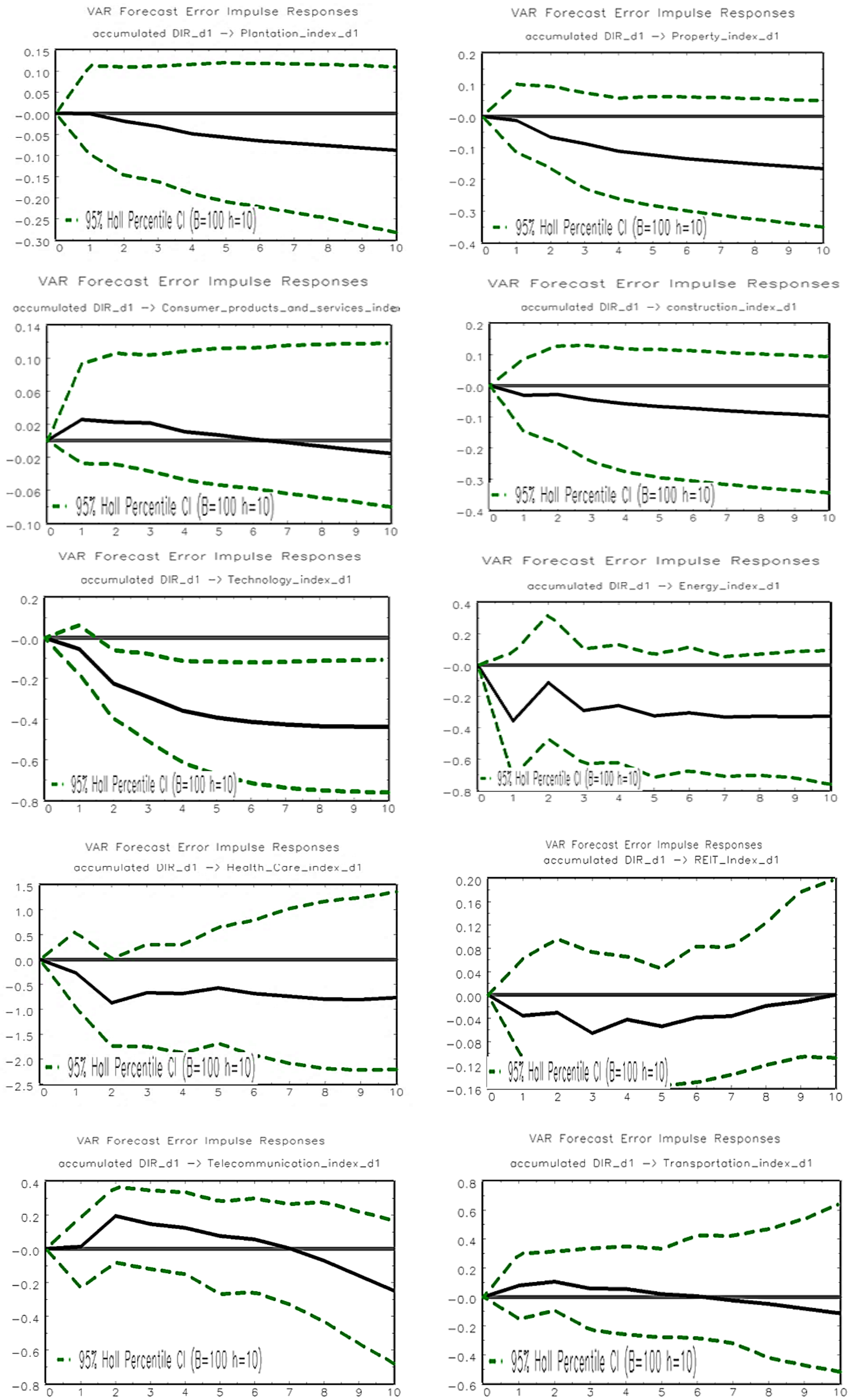
Notes: *, ** and*** represent significance at 10%,5% and 1% levels, respectively.
Source: Author's work

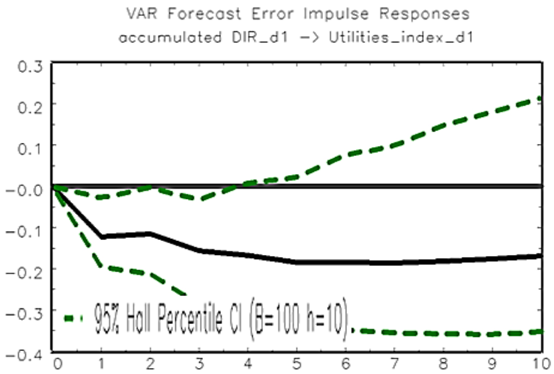
Impulse Response Function

Here, the focus is on how domestic and foreign monetary policies influence 13 sectoral indices. The analysis will use a 2-lag identification to assess the impact of time lags on monetary variables. These confidence bands are constructed using Hall's bootstrap method, based on a 95% confidence interval. Using 8-12 lags in the impulse response could cause issues with six sectoral indices due to limited data and insufficient coefficients. JMULTI software will be used to compute the impulse response as follows.






VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Plantation_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Property_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Consumer_products_and_services_index
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> construction_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Technology_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Energy_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Health_Care_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> REIT_Index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Telecommunication_index_d1
95% Hall Percentile CI (B=100 h=10)
VAR Forecast Error Impulse Responses
accumulated DIR_d1 -> Transportation_index_d1
95% Hall Percentile CI (B=100 h=10)



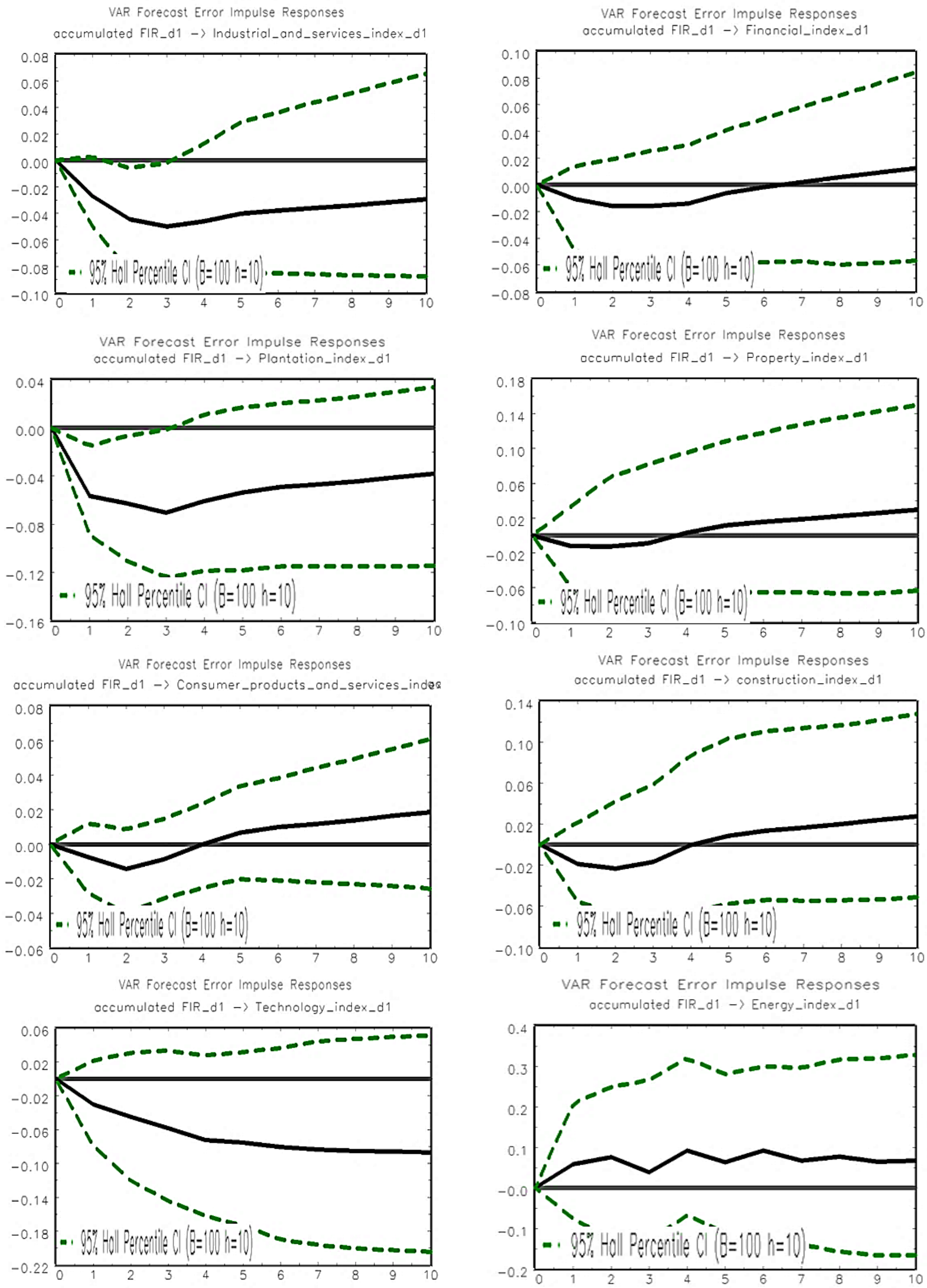
Source: Author's work

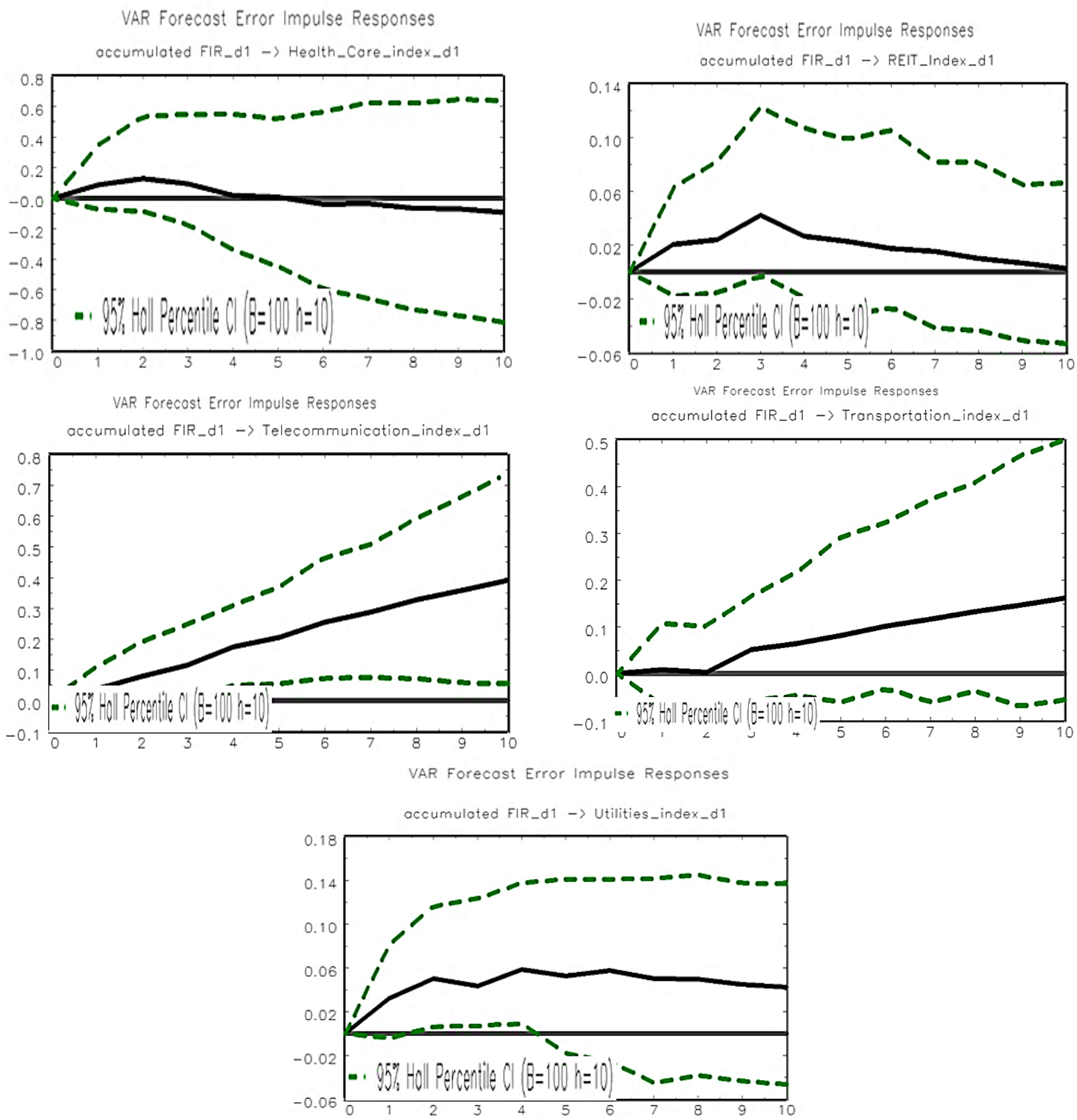
Figure 4. Result of domestic monetary shocks from the sectoral stock index

Figure 4 shows sectoral stock index responses to changes in the domestic interest rate (DINR) in Malaysia, revealing varied sectoral impacts. The Industrial & Services Index showed a significant negative response from months 3-10 (-0.09% in month 2), while the Financial Index had a minor, insignificant negative response (-0.03% in month 2) with limited impact from interest rate. The Plantation Index exhibited a negative but insignificant response (-0.01% in month 2), influenced more by factors beyond interest rate. The Property Index also showed a negative, insignificant relationship (-0.06% in month 2), likely driven by external factors, while the Consumer & Services Index had a mixed response, positive from months 1-5 and negative from months 6-10, though insignificant (0.03% in month 1). The Construction Index experienced a negative but insignificant response (-0.003% in month 2), and the Technology Index showed a negative response across all periods, significant from months 3-10 (-0.4% in month 4). The Energy Index had a negative but insignificant response (-0.3% in month 6), primarily driven by foreign energy supply, while the Healthcare Index showed a negative but insignificant response (-0.5% in month 5), possibly influenced by COVID-19. The REIT Index initially showed a negative, insignificant response (-0.02% in month 8), while the Telecommunication Media Index showed a mixed, insignificant response (0.02% in month 2). The Transportation & Logistics Index had a positive response from months 1-6, followed by a negative response from months 6-10, remaining insignificant overall (0.1% in month 2), and the Utilities Index experienced a negative significant response initially but became insignificant later (-0.1% in month 2). In summary, sectoral responses to domestic interest rates varied, with some sectors showing significant impacts and others influenced more by other factors.

Figure 5 presents the accumulated responses of various Malaysian sectoral stock indices to foreign interest rate changes from the country's top 5 trading partners. The Industrial & Services Index showed a negative relationship with foreign interest rates, significant only in month 2 (-0.05% in month 3), suggesting a limited impact. The Financial Index showed a negative response in months 1-6, then turned positive in months 7-10, but remained insignificant (-0.01% in month 4). The Plantation Index had a significant negative response in months 1-3 (-0.07% in month 3), becoming insignificant thereafter. The Property Index showed a negative but insignificant response throughout the period (-0.01% in month 1), indicating limited influence. The Consumer & Services Index displayed a negative response from months 1-4, followed by a positive response, but remained insignificant (-0.015% in month 2). The Construction Index exhibited a similar pattern, characterised by negative responses in the first four months and positive responses in the subsequent six months, yet it remained insignificant (-0.02% in month 2). The Technology Index showed a consistent negative response throughout, with insignificant results (-0.07% in month 4). The Energy Index exhibited a positive but insignificant relationship across the entire period (0.05% in month 7), reflecting limited influence from foreign interest rates. The Healthcare Index showed a positive response from months 1-5, followed by a negative response from months 5-10, with insignificant results (-0.02% in month 2). The REIT Index followed a positive trend throughout the period. However, it remained insignificant (-0.04% in month 3), while the Telecommunication

Media Index showed an insignificant positive relationship in months 1-3 and a significant positive relationship in months 3-10 (0.2% in month 4). The Transportation & Logistics Index exhibited a consistent positive relationship with foreign interest rates (0.05% in month 3), though it was insignificant, likely influenced by external production and border constraints. The Utilities Index showed a positive relationship throughout, with significant responses in months 2-4 (0.06% in month 6), but became insignificant thereafter, possibly due to energy-sector dynamics and temporary shocks. These results were analysed using the same method as described by Zaidi (2011), Karim and Karim (2016), Zaidi et al. (2017), and Zaidi and Fisher (2010).





Source: Author's work

Figure 5. Result of foreign monetary policy on the sectoral stock index

The findings suggest that foreign monetary policy, particularly changes in the US Federal Funds Rate (FFR), exerts a greater and more immediate influence on Malaysia's sectoral stock markets than domestic monetary policy. This reflects Malaysia's high degree of financial openness and integration into global capital markets, where shifts in major economies' monetary policy are quickly transmitted through cross-border investment flows, interest rate differentials, and exchange rate movements. Firstly, US interest rate changes often drive global stock index movements by influencing international liquidity, credit conditions, and investor risk appetite, as referred to by Kahler (2008). These global dynamics affect capital allocation decisions across countries, leading to synchronised stock market reactions, especially in emerging markets. Secondly, investors routinely compare the relative profitability and risk-adjusted returns of different countries' stock markets. A rise in US interest rates increases the opportunity cost of investing in foreign equities, prompting reallocation of funds back to the US and exerting downward pressure on Malaysian stock prices. This capital rebalancing disproportionately impacts sectors with high foreign ownership or those sensitive to external financing. Thirdly, as a small open economy, Malaysia's equity markets are highly susceptible to shifts in global sentiment and international monetary policy. Even when domestic policy remains unchanged, a tightening of foreign policy can indirectly affect

Malaysia through exchange rate depreciation, higher import costs, and reduced investment confidence. This highlights the theoretical importance of incorporating both external and internal monetary drivers in models of financial market behaviour in emerging markets. Overall, these findings challenge the traditional view that domestic monetary policy is the primary driver of financial conditions and instead support a more nuanced framework that acknowledges the dominant role of global financial cycles in shaping sectoral market dynamics, as referred to by Ha (2021) and Ali and Hatta (2013).

Robustness Checking

In the SVAR model, the 13 substitution stock indices were ordered correctly, revealing that the Malaysian substitution sectoral index is more influenced by foreign monetary shocks and by significant terms, and is more affected by domestic monetary shocks in terms of magnitude. This sectoral index is considered block exogenous, as confirmed by the checks, and this result forms the basis of the analysis.

Focusing on the Impulse Response Function (IRF), the analysis shows that four sectoral stock indices are significantly affected by foreign monetary policy shocks in both the short- and long-term. In contrast, three sectoral stock indices are significantly influenced by domestic monetary policy shocks over both periods. The model's structural identification aligns with theoretical expectations, though it remains unstable across circumstances, with the indices primarily affected by their own shocks. Thus, there is a heterogeneous effect of monetary policy shocks have heterogeneous effects across sectoral stock indices.

Conclusion

This paper examined the impact of domestic and foreign monetary policy shocks on the sectoral stock prices in Malaysia using a Structural Vector Autoregression (SVAR) approach. The analysis revealed how changes in domestic and international monetary policies (interest rate) adjustments by the Bank Negara Malaysia (BNM) and major foreign central banks like the US Federal Reserve, and the combination of other countries interest rates, which were top 5 trading partners with Malaysia, including the United States, affect the stock price of different sectors in the Malaysian economy. Specifically, the Impulse Response Function (IRF) shows the impact of domestic and foreign monetary policy shocks on sectoral stock indices in Malaysia.

The main findings of the study can be divided into two aspects. First, the results reveal that four sector indices — industrial and services, plantation, telecommunications, and utilities — were significantly affected by foreign interest rates. This analysis shows that the results are consistent with the theory for specific sectoral stock indices, while other factors may influence others. Second, regarding domestic interest rates, three sector indices — industrial and services, technology, and utilities — showed significant responses. As mentioned above, it is like the explanation provided by foreign monetary policy.

The finding suggests that Bank Negara Malaysia (BNM) should be aware of the varying impacts of foreign and domestic interest rates on different sectors. As sectors such as industrial, services, plantation, telecommunications, and utilities are significantly influenced by foreign interest rates, BNM should consider these sectors when formulating monetary policies. Moreover, recognising the significant impact of domestic interest rates on sectors such as industry, services, technology, and utilities can help BNM craft targeted policy interventions to mitigate adverse effects on vulnerable sectors. BNM, as usual, cannot implement decisions that affect all sectoral stock indices equally. While theory suggests that monetary policy should significantly impact all sectors, creating a complete transmission channel via the stock market, this study highlights three sectoral indices most influenced by monetary policy. However, these sectors may not be of primary concern to BNM. For investors, the results highlight the importance of closely monitoring both foreign and domestic interest rate changes to adjust their portfolios strategically. Sectors such as telecommunications, technology, and utilities respond significantly to domestic interest rate shifts, offering opportunities for investors to align their investments with sectoral trends driven by monetary policy changes.

Nevertheless, the study has some limitations. Firstly, the study's findings are limited to the specific period analysed. The impact of monetary policy shocks may vary under different economic conditions or in future periods, and the role of interest rates can differ across different time periods in Malaysia. While the research does not delve deeply into the causal relationships, particularly in terms of how potential feedback effects between sectoral stock indices and domestic or international monetary policies, in this study focus on how domestic and international monetary policy directly affects sectoral stock indices; there have been studies that include other monetary policy effects on several sectoral stock indices that can refer to Bhatti et al. (2015).

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Author contributions
Both authors contributed equally to the conception, design, analysis, and interpretation of the study, as well as to the drafting and revision of the manuscript. Both authors have read and approved the final version of the manuscript.

Use of AI tools declaration
The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest
The authors declare no conflicts of interest.

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Thrift-growth nexus for the regional comprehensive economic partnership countries

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Abstract

Purpose — This study examines the relationship between savings and economic growth, accounting for the mediating role of financial development across the selected Regional Comprehensive Economic Partnership (RCEP) countries.

Methods — Using a panel data set spanning 1986 to 2022, the long-run interaction among the variables is investigated with panel cointegration methods that account for cross-sectional dependence. Moreover, the associated long-run elasticities were estimated using the augmented mean group estimation method. The causal nexus was examined for each country in the sample.

Findings — In addition to the presence of a long-run relationship, the findings revealed that both thrifts and growth have a positive influence on each other in the long run. In addition, bidirectional causality tends to exist between thrifts and growth.

Implication — Since the findings disclose the validity of two mainstream macroeconomic views, policymakers should rely on developing economic policies aiming at fostering thrift and economic growth, which may include support of institutional quality and financial deepening in those economies.

Originality — The originality and added value of the study stem from the development of a new perspective, particularly in the examination of causal relationships. Furthermore, this is one of the primary efforts focused on the RCEP bloc, which has significant potential in terms of trade, finance, thrifts, and economic size in the contemporary world economy.

Keywords — Economic growth, panel data, RCEP countries, thrifts.

Introduction

In the realm of macroeconomics and development economics, the interplay between thrift and economic growth is considered a conventional subject. The theoretical underpinnings of this phenomenon encompass two pivotal macroeconomic perspectives: the Classical and the Keynesian. The savings-led growth hypothesis, also known as the Classical view, argues that thrift is the primary driver of economic growth. Hence, higher thrifts lead to higher short-run growth. However, as diminishing marginal returns to the accumulable factors of production set in, the long-run growth rate of income per worker will slow. It is also postulated that, with analogous savings rates, population growth, and technology, economies with low income per worker would grow faster than those with initially higher income or capital per worker. The flip side of the coin argues

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that, rather than thrifts, income and aggregate demand are the power horses of the overall economy. Based on the Keynesian view, savings could be enhanced by boosting disposable income, which is then allocated between consumption and thrift. Thus, it is essential to implement expansionary monetary and fiscal policies to stimulate aggregate demand.

Apart from the foregoing theoretical discussions, the topic is also notable for developmental aspects. For developing economies, one of the primary long-run objectives is to combat income disparities with developed economies. One possible solution to overcome this issue is to accelerate capital accumulation. In this regard, thrifts should not only be encouraged but also directed toward productive investments to stimulate effective demand. Thus, the necessity for the mediating role of the financial sector arises. In an efficiently functioning financial system, thrifts can be rapidly directed to feasible investment projects.

On the other hand, income growth might lead to greater thrift. In this setting, a well-functioning financial system can channel the excessive amount of income into the thrifts by means of various financial instruments. Besides the economic policies, either in favour of growth or in favour of savings, policymakers should account for the essentiality of the institutional regulations (in terms of financial inclusion and financial development) in achieving developmental goals.

Except today, some major Asian economies, most of the developing and less-developed nations have failed to narrow the income disparities with the industrialised countries. Thus, Asian economies have become the leading economies in terms of global trade and production hubs. In this regard, the fifteen Southeast Asian and Pacific countries (ten of which are members of ASEAN) signed the "Regional Comprehensive Economic Partnership" (RCEP) agreement on 15 November 2020, forming the largest free trade zone beyond the World Trade Organisation (WTO). It should be highlighted that the RCEP agreement covers almost 30% of the world population and accounts for 30% of the global GDP (Flach et al., 2021). To this end, Figure 1 presents the dynamics of real gross domestic savings and real GDP for the selected RCEP economies, which are also incorporated into the empirical analysis.[1] Accordingly, both variables tend to decline in the Asian countries of the sample due to the adverse severe effects associated with the emergence of the Asian financial turmoil in 1997. From then onwards, both variables are inclined to rise over the years, except for the global economic crisis, which emerged in the US economy in 2008, and the Covid-19 pandemic.

Source: Author calculation

Figure 1. Dynamics of Real Gross Domestic Product and Real Gross Domestic Savings
Note: Variables are expressed in natural logarithmic form.

[1] See Appendix 1 for the list of countries.

Building on prior theoretical debates, empirical research offers mixed results, supporting either the Classical or Keynesian view. Owing to differences in samples and methods, no consistent relationship has been established. The empirical studies that are in favour of the Classical approach suggest that savings precede economic growth. Recent empirical studies by Tang and Tan (2017), Nguyen and Nguyen (2017), Patra et al. (2017), and Chakraborty (2023) have provided evidence in support of the savings-led growth hypothesis. Conversely, the argument that economic growth precedes the accumulation of thrifts, known as the Keynesian view, has also garnered substantial support among scholars. The compelling evidence presented in the studies by Anoruo and Ahmad (2001), Agrawal (2001), Baharumshah et al. (2003) and more recently, Van Wyk and Kapingura (2021) and Brueckner et al. (2023) substantiates this assertion. Apart from the studies that delve into the Classical or Keynesian approach individually, the empirical corpus also provides a range of evidence. In this context, some empirical studies have concluded the validity of both methods (Adeleke, 2014; Jouini, 2016; Singh, 2010; Šubová et al., 2024). Strikingly, some empirical studies have declared no relationship or even a negative relationship (Chigozie & Omolade, 2021; Hundie, 2014; Joshi et al., 2019).

Despite the lack of direct research on the RCEP countries, there is a glut of studies that focus on individual RCEP countries or groups of Asian countries covered by the RCEP agreement. For instance, Tang and Chua (2012) and Tang (2015) confirmed the existence of a reciprocal feedback mechanism for the Malaysian economy. In contrast, Lean and Song (2009) reached the same conclusion for the Chinese economy, and their findings acknowledged the growth-inducing effects of thrift and investment. In a similar vein, Liu and Ma (2022) verified the Classical view for East and South Asian countries by utilising panel data analysis across a large group of countries from 1960 to 2021. It should also be noted that some studies have declared no causal link between thrifts and economic growth. Accordingly, Tang (2009) reported the insensitivity of the causal nexus to the selection of the technique, whether parametric or nonparametric. The findings, based on an analysis of quarterly time-series data spanning 1991: Q1 to 2006: Q3, indicated that the two variables are independent of each other. Analogously, Sothan (2014) derived no correlation between thrifts and growth for Cambodia.

On the other hand, the empirical literature also includes cross-country studies involving some Asian countries that are part of the RCEP agreement to date. However, the discoveries in those studies do not reveal a uniform tendency due to issues with the sample, data span, and methodology. For instance, Agrawal (2001) investigated the direction of the nexus for seven Asian countries, and, except for Malaysia, the Keynesian hypothesis is vindicated for the rest of the sample. In the latest evidence, Athukorala and Suanin (2024) have declared the validity of the Keynesian view for a group of Asian economies. Baharumshah et al. (2003) scrutinised the factors influencing thrifts across Asian countries and found that thrifts and growth are uncorrelated, except in Singapore, where thrifts precede growth. Tang and Ch'ng (2012) examined the association between thrift and growth across the five ASEAN founding economies from 1970 to 2010. The findings revealed a causal nexus between thrifts and economic growth, suggesting that the Classical view holds for those ASEAN economies. In recent research, Tang and Tan (2017) examined the interaction between thrifts and growth across seven East Asian economies and documented the existence of bidirectional causality between the tandems. Despite the plethora of empirical studies that specifically address the savings-growth nexus across a wide array of Asian countries, both individually and in terms of groupings, no empirical research has been undertaken that directly focuses on the RCEP countries. Accordingly, the present study aims to fill this gap and contribute to the existing empirical corpus by deploying novel panel time-series methods to address the nexus between thrifts and growth. The following section will rigorously discuss the empirical strategy.

Methods

This paper is founded on the approach proposed by Odhiambo (2008) to expose the dynamic interaction between real gross domestic savings and real GDP by accounting for the following linear equations:

$$Y_{it}=\gamma_{0i}+\gamma_{1i}S_{it}+M2_{it}+\varepsilon_{it} \tag{1}$$
$$S_{it}=\gamma_{0i}+\gamma_{1i}Y_{it}+M2_{it}+\varepsilon_{it} \tag{2}$$

Where Y_{it} denotes real GDP, Sit denotes real gross domestic savings, and M2it denotes financial depth, proxied by the real broad money supply. Finally, ε_{it} represents the conventional disturbance term, where superscripts i and t denote the country and time periods, respectively. Due to the limited availability of suitable data, empirical analysis is conducted for the selected RCEP countries using an annual panel data set from 1986 to 2022. All the data regarding the variables above were procured from the World Bank every year[2]. Except for the broad money supply, the rest of the variables are expressed in current US$. It should also be noted that all the variables were deflated by using the consumer price index.

Given the baseline model specifications, the projection of the current study scrutinises the validity of the following hypotheses:
H_1: Real gross domestic savings positively influence real GDP in the RCEP countries.
H_2: Real GDP positively influences real gross domestic savings in the RCEP countries.

In accordance with the foregoing main hypotheses of the present study, the expected signs in estimating the elasticities for real gross domestic savings and real GDP are positive. On the other hand, the anticipated proxy for financial depth is expected to be positive, thereby aligning with the observations of McKinnon (1973) and Shaw (1973). This is because economic development has been shown to stimulate growth by accelerating thrifts through enhanced efficiency in financial intermediation services (Odhiambo, 2008).

The empirical treatment commences with cointegration analyses, which reveal the dynamic interactions among the variables of interest. Anchored on the error-correction mechanism (ECM), Westerlund (2007) suggests the following data-generating process:

$$y_{it}=\varphi_{1i}+\varphi_{2i}t+z_{it} \tag{3}$$

Where φ_{1i}, $\varphi_{2i}t$ represent the deterministic terms, while z_{it} shows the stochastic component of equation 1. To construct the group and panel test statistics, Westerlund (2007) derives the following error-correction representation:

$$\Delta y_{it}=\delta_i' d_t+\alpha_i y_{it-1}+\theta_i' x_{it-1}+\sum_{j=1}^{p_i}\alpha_{ij}\Delta y_{it-j}+\sum_{j=0}^{p_i}\gamma_{ij}\,\Delta x_{it-j}+e_{it} \tag{4}$$

Where d_t denotes the deterministic components, which are embedded in the vector of parameters (δ_i'). Moreover, β_i is the error-correction term, and p_i denotes the lag length, which is allowed to vary across the individual units. By the estimation of equation 4 for each unit, the following group mean test statistics are obtained (Westerlund, 2007):

$$G_\tau=N^{-1}\sum_{i=1}^{N}\frac{\hat{\alpha}_i}{\sigma(\hat{\alpha}_i)} \text{ and } G_\alpha=N^{-1}\sum_{i=1}^{N}\frac{T\hat{\alpha}_i}{\hat{\alpha}_i(1)} \tag{5}$$

Where σ shows the standard error of the error-correction term, in a similar vein, Westerlund (2007) calculates the panel test statistics in the following form:

$$P_\tau=\frac{\hat{\alpha}}{\sigma(\hat{\alpha})} \text{ and } P_\alpha=T\hat{\alpha} \tag{6}$$

Even though Westerlund (2007)'s ECM-based test works efficiently under the existence of CD and slope heterogeneity, it requires variables to be integrated at the same order. In this regard, Westerlund (2008) develops the Durbin-Hausman (DH) type test, which is feasible to the extent that independent variables are integrated at different orders, while the dependent variable must contain a unit root. Anchored on the common factor approach for cross-sectional dependency, Westerlund (2008) suggests a pair of statistics regarding the homogeneity and heterogeneity of the autoregressive component (Φ_i), which is obtained by using the common factor

[2] (World Bank, 2023). World Development Indicators. https://databank.worldbank.org/source/world-development-indicators (accessed 20.10.2024). See Appendix 2 for the definitions and sources of data used in this paper.

for the disturbance term. To this end, for the homogeneity and heterogeneity of the autoregressive parameter, Westerlund (2008) develops the following two types of DH test statistics:

$$DH_g = \sum_{i=1}^{n} \widehat{S}_i \left(\tilde{\emptyset}_i - \widehat{\emptyset}_i\right)^2 \sum_{t=2}^{T} \hat{e}_{it-1}^2 \text{ and } DH_p = \sum_{i=1}^{n} \widehat{S}_n \left(\tilde{\emptyset} - \widehat{\emptyset}\right)^2 \sum_{i=1}^{n} \sum_{t=2}^{T} \hat{e}_{it-1}^2 \quad (7)$$

Where DH_g denotes the group-mean DH statistics, while DH_p denotes the panel statistics, respectively.

Of investigate parameter elasticity, the present paper utilises the Augmented Mean Group (AMG) estimator. Eberhardt and Teal (2010) constructed a two-step estimator, which handles CD by accounting for the standard dynamic process. Moreover, this estimator produces efficient results in small-sample settings. Eberhardt and Teal (2010) derive the AMG estimator by utilising the following two equations:

$$\text{AMG (i) } \Delta y_{it} = b'\Delta x_{it} + \sum_{t=2}^{T} c_t D_t + e_{it} \Rightarrow \hat{c}_t \equiv \widehat{\mu}_t^{\bullet} \quad (8)$$

$$\text{AMG (ii) } y_{it} = \alpha_i + b_i' x_{it} + c_i t + d_i \widehat{\mu}_t^{\bullet} + e_{it} \quad (9)$$

Equation 8 represents the first step for the AMG estimator, which points out the first-differenced pooled regression by including the time dummies ($\widehat{\mu}_t^{\bullet}$). Equation 9 represents the second step, in which deterministic trends for each cross-sectional unit are included in the regression. By the estimation of each equation, the following AMG estimator can be derived for each cross-sectional unit:

$$\widehat{b_{AMG}} = N^{-1} \sum_{i=1}^{N} \widehat{b}_i \quad (10)$$

The present study proposes a novel approach to causality analysis that differs from the empirical approaches suggested thus far. In this regard, Juodis et al. (2021) proposed a strategy that yields efficient outcomes even in the presence of CD, despite the potential for homogeneity or heterogeneity in panel data. In addition, this test has a power advantage for large N and small T cases (Xiao et al., 2023). By establishing a linear restriction on causation parameters, examination of the null hypothesis is done by the pooled least-squares estimator Juodis et al. (2021):

$$\hat{\beta} = \left(\sum_{i=1}^{N} X_i' M_{Zi} X_i\right)^{-1} \left(\sum_{i=1}^{N} X_i' M_{Zi} y_i\right) \quad (11)$$

Where $M_{Zi} = I_T - Z_i\left(Z_i' Z_i\right)^{-1} Z_i'$. Since this estimator suffers from Nickell bias, Juodis et al. (2021) introduced the following bias-corrected version of the Wald test (Xiao et al., 2023):

$$\widehat{W}_{HPJ} = NT\tilde{\beta}' \left(\hat{J}^{-1}\hat{V}\hat{J}^{-1}\right)^{-1} \tilde{\beta} \quad (12)$$

Where $\hat{J} = \frac{1}{NT} \sum_{i=1}^{N} X_i' M_{Zi} X_i$ and $\tilde{\beta}$ is the estimator that wipes out the Nickell bias under the homogeneity restriction. In accordance with the foregoing methodological presentation, the following section provides the empirical findings on the nexus between thrifts and economic growth, incorporating the baseline model specifications.

Results and Discussion

Before a more detailed empirical investigation, descriptive statistics for the variables employed in the study are reported in Table 1. The mean and median values of each variable are closely aligned, suggesting that the estimation of the coefficients may result in more reliable outcomes. Meanwhile, the series of financial depth displays greater variability than those of real GDP and real gross domestic savings, reflecting differences in their maximum and minimum values. Since the mean exceeds the median, the skewness statistic is positive. Panel B is devoted to the computation of the correlation matrix. Evidently, a high degree of correlation is observed between real gross domestic savings and real GDP across the sample. Moreover, the broad money supply is highly correlated with real GDP.

Table 1. Summary Statistics and Correlation Matrix

Panel A: Summary Statistics			
Variables	Y	S	M2
Observations	370	370	370
Mean	22.302	21.130	25.376
Median	22.080	20.980	24.640
Standard Deviation	1.341	1.421	3.512
Minimum	19.449	18.551	19.764
Maximum	25.649	24.882	31.588
Skewness	0.477	0.480	0.305
Panel B: Correlation Matrix			
Variables	Y	S	M2
Y	1.000		
S	0.975	1.000	
M2	0.725	0.688	1.000

Note: Computation of the summary statistics for all variables is conducted in their natural logarithms.

It is conventional wisdom to begin the analysis by testing for the slope homogeneity and CD to identify the dynamic interplay across variables. Given that the countries in the sample have recently formed an integration agreement under the RCEP, which aims to expand trade volumes across the signatories, it is plausible that shocks (financial, commercial, etc.) in any country may affect the rest of the integration bloc. Accordingly, the CD tests developed by Pesaran (2004) and Pesaran (2007) were applied, with the respective results presented in the preceding section of Table 2. Apart from the initial test, the remaining tests consistently indicate the presence of cross-country dependence, rejecting the null of independence across all specifications. The results of the homogeneity tests, shown beneath the CD tests, rely on the procedure of Pesaran and Yamagata (2008) and provide strong evidence against slope homogeneity across units for each baseline specification.

Table 2. Cross-Sectional Dependency and Slope Homogeneity Tests

	A-Cross-Sectional Dependency Tests	
	Y	S
CD Test	0.543	-0.757
CD_{LM} Test	11.697*	9.155*
LM Test	155.963*	131.853*
$LM_{adj.}$ Test	15.421*	13.915*
	B-Homogeneity Tests	
	Y	S
$\tilde{\Delta}$ Test	15.926*	18.680*
$\tilde{\Delta}_{adj.}$ Test	16.863*	19.780*

Notes: * indicates the significance level at 1%.

Exploration of cointegration among non-stationary series requires pre-testing to identify the presence of a unit root, so that further estimation may yield spurious results. Table 3 reports the outcomes associated with various types of tests. The outcomes from the test introduced by Im et al. (2003) (hereafter, IPS) produce estimates that are consistent with those of other tests, to the extent that all series become stationary after first differencing. However, in the presence of CD IPS (2003), the test may not generate reliable results. Thus, we employed the so-called second-generation unit root tests at both levels and in first differences of the series. The estimates of both tests reveal analogous results. In this respect, the results of the bootstrap-IPS test indicate that the series are non-stationary at their levels (Smith et al., 2004). This outcome is also verified using the CIPS test proposed by Pesaran (2007). Hence, the results of the CIPS test clearly indicate that the series are integrated of order I (1).

Table 3. Panel Unit Root Tests

	IPS	BootstrapIPS	CIPS	Decision
Y	-0.808	-2.281	-2.465	
ΔY	-10.477*	-4.452*	-4.691*	I(1)
S	-0.773	-2.514	-2.544	
ΔS	-10.149*	-4.564*	-4.529*	I(1)
M2	-0.489	-2.007	-2.415	
ΔM2	-10.719*	-4.717*	-4.818*	I(1)

Notes: * indicates the significance level at 1%.

Having manifested that I (1) process is valid for all series, the long-run interaction among the variables is examined by conducting the panel cointegration tests. The above section of Table 4 presents the results of the panel ECM test, which reports four types of test statistics depending on whether heterogeneity of present. The results warrant that all the test statistics can indicate the existence of cointegration with respect to baseline specification in which real GDP is chosen as the dependent variable. A similar outcome is achieved when real gross domestic savings are selected as the dependent variable. To this end, except for G_{tau} statistics, the rest of the test statistics are inclined to accept the cointegration at a 10% significance level. The section below of Table 4 is dedicated to the results of the DH test. The results of each type of test statistic clearly indicate the existence of a cointegration nexus among the variables of interest for each specification.

Table 4. Panel Cointegration Tests

A. Panel ECM Cointegration Test				
Test Statistics	Y	Cointegration	S	Cointegration
G_{tau}	-5.034**	Yes	-4.554**	Yes
G_{alpha}	-7.321**	Yes	-5.838***	Yes
P_{tau}	-4.290**	Yes	-3.687***	Yes
P_{alpha}	-6.634**	Yes	-4.605***	Yes
B. Panel Durbin-Hausman Cointegration Test				
Test Statistics	Y	Cointegration	S	Cointegration
DH-group	-1.261***	Yes	-1.451***	Yes
DH-panel	-1.879**	Yes	-1.390***	Yes

Notes: ** and *** indicate significant levels of 5% and 10%, respectively.

After determining the presence of cointegration, the next stage is of estimate the parameter elasticities. To this end, the AMG estimator was employed, and the findings are presented in Table 5. As the real GDP becomes a dependent variable, the results firmly endorse the validity of the hypothesis that thrifts drive growth, and this tendency is consistently demonstrated across the entire sample. In this context, the coefficients of real gross domestic savings are positive throughout the countries in the sample. In a similar vein, the panel test statistic is positive to the extent that 1% rise in real gross domestic savings upswings real GDP by 0.751.

Meanwhile, the impact of broad money on real GDP is insignificant and exhibits sporadic tendencies across the sample. Accordingly, rises in broad money supply promote economic growth only in Singapore and the Republic of Korea. In contrast, negative interplay exists between these tandems of variables in Indonesia, Malaysia, and Thailand.

The following segment presents the estimation results, with real gross domestic savings as the dependent variable. The results clearly support the validity of the income-led savings hypothesis across the sample. To this end, the panel test statistics are positive, indicating that a 1% rise in real GDP is associated with a 1.131 increase in real gross domestic savings. Nonetheless, the most prominent effect of income on thrifts was observed in Singapore, where 1a % rise in income causes a 1.311% upswing in thrifts1.311. However, the weakest effect of income on thrifts was observed in Japan, where a 1% rise in income promotes thrifts by 0.695. On the other hand, the results also show that financial depth positively influences the thrift's overall performance. According to the panel test statistics, 1% rise in M2 triggers a real gross domestic savings of 0.100.

Table 5. Augmented Mean Group Estimation Results

Dependent Variable: Y	Variables			
	S (Coef.)	S (t-stat)	M2 (Coef.)	M2 (t-stat)
Australia	0.752	(0.042)**	0.038	(0.049)**
China	0.938	(0.045)**	-0.033	(0.032)**
Indonesia	0.767	(0.032)**	-0.145	(0.049)**
Japan	0.950	(0.093)*	-0.069	(0.143)
Republic of Korea	0.918	(0.036)**	0.082	(0.012)**
Malaysia	0.578	(0.030)**	-0.109	(0.028)**
New Zealand	0.750	(0.058)*	-0.029	(0.045)*
Philippines	0.479	(0.076)*	-0.131	(0.088)
Singapore	0.729	(0.022)**	0.191	(0.020)**
Thailand	0.649	(0.016)**	-0.078	(0.015)**
Panel	0.751	(0.049)**	-0.028	(0.033)**
Wald χ^2 274.98[0.000]*				
RMSE 0.0539				
Dependent Variable: S	Variables			
	S (Coef.)	S (t-stat)	M2 (Coef.)	M2 (t-stat)
Australia	1.096	(0.066)*	0.147	(0.040)**
China	1.030	(0.031)**	0.125	(0.021)**
Indonesia	1.236	(0.040)**	0.261	(0.019)**
Japan	0.695	(0.067)*	0.138	(0.073)*
Republic of Korea	1.110	(0.059)*	0.029	(0.020)**
Malaysia	0.785	(0.116)*	0.237	(0.084)*
New Zealand	1.098	(0.064)*	0.109	(0.050)**
Philippines	1.121	(0.103)	-0.079	(0.035)**
Singapore	1.311	(0.062)*	-0.038	(0.056)*
Thailand	1.158	(0.044)**	0.059	(0.027)***
Panel	1.131	(0.041)**	0.100	(0.038)*
Wald χ^2 758.51[0.000]*				
RMSE 0.0787				

Notes: ***, **, and * indicate significant levels at 1%, 5%, and 10%, respectively. The standard errors are indicated within parentheses, while the p-values of the Wald test are appended in brackets.

Table 6. Panel Granger Non-Causality Test

Direction	Test Statistics	Decision on H_0
S→Y	10.882**	Rejected
M2→Y	2.845	Accepted
Y→S	7.473***	Rejected
M2→S	12.785**	Rejected
Y→M2	0.751	Accepted
S→M2	8.220**	Rejected

Note: ** and *** indicate significance levels of 5% and 10%, respectively.

Having attested to the presence of long-run interaction, our empirical treatment concludes with a panel Granger non-causality test proposed by Juodis et al. (2021). In line with the results from the long-run parameter estimations, the panel Granger non-causality test supports bidirectional feedback between thrifts and growth (Table 6). In this respect, the null hypothesis of the impact of real gross domestic savings on real GDP is rejected at 5% significance level. In contrast, the null hypothesis that real GDP does not cause real gross domestic savings is rejected at the 10% significance level. It should also be noted that there is a bidirectional relationship between financial development and thrift. In this context, the null hypotheses of which neither variable causes the other are rejected at 5% significance level. However, the results show no significant causal nexus between financial development and growth, as the null hypotheses for each case were not rejected at any significance level.

The discoveries in the preceding section attest to the cointegration relationship. The primary objective of this paper is to examine which perspective, Classical or Keynesian, prevails overall. In this respect, the AMG method was utilised, and the findings bolster the validity of both views. The feedback mechanism between thrifts and growth is also supported by the results of Juodis et al. (2021), which show that real gross domestic savings and real GDP are mutually causative. Even though the present paper focuses specifically on the RCEP countries, the findings presented here can be compared with empirical studies that focus on individual RCEP countries or groups of countries. The validity of both mainstream views and the existence of a bidirectional causal nexus align with Tang (2009), Lean and Song (2009), Singh (2010), Tang & Chua (2012), Adeleke (2014), Tang (2015), Jouini (2016), Tang and Tan (2017), Šubová et al. (2024).

It can be inferred that the effect of growth on thrifts is more pronounced as the long-run coefficient of elasticity for real GDP is higher than the coefficient of the real gross domestic savings. Accordingly, income-led policies are more efficient in the long run than savings-led policies. Nonetheless, thrifts are also positively influenced by developments in the financial system. Therefore, the success of the income-led policies is also closely tied to economic growth in most of those countries. Notably, most East and Southeast Asian countries have shifted their development and industrialisation strategies in recent decades. In fact, with the introduction of the export-oriented development strategy in the 1970s and 1980s, most of those countries recorded high output growth rates, which inevitably led to higher disposable personal income. Thus, this has contributed to the development of the financial sector by fostering thrift and increasing demand for financial instruments. In this respect, co-movement or a feedback mechanism between thrifts and growth exists in those countries. Under a well-functioning financial system, the mobilisation of the thrifts into efficient investment projects results in a higher capital accumulation process and ultimately higher economic growth in those countries.

Conclusion

This study examined the causal link between real gross domestic savings and real GDP, a topic of great importance in the terrain of macroeconomics and development economics. By considering the mediating role of financial development, empirical analyses were conducted on the selected countries of the RCEP agreement. Since most of the signatory parties of the RCEP agreement are the leading industrialised economies, empirical research on the nexus between thrifts and growth is both worthy and timely. Under the existence of a cointegration relationship, it is documented that both variables are positively interacting, besides the prevalence of bidirectional feedback. Theoretical considerations confirmed the validity of either the Classical or the Keynesian view for the RCEP countries.

Since the findings verified the validity of both approaches, economic policies can be in favour of saving or of growth-enhancing policies. As previously indicated, the effect of real GDP on real gross domestic savings is greater than the effect of the latter on the former. Therefore, demand-led policies can be more efficient than the policies that aim to stimulate thrift in the long run. In this respect, demand-led policies should be at the forefront of policymakers' minds. Hence, conducting the tools of either monetary policy or fiscal policy in an expansionary way would result in higher disposable personal income. Thus, it can facilitate the growth of thrifts under a well-functioning financial system. With the existence of various types of financial instruments, individuals or households can hold their thrifts in the financial system, enabling borrowers and investors to easily access suitable credit and loans to finance their investment projects. Therefore, policymakers should not rely solely on economic policies but also resort to institutional regulations, especially to improve the performance of the financial sector.

As discussed in the previous section, some structural policy changes might play a crucial role in explaining the mysterious growth rates of some Asian economies in the sample. Thereby, future research can be extended by considering the role of the structural changes in individual economies of the RCEP agreement. Moreover, future research can be improved by including the various determinants of thrift behaviour and economic growth. Accordingly, most of the countries in the sample are prominent global actors in terms of foreign direct investment, foreign trade, and population, which are also key elements of growth and thrift in empirical literature. Finally, further empirical analyses can be upgraded by including the proxies of both conventional economic policies.

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The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest
The authors declare no conflicts of interest.

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Appendices

Appendix 1. List of the Selected RCEP Countries

Australia, China, Indonesia
Japan, the Republic of Korea, Malaysia
New Zealand, Philippines, Singapore
Thailand

Appendix 2. List of Variables and Data Sources

Variables	Definition	Data Source
Real GDP (Y)	Value in current $ deflated by CPI	WDI
Real Gross Domestic Saving (S)	Value in current $ deflated by CPI	WDI
Real Broad Money Supply (M2)	Value in current national currency deflated by CPI	WDI

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Measuring Islamic Banking resilience: A case study of *Nusa Tenggara Barat* Province, Indonesia

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Abstract

Purpose — Islamic banks in *Nusa Tenggara Barat* (NTB) province have experienced positive developments in assets, branches, and financing. This study aims to measure the resilience of Islamic banking in NTB using a composite bank variable and to determine how effectively the institution manages and absorbs various risks.

Method — The data used consisted of monthly data from 2010 to 2023, covering several banking variables, including the Financing to Deposits Ratio (FDR), Non-Performing Financing (NPF), Bank Size (BS), and Third-Party Fund (TPF). The analysis method employed in this study was the early warning system (EWS), utilising a non-parametric signal extraction approach.

Findings — All selected banking variables are used to measure the resilience of Islamic banking in NTB through the composite index of bank (CIB). The signal extraction method provides optimal thresholds for each selected banking variable and for the composite index (CIB). Visualisation results show the interval values that can absorb risk and maintain the resilience of Islamic banking as follows: (1) FDR between 81% and 102%; (2) NPF between 1.29% and 1.89%; (3) BS between 3.79% and 4.59%; (4) TPF between 4.16% and 4.58%; and (5) CIB between 10.66 and 28.14.

Implications — Assessing the resilience of Islamic banking in NTB involves identifying key banking variables to pinpoint sources of risk exposure, determining the optimal time horizon for policy interventions, and setting appropriate thresholds for the surveillance mechanism.

Originality — Currently, the resilience of Islamic banks at the provincial level has not been widely studied, particularly in NTB Province, where there has been a notable increase in Islamic banking offices and assets.

Keywords — Islamic banking, resilience, optimal thresholds, optimal time horizon, early warning system approach

Introduction

As a country that adheres to the concept of an open and developing economy, Indonesia has been significantly affected by external shocks, including financial crises, since the 1990s. The two major financial crises were the Asian Financial Crisis (AFC) in 1998 and the Global Financial Crisis (GFC) in 2008 (Azwar & Tyers, 2015). Since then, Islamic banks in Indonesia have continued to grow. At the provincial level, the development of Islamic banks has shown significant growth, as reflected

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in the increase in each Islamic bank's total assets in each province. Jakarta became a special province, with assets increasing from IDR 491 billion in August 2022 to IDR 554.913 billion in August 2023, a 13% increase. In addition, of the ten provinces that dominate Islamic banking assets nationally, nine of them come from provinces located in the islands of Java and Sumatra. In this case, *Nusa Tenggara Barat* (NTB) is the only province outside the two islands to compete, with asset growth from IDR 20 billion in August 2022 to IDR 22 billion in August 2023, a 10% increase. This indicates that NTB Province can compete with the larger provinces on the two main islands, despite its smaller area and lower economic resources.

In addition to assets, Islamic banks in NTB have also experienced positive developments in terms of branch offices and financing. From 2018 to 2022, the number of Islamic banks' branches in NTB increased from 41 in the first quarter of 2018 to 72 in the fourth quarter of 2022, a 43% increase. Meanwhile, Islamic banks' financing in NTB increased from IDR 132 billion in August 2022 to IDR 189 billion in December 2023, reflecting an increase of 23% in 2023. This development indicates that the regional economy is showing a positive trend and that operational risks of banking are well controlled, allowing Islamic banks to gradually expand their branch networks.

In terms of risk management, several main risk assessments conducted by Bank Indonesia (BI) on Islamic banks focus on financing risk and liquidity risk. In terms of financing, in the first quarter of 2021, the Non-Performing Financing (NPF) rate of Islamic banks in NTB decreased from 1.71% to 1.32% in the first quarter of 2022. In comparison, in terms of liquidity, in the first quarter of 2021, the Financing to Deposit Ratio (FDR) decreased from 104.98% to 98% in the first quarter of 2022. Although the NPF level remains in an ideal condition based on Central Bank regulations, the FDR is still in a less-than-ideal condition. A decrease in FDR aligns with the slow regional economic growth.

The relatively low regional economic growth conditions indicate the need for greater support from the Islamic banking sector in encouraging the development of the real sector in NTB. However, the presence of financing and liquidity risks in Islamic banks requires greater attention, as they create obstacles and undermine resilience in banking operations. Technically, the strength of Islamic banking is in an ideal condition, where some selected banking variables, such as the Non-Performing Financing ratio and financing to deposit ratio, remain within the tolerance thresholds of Bank Indonesia (BI) around a maximum of 2% - 5% and 78-92%, respectively (Bank Indonesia, 2015). Thus, any movement of the selected variable serves as a signal of the banking sector's resilience, reflecting the escalation of risk during banking operations.

Banking sector resilience can be defined as the ability of the banking sector to survive and adapt to both short- and long-term shocks while continuing to fulfil its role in supporting real economic growth (Caldera-Sánchez, 2017; Wiranatakusuma, 2018). Unfortunately, current studies on the resilience of Islamic banks mainly discuss issues on a national and international scale (Maliha & Marlina, 2019; Khan et al., 2019; Pratama & Rizal, 2019; Setyawati et al., 2019; Ghosh et al., 2020; Nugroho et al., 2020; Ahmad et al., 2022; Setyawati et al., 2022; Fitri & Hafiz, 2022; Albaity et al., 2023). Meanwhile, the resilience of Islamic banks within the provincial scope has not been widely studied, especially in NTB Province.

Therefore, this study focuses on the resilience of Islamic banking in NTB by adopting several selected banking variables. The significance of this study is to select leading variables, calculate optimal thresholds, select optimal time horizons for all variables, and compare them with existing regulations and ongoing data. Those attempts finally enable banking practitioners and policymakers to implement surveillance mechanisms as part of a significant effort to achieve and maintain financial system stability.

Methods

This quantitative study analysed the resilience of Islamic banks in West Nusa Tenggara Province from 2010 to 2023, using secondary data from the Financial Services Authority (OJK). The sample period spanned January 2010 to August 2023, utilising four variables, including Financing to

Deposit Ratio (FDR), Non-Performing Financing (NPF), Bank Size (BS), and Third-Party Fund (TPF).

The data were analysed using the early warning system (EWS) method with a signal extraction approach. Technically, the analysis was conducted in Excel 365 using the Islamic Banking Resilience Index (IBRI) modelled by Wiranatakusuma (2018).

The study involved 15 sequential steps to select the leading variables, calculate the optimal threshold, and select the optimal time horizon (Figure 1).

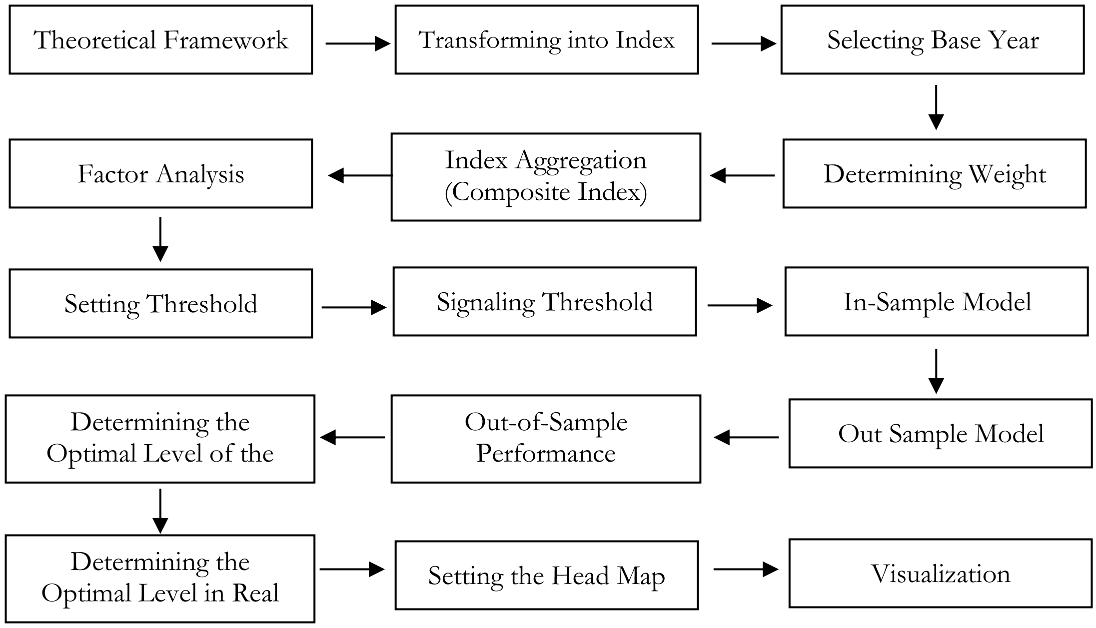


Figure 1. Stages of the early warning system (EWS) method
Source: Modified from Kaminsky and Reinhart (1999)

Transforming Into an Index

Index data provides a numerical method to measure relative changes across variables over time, simplifying data by converting absolute values into relative figures. In this study, index data were used to evaluate the resilience of Islamic banking, with normalisation required to standardise units across variables.

$$I_{it} = \frac{(Xit - \bar{X}i)}{\sigma i} \quad (1)$$

where:
I_{it}= The single index value of variable i at time t.
X_{it}= The value of variable i at time t.
$\bar{X}i$ = The mean value of variable i.
σi = Standard deviation of variable i.

Selecting The Base Year

The base year is the year used as a benchmark or reference when collecting, combining, or analysing data. The selected year is considered the base year, in which there is a fundamental balance and the smallest possible deviation across all index data.

$$S = \sqrt{\frac{1}{N-1} \sum_{i'}^{N} (X_t - \bar{X})^2} \quad (2)$$

Where: S = Standard deviation; Xt = Value of each observation in the sample; X = Sample mean in the base year; N = Number of observations in the sample

$$\bar{X} = \frac{\sum X}{N} \tag{3}$$

Where: X = Sample mean for the base year; $\sum x$ = Number of X values in the sample; N = Number of values in the sample

Determining Weight

The weighted average method is used, in which each variable is assigned a weight reflecting the importance of the parameter it proxies. A higher weight indicates that the role of the variable is becoming increasingly essential in establishing the resilience of Islamic banking.

$$Weighted\ Index_{ij} = \frac{\text{Average of Varianceij}}{Total\ Variance} \tag{4}$$

Aggregate Index (Composite index)

An aggregate or composite index is an index created by combining several measures to capture a broader concept or phenomenon. In this study, one composite index was determined by including four single indices with the formula:

$$\text{Composite Index}_t = w * IFDR_t + w * INPF_t + w * IBS_t + w * ITPF_t \tag{5}$$

Where: I = Variable index, w = Weighted index, t = Time observation

Factor Analysis

This study uses factor analysis to test data and variable coherence, ensuring consistency in relationships and meeting research objectives to produce reliable results (Saaty, 1990). The Hierarchical Consistency Ratio (HCR) was used to prioritise variables in the composite index and to assess coherence, with an ideal HCR of 0.1 or less (10%).

$$HCR = CI\ RI \tag{6}$$
$$CI = \lambda maks - n\ n - 1 \tag{7}$$
$$RI = 1.98 * (N - (n - 1))\ N \tag{8}$$

Where: CR = Consistency ratio; CI = Consistency index; RI = Random index; N : Number of criteria or sub-criteria

The value of the random index can be determined directly from the random index table. In this study, four variables were used, yielding a random index value of 0.9 (Table 1).

Table 1. Random index values.

N	2	3	4	5	6	7	8	9	10
RI	0.00	0.58	0.90	1.12	1.24	1.32	1.41	1.45	1.49

Source: Saaty (1990)

Setting Threshold

Evaluating financial institutions, identifying crises, analysing risks, and setting priorities requires a threshold to filter out less impactful decisions. The Decision-Making Trial and Evaluation Laboratory (DEMATEL) method is effective for analysing interrelated problems by categorising them into causal and effect groups (Fu et al., 2017; Shakeri et al., 2021). DEMATEL clarifies cause-and-effect relationships, reduces complexity, and provides objective recommendations (Seker & Zavadskas, 2017). In this study, DEMATEL-generated thresholds were used to assess Islamic banking resilience (Kaminsky & Reinhart, 1999). Its systematic methodology prioritises criteria based on direct and indirect effects, thereby enhancing the trustworthiness of decision-making (Özdemirci et al., 2023). Before using the DEMATEL method, it is necessary to find the causal relationship of all variables using the Granger Causality method (Granger, 1969) (Table 2).

Table 2. Criteria for the DEMATEL method and Granger causality

DEMATEL Scale	Criteria	Granger Causality
1	Low Influence	>10%
2	Medium Influence	5%-10%
3	High Influence	1%-5%
4	Very High Influence	<1%

Source: Granger (1969), Gabus and Fontela (1973)

$$\alpha = \frac{\sum_{t=1'}^{n} \sum_{j=1'}^{n} [t_{ij}]}{N} \quad (9)$$

$$Ri = \left[\sum_{j=1}^{n} t_{ij}\right] n * 1 = [ti\ ]n * 1 \quad (10)$$

$$Cj = \left[\sum_{i=1}^{n} t_{ij}\right] n * 1 = [ti\ ]n * 1 \quad (11)$$

Where:
α = Threshold value
Ri = Row in matrix T
Cj = Column in matrix T
I = DEMATEL Scale Row
J = DEMATEL Scale Column

Signalling Threshold

The signalling threshold is part of the early warning system (EWS), which helps detect potential vulnerabilities in the economy. This signalling threshold was implemented by assigning codes 0 and 1 to selected variables in a given month based on the calculated threshold multiplier (Berg & Pattillo, 1999). To carry out threshold signalling, consider the following criteria (Table 3).

Table 3. Threshold signaling

Threshold	Criteria Index for a Certain Variable (l_{it})	Signal
Upper	= l_{it} > Threshold Multiplier Value	1 (Yes)
	= l_{it} < Threshold Multiplier Value	0 (No)
Lower	= l_{it} < Threshold Multiplier Value	1 (Yes)
	= l_{it} > Threshold Multiplier Value	0 (No)

Source: Kaminsky and Reinhart (1999), Berg and Pattillo (1999)

In-Sample Model

The in-sample model calculates the estimated trend of Islamic banking resilience using the Hodrick-Prescott filter. The HP separates cycle components of optimising the loss function (Kanas et al., 2012). In the Early Warning System (EWS) approach, setting the lambda (λ) value is critical, as it influences prediction accuracy across different time horizons (e.g., 12 or 24 months) (Kaminsky & Reinhart, 1999; Ravn & Uhlig, 2002; Domala & Kim, 2023). The λ function, a smoothing parameter, was applied as per Ravn and Uhlig (2002)'s formula:

Lambda (λ) = $(14.400 * (\text{Number of Observations in One Month}))^2$ (12)

Out Sample Model

The Out Sample Model was conducted using the crisis-signal matrix framework, or early warning system (EWS), to evaluate the previous signals. In the crisis-signal matrix framework, A is the number of months the variable produces a good signal, B is the number of months the variable produces a bad signal because it sends a false signal, C is the number of months the variable fails to issue a warning signal, and D is the number of months when the variable does not give a signal or there is no crisis (Table 4) (Kaminsky & Reinhart, 1999).

Table 4. True and false signal assessment using the crisis-signal matrix

Items	Stress occurs in the next n months (C = 1) (Pre-stress periods)	No stress occurs in the next n months (C=0) (Pre-stress periods)
Signal (S=1)	A (Number of true imbalance signals)	B (Number of true imbalance signal types two error)
No Signal (S=0)	C (Number of false balance signals-types one error)	D (Number of true balance signals-)

Out Sample Model Performance

At this stage, the selection of the best time horizon was taken as the one with the least Quadratic Probability Score (QPS), representing accuracy, and the Global Square Bias (GSB), representing calibration. The QPS and GBS values range from 0 to 2, where 0 represents perfect accuracy or calibration. The robustness analysis in this study relies on the out sample performance stage to evaluate the reliability of the early warning system model. Accuracy refers to whether forecasted values are sufficiently accurate to reflect the reality being captured and to the closeness of computations or estimates to exact values (Kaminsky & Reinhart, 1999; Berg & Pattillo).

Determining the Optimal Level of Index Value

This stage aims to assess variable index values within optimal, tolerant, stagnant, or vulnerable conditions, with "optimal" as the most preferred state. A condition reflects each variable's characteristics, with "High is Good" for FDR, BS, and TPF. At the same time, "Low is Good" applies to NPF in the Islamic banking sector (Table 5).

Table 5. Definition of optimal, tolerant, stagnant, and vulnerable conditions

A Variable with the Characteristic of "High is Good."	
Optimal	Average ≤ Variable ≤ Upper Threshold
Tolerance	Average > Variable ≥ Lower Threshold
Stagnant	Variable > Upper Threshold
Vulnerability	Variable < Lower Threshold
A Variable with the Characteristic of "Low is Good."	
Optimal	Average ≥ Variable ≥ Lower Threshold
Tolerance	Average < Variable ≤ Upper Threshold
Stagnant	Variable < Lower Threshold
Vulnerability	Variable > Upper Threshold

Calculating the Optimal Level in Real Value

After finding the best model for calculating the resilience of Islamic banking by using data index values, the next step was to convert the model back to the original data.

Using the Head Map

A head map is a data visualisation method that uses different colour representations to display data conditions. The use of various colours enables the identification of the sources of vulnerabilities affecting the resilience of Islamic banking (Table 6).

Table 6. Use of colours in the head map

Vulnerable/Excessive
Tolerance/Healthy Enough
Expected/Optimal/Healthy
Stagnant/Strict/Prudent

Visualisation

The visualisation stage can present research results effectively and make them easier to understand by summarising the key points.

Results and Discussion

The resilience analysis of Islamic banking in West Nusa Tenggara Province began with assessing financial performance metrics, including FDR, NPF, BS, and TPF, which represent banking vulnerabilities. These variables reflect liquidity risk (FDR), credit risk (NPF), market risk (BS), and operational risk (TPF). The next step was to determine the threshold ranges (optimal, tolerant, stagnant, and vulnerable) for each variable to gauge resilience. For "high is good" variables, the optimal range lies between the average and upper threshold; for "low is good" variables, it lies between the average and lower threshold. An Early Warning System (EWS) approach with signal extraction was used to assess the resilience level of each variable.

Transforming into index data is intended to facilitate measuring relative changes, comparing variables or groups, analysing, providing baselines, and measuring relative performance. The following is a detailed explanation of the data from each variable of Islamic banking after converting absolute data into index data (Figure 2).

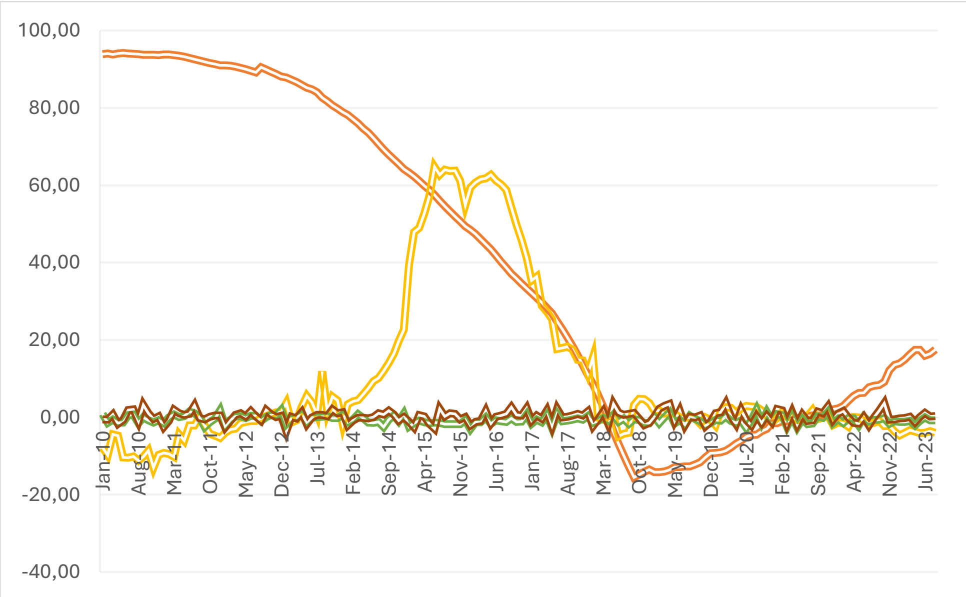


Figure 2. Transformation of Islamic banking index, where: Blue is NPF, Orange is FDR, Red is BS, Green is TPF,

The base year, selected as the year with the smallest standard deviation for each variable, was used as a benchmark or reference for analysing the data (Table 7).

Table 7. Selected base years

Sharia Bank Variables	Base Year
FDR	2021
NPF	2021
BS	2015
TPF	2014

Source: Processed Data, Excel 365

Determining the weight of each variable aimed to provide a relative value to each variable. Determining the weight demonstrates the more critical variables and contributes more to the formation of banking resilience (Table 8).

Table 8. Weight of Each Variable (Decimal)

Variable	FDR	NPF	BS	TPF
Weight	0.156	0.642	0.187	0.015

Source: Processed Data, Excel 365

The creation of a composite index in this study involved combining several variables to measure a broader concept or phenomenon (Figure 3). This composite index was created by combining several variables into a simpler, easier-to-understand aggregate measure. In calculating the composite index, the formula is as follows:

$\textit{Composite Bank Index (CIB)}_t = w * IFDR_t + w * INPF_t + w * IBS_t + w * ITPF_t$

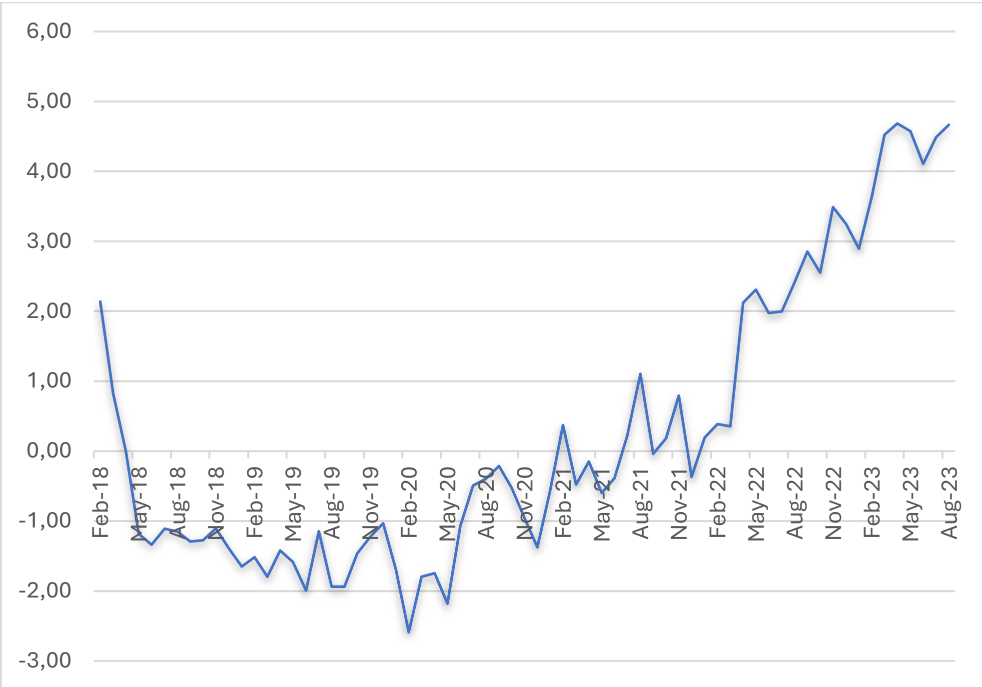


Figure 3. Composite Islamic Source: Processed Data, Excel 365
Where: Blue is the Composite index

The factor analysis test was based on the degree of coherence between the data and the variables. A good level of coherence between variables can be indicated by an R-value of no more than 0.1 (10%). The HCR values of Islamic banking in West Nusa Tenggara were 7%. This indicates that all variables were logically and consistently connected, with no contradictions in the research results (Table 9).

Table 9. Factor analysis

Factor Analysis	Sharia Bank
Consistency Index (CI)	0.16
Random Index (RI)	1.51
Consistency Ratio (CR)	0.11
HCR (Hierarchy Consistency ratio)	7%

Source: Processed by Author

The multiplier threshold was determined based on the Decision-Making Trial and Evaluation Laboratory (DEMATEL) method. This DEMATEL method converts Granger causality values between variables into the DEMATEL scale to address the core problems of complex systems and facilitate decision-making. The value of the threshold calculation was used as the signalling threshold. The results of the calculations in this study were 0.74 for Islamic banking.

At the signalling stage, the threshold was determined by assigning codes 0 and 1 to the index variables for a given month based on the calculated multiplier threshold. Code 0 indicates optimal conditions or values within the ideal threshold, and code 1 indicates values outside the ideal threshold. In addition, this stage is intended to detect threats earlier, provide warnings, and enable rapid preventive or mitigating actions. The upper threshold is 1.20, and the lower threshold is -1.20 (Table 10).

Table 10. Signalling threshold variables of Islamic Banking in January 2018

Base Year/Month	Upper IFDR	Multiplier Threshold	Signal	Base Year	Lower IFDR	Multiplier Threshold	Signal
January 2018	8.59	1.20	1	January 2018	8.59	-1.20	0
Base Year/Month	Upper INPF	Multiplier Threshold	Signal	Base Year	Lower INPF	Multiplier Threshold	Signal
January 2018	0.59	1.20	0	January 2018	0.59	-1.20	0
Base Year/Month	Upper IBS	Multiplier Threshold	Signal	Base Year	Lower IBS	Multiplier Threshold	Signal
January 2018	-1.44	1.20	0	January 2018	-1.44	-1.20	1
Base Year/Month	Upper ITPF	Multiplier Threshold	Signal	Base Year	Lower ITPF	Multiplier Threshold	Signal
January 2018	-2.26	1.20	0	January 2018	-2.26	-1.20	1

Meanwhile, the in-sample model stage was carried out by assigning 0 or 1 to the Hodrick-Prescott (HP) Filter data or index variable forecast data for a given month based on the multiplier threshold value (Table 11). The goal was to calculate the long-term trend of the operational resilience cycle of Islamic banking. Codes 0 and 1 indicate the degree of data suitability with the cycle trend, where code 0 indicates the condition of the variable is at the ideal threshold and vice versa for code 1. The in-sample model is said to be good if it can predict the amount of data in the future and has a goodness-of-fit model close to 0.

Table 11. In-sample model of Islamic Banking variables

Base Year/Month	Upper HP IFDR	Multiplier Threshold	Signal	Base Year	HP Lower HP IFDR	Multiplier Threshold	Signal
January 2018	19.01	1.20	1	January 2018	19.01	-1.20	0
Base Year/Month	Upper HP INPF	Multiplier Threshold	Signal	Base Year	Lower HP INPF	Multiplier Threshold	Signal
January 2018	1.80	1.20	1	January 2018	1.80	-1.20	0
Base Year/Month	Upper HP IBS	Multiplier Threshold	Signal	Base Year	Lower HP IBS	Multiplier Threshold	Signal
January 2018	-0.78	1.20	0	January 2018	-0.78	-1.20	0
Base Year/Month	Upper HP ITPF	Multiplier Threshold	Signal	Base Year	Lower HP ITPF	Multiplier Threshold	Signal
January 2018	0.40	1.20	0	January 2018	0.40	-1.20	0

This out-sample stage was carried out to evaluate the previous signalling using the signal-crisis matrix framework. The evaluation was conducted by accumulating the number of months of indicators that produced positive or correct signals (A), negative or incorrect signals (B), failed to provide warning signals (C), and did not provide signals or were not in crisis (D). The determination of this crisis signal was based on time horizons of 1, 3, 6, 12, and 24 months (Table 12).

Table 12. Evaluation of signalling based on the time horizon

Time Horizon		Upper				Lower			
		A	B	C	D	A	B	C	D
FDR	1 Month	106	20	0	37	3	11	19	130
	3 Months	104	20	0	37	3	11	19	128
	6 Months	101	20	0	37	3	11	19	125
	12 Months	95	20	0	37	3	11	19	119
	24 Months	83	20	37	37	3	11	107	107
NPF	1 Month	63	15	12	73	43	17	8	95
	3 Months	63	15	12	71	41	17	8	95
	6 Months	63	15	12	68	38	17	8	95
	12 Months	63	15	12	62	32	17	8	95
	24 Months	63	15	12	50	21	16	8	95
BS	1 Month	0	0	6	157	0	0	55	108
	3 Months	0	0	6	155	0	0	54	107
	6 Months	0	0	6	152	0	0	53	105
	12 Months	0	0	6	146	0	0	50	102
	24 Months	0	0	4	136	0	0	47	93
TPF	1 Month	0	0	35	128	0	0	26	137
	3 Months	0	0	35	126	0	0	25	136
	6 Months	0	0	33	125	0	0	25	133
	12 Months	0	0	32	128	0	0	24	128
	24 Months	0	0	30	110	0	0	23	117
CIB	1 Month	88	12	6	57	2	19	12	130
	3 Months	88	12	6	55	2	17	24	127
	6 Months	88	12	6	52	2	17	12	127
	12 Months	88	12	6	46	2	17	10	123
	24 Months	78	12	6	44	2	17	10	111

Table 12 evaluates the crisis-signalling effectiveness of key Islamic banking variables across multiple time horizons (1, 3, 6, 12, and 24 months). The variables assessed were FDR, NPF, BS, TPF, and the Composite Bank Index (CIB), each categorised as accurate imbalance signals (A), false signals (B), missed signals (C), and proper balance signals (D).

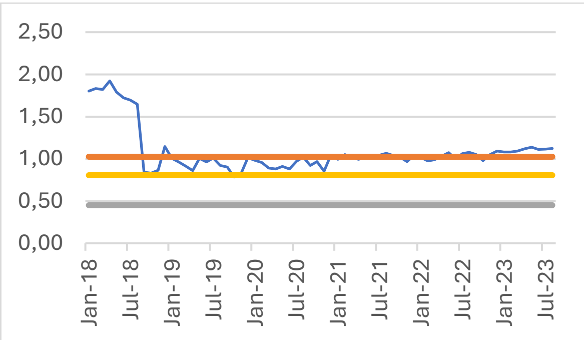
FDR's effectiveness declined with time; it provided 106 correct signals at a 1-month horizon, dropping to 83 at 24 months, while false signals remained stable at 20 across horizons. NPF produced stable, correct signals (around 63) for up to 12 months, declining slightly to 50 at 24 months. BS and TPF were less effective overall, generating minimal positive signals and mainly producing accurate balance signals, especially in shorter horizons. The CIB showed a decrease in positive signals with longer horizons, indicating that a combined index offers a broad perspective but with declining accuracy over time. In summary, FDR and NPF were more effective for short-term imbalance detection, whereas BS and TPF lacked sensitivity in both the short and long term. The CIB, while helpful, also saw reduced effectiveness over time. This indicates that FDR and NPF are exceptionally responsive to resilience challenges in NTB's Islamic banking, while BS and TPF, though cautious, require closer monitoring for early warning purposes.

The out-sample performance stage assessed the prior signalling threshold and the in-sample model stages using the crisis-signal matrix framework. Out-sample model results identified crisis signals based on the optimal time horizon and the smallest Quadratic Probability Score (QPS) for accuracy or the smallest Global Square Bias (GSB) for calibration. The out sample model results are used to detect crisis signals by determining the optimal time horizon, focusing on the minimal Quadratic Probability Score (QPS) for accuracy assessment or the minimal Global Square Bias (GSB) for calibration. This methodology guarantees that the constructed index accurately reflects the dynamics of the financial system and is applicable across diverse economic scenarios (Bespalova, 2015; Handoyo et al., 2020; Gupta & Kumar, 2022). Table 13 presents out-sample model performance for each variable's optimal horizon and GSB/QPS values at the upper and lower thresholds. A shorter optimal time horizon suggests a rapid response to vulnerabilities when risk levels are high.

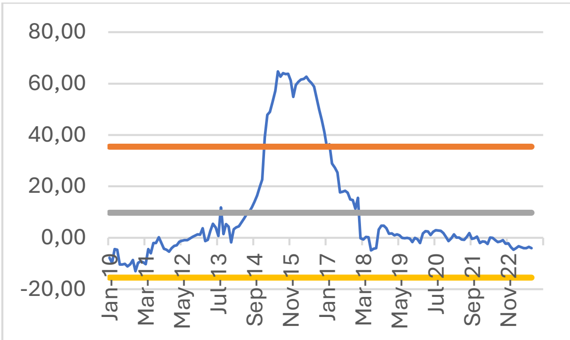
Table 13. Out-sample of Islamic Banking Performance

VARIABLE	TIME HORIZON	QPS	GSB	THRESHOLD
FDR	1 Month	0.8589	0.0067	Upper
	1 Month	0.9080	0.0301	Lower
NPF	1 Month	0.5890	0.0003	Upper
	3 Months	0.6135	0.0067	Lower
BS	1 Month	0.147	0.003	Upper
	6 Months	1.342	0.231	Lower
TPF	1 Month	0.8589	0.1020	Upper
	6 Months	0.684	0.060	Lower
CIB	12 Months	1.053	0.002	Upper
	3 Months	0.9882	0.0031	Lower

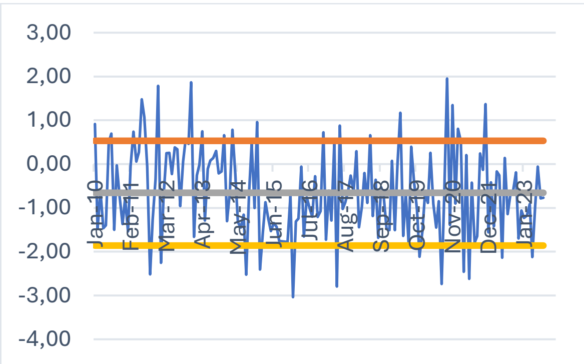
Figure 4 is a visualisation of the individual movements of index variables against static thresholds in Islamic banking. The threshold values indicate that the performance of each variable and the period of its movement are at the optimal threshold (resilience) or exceed the lower or upper thresholds (vulnerable or stagnant). In the context of banking resilience, visualisation is essential as a surveillance tool for policymakers to monitor the development of each variable, mitigate the impact of risk, and prevent the probability of banking risk from increasing.



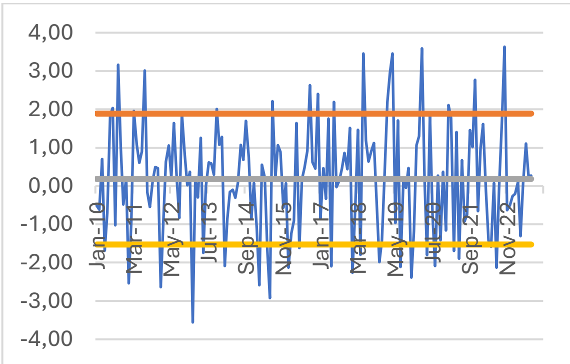
Optimal Level Index IFDR



Optimal Level Index INPF



Optimal Level Index ITPF



Optimal Level Index IBS

Figure 4. Optimal level in the Islamic Banking variable index.

Figures 5 depict the movement of individual variables in the form of their absolute values or actual values (in percentage). The results revealed that each variable fluctuated between optimal (between the ideal thresholds) and non-optimal (stagnant and vulnerable) positions. The optimal position depends on the variable's characteristics—whether "high is good" or "low is good"—and on the choice of scale.

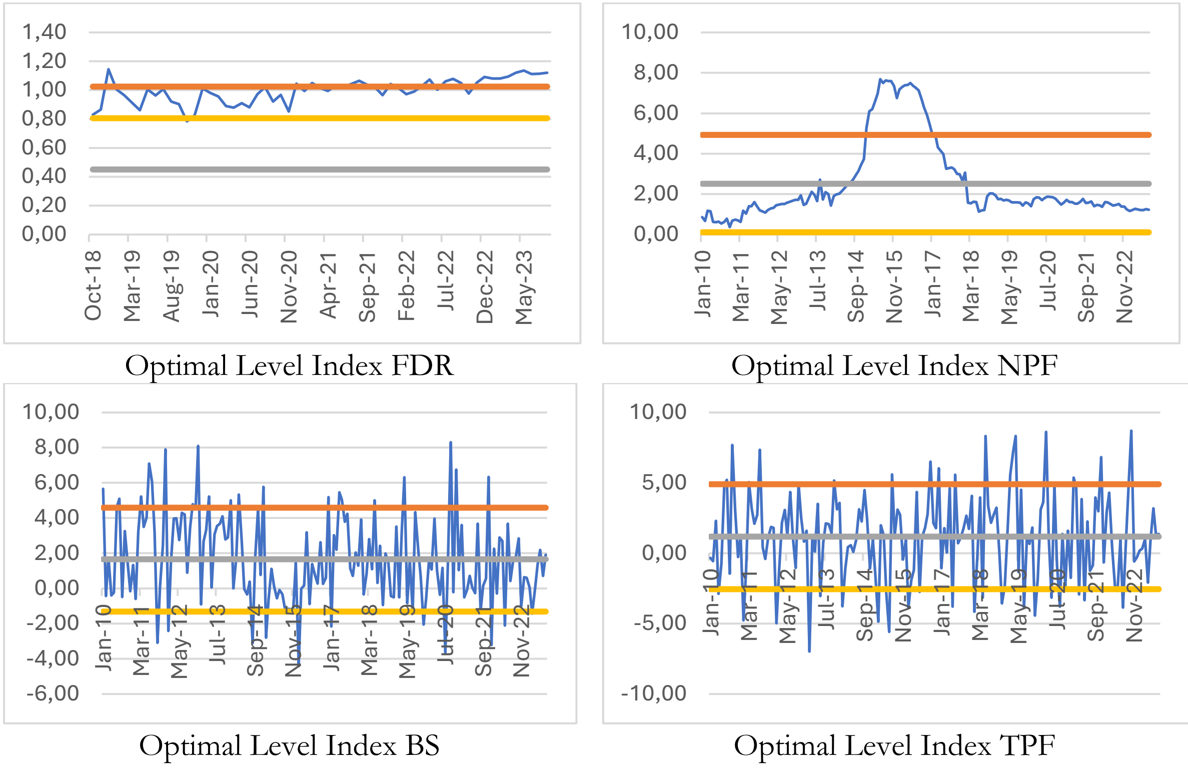


Figure 5. Optimal level in real variables of Islamic banking

In the banking variables, Bank Size (BS), Third-Party Fund (TPF), and FDR variables are coded as "high is good," while only NPF is coded as "low is good." This categorisation affects the position labelled "optimal" in Figure 5. For the “high is good” variable, the optimal variable movement condition is between the middle and upper thresholds.

Table 14 demonstrates the stages of the variable conditions per month during the observation period, illustrated by colour variations. In the context of individual variables, such as FDR, NPF, BS, and TPF, the head map shows the distribution of variable conditions at a given time. The head maps can identify areas with high intensity (vulnerable or stagnant) or low intensity (resilient or tolerant) in the data they represent.

Table 14. Head map of Islamic Banking

Years	Month	FDR	NPF	BS	TPF	CIB
2010	Jan-10	141.30	0.86	5.66	-0.33	32.42
	Feb-10	141.36	0.67	-0.43	-0.57	31.35
	Mar-10	141.29	1.18	1.75	2.31	34.26
	Apr-10	141.36	1.15	-0.40	-2.89	34.04
	May-10	141.42	0.62	-0.23	-0.74	31.10
	Jun-10	141.37	0.61	4.63	4.82	31.15
	Jul-10	141.33	0.64	5.10	5.21	31.31
	Aug-10	141.31	0.54	-0.49	-1.46	30.51
	Sep-10	141.23	0.63	3.25	7.69	31.18
	Oct-10	141.24	0.78	1.42	3.02	31.93
	Nov-10	141.25	0.37	-0.16	-0.27	29.50
	Dec-10	141.22	0.68	1.34	1.23	31.30

Description: Red indicates variables in vulnerable conditions; Yellow indicates variables in the tolerance conditions; Green indicates variables in resilience conditions; Blue indicates variables in stagnant conditions.

This visualisation stage evaluated prior signalling thresholds and in-sample models using the signal-crisis matrix framework. The out-sample model results identified crisis signals based on the optimal time horizon with the smallest Quadratic Probability Score (QPS) for accuracy and Global Square Bias (GSB) for calibration. Additionally, an optimal level assessment determined each variable's status (optimal, tolerant, stagnant, or vulnerable) across Islamic banking, as shown in Table 15. The short-term horizons suggest that risks in the banking sector escalate quickly, necessitating prompt responses to limit potential losses and bolster resilience in NTB's Islamic banking. The significance of an optimal time horizon is to help policymakers strike a balance between preventive and mitigative measures in confronting Islamic banking vulnerability.

Table 15. Visualisation of Islamic Banking Variables

Variable	Optimal	Tolerance	Stagnant	Vulnerable	Time Horizon	QPS	GSB	Threshold
FDR (%)	81 ≤ FDR ≤ 102	45 ≤FDR<81	FDR <81	FDR>102	1 Month	0.85	0.01	Upper
					1 Month	0.90	0.03	Lower
NPF (%)	1.29 ≤ NPF ≤ 1.89	1.59 < NPF ≤ 1.89	NPF < 1.29	NPF > 1.89	1 Month	0.58	0.01	Upper
					3 Months	0.61	0.01	Lower
BS (%)	3.79 ≤ BS ≤ 4.59	2.98 ≤ BI < 4.59	BS <2.98	BS>4.59	1 Month	0.15	0.01	Upper
					6 Months	1.34	0.23	Lower
TPF (%)	4.16 ≤ TPF ≤ 4.58	3.73 ≤ TPF < 4.16	TPF < 3.73	TPF > 4.58	1 Month	0.86	0.10	Upper
					6 Months	0.68	0.06	Lower
CIB	10.66 ≤ CI ≤ 28.14	-6.61 ≤ CI < 10.66	CI < -6.61	CI > 28.14	12 Months	1.05	0.01	Upper
					3 Months	0.99	0.01	Lower

In analysing Islamic banking resilience, calculated threshold conditions were categorised as underestimated, appropriate, or overestimated and compared with government and monetary authority standards for each variable. The FDR was overestimated, exceeding the ideal government range of 78% to 92% by ranging from 81% to 102%, indicating a high liquidity risk due to an imbalance between financing and fund collection. The NPF was within an appropriate range, aligning with the government's maximum of 1.59% to 29%. This indicates that Islamic banking maintained sound financing quality at optimal levels. The Bank Size (BS) was underestimated, falling below the expected 3.64% to 5.12% range, with a value between 2.98% and 4.59%. This suggests limited capacity to fully support economic growth. Similarly, the TPF was also underestimated, with growth from 2.98% to 4.58%, below the target range of 7.32% to 8.15%. Islamic banking should be more innovative and attractive to attract a large number of depositors (Table 16).

Table 16. Visualisation of research results threshold with the threshold of the government and the monetary authority (Monthly Percentage)

Variable	Government Provisions	Research Result	Condition
FDR* (Percentage)	78 ≤ FDR ≤ 92	81 ≤ FDR ≤ 102	Not yet appropriate (Overestimate)
NPF** (Percentage)	NPF ≤ 5	1.59 ≤ NPF ≤ 1.29	Appropriate
BS*** (Percentage)	3.64 ≤ BS ≤ 5.12	2.98 ≤ BS ≤ 4.59	Not yet appropriate (Underestimate)
TPF****	7.32 ≤ TPF ≤ 8.15	2.98 ≤ TPF ≤ 4.58	Not yet appropriate (Underestimate)

*FDR, based on Bank Indonesia provisions 2013; **NPF, based on Bank Indonesia provisions 2015; **INF, based on Regional Inflation assumptions and Bank Indonesia Provisions 2021; ***BS, based on Bank Indonesia provisions 2022; ****TPF, based on the assumptions of the Ministry of Finance of the NTB Regional Office 2023

Conclusion

The study of the resilience of Islamic banking in West Nusa Tenggara Province is essential, given the gradual expansion of its offices and assets. However, rising risks in banking operations can increase pressure, leading to banking vulnerability. Consequently, a prolonged period of imbalance can weaken banking resilience, as its balance sheets cannot absorb the escalating risks. The Early Warning System (EWS) approach, involving 15 integrated steps, helps monitor the resilience by using variables of FDR, BS, NPF, and TPF against their optimal thresholds.

The results show that the optimal time horizon for these variables is mostly under one year (short term), highlighting the need for rapid policy responses to prevent systemic vulnerability and mitigate its impact. This brief time horizon underscores the urgency for prompt decision-making to maintain resilience. The findings also show that aligning resilience levels with regulatory thresholds is vital. Deviations—whether above or below targets—indicate excessive or cautious trends, suggesting the need for adjustments. Thus, resilience requires leading variables to assess risks, an optimal time horizon for policy action and reaction, and a defined threshold to guide monitoring and ensure stability. This study is significant for evaluating the degree of compliance by comparing the stipulated regulatory thresholds with estimated thresholds. Islamic banking in NTB province can effectively control NPF, provided its movements remain within the specified thresholds. However, FDR, BS, and TFT need to be adjusted because they have exceeded regulatory thresholds. Therefore, this study is essential for monitoring operational aspects in the banking industry and would be a significant tool for addressing issues arising from escalating risks and a gradual decrease in resilience levels.

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Author contributions

Both authors contributed equally to the conception, design, analysis, and interpretation of the study, as well as to the drafting and revision of the manuscript. Both authors have read and approved the final version of the manuscript.

Use of AI tools declaration

The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest

The authors declare no conflicts of interest.

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Green bond underdevelopment in emerging economies: Exploring the dynamic roles of institutional quality

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Abstract

Purpose — The green bond market, as a financial innovation, faces institutional quality challenges common to emerging economies. This study examines the influence of institutional quality on the development of green bonds in emerging economies.

Method — The study relied on the panel Fully Modified Ordinary Least Squares (PFMOLS) to estimate long-run panel cointegration relationships between institutional quality and green bond development. The observation consists of twenty-one (21) emerging economies from 2010 to 2023.

Findings — The findings show that institutional factors can hinder the development of green bonds in emerging economies. This is mainly due to the adverse effect of voice and accountability, even though regulatory quality and the rule of law provide firm support. Policy efforts that improve overall institutional quality, along with measures for macroeconomic stability, will benefit green bond development.

Implications — The results recommend that emerging economies strengthen their existing institutional structures. In addition to reforms in voice and accountability, they need to improve enforcement in promoting regulatory quality and the rule of law. Encouraging inclusive institutional quality can build legitimacy and green investors' confidence, thereby supporting long-run green finance outcomes.

Originality — Providing a new and detailed understanding of the green bond market, this study examines the distinct and collective effects of regulatory quality, voice and accountability, and the rule of law on green bond development in emerging economies.

Keywords — Green bond, institutional quality, emerging economies, panel FMOLS, finance outcomes

Introduction

Green bonds have gained prominence within the concept of green finance, which refers to financial commitments aimed at mitigating the adverse effects of climate change on the world's economies (Xing et al., 2022; Zheng et al., 2021). This financial initiative is a global consensus leading from the Paris Climate Change Agreement. As the most significant financing option in the green finance initiative, green bonds are fixed-income securities designed to fund eco-friendly projects (García et al., 2023; Kanamura, 2020; Mertzanis, 2023; Smaoui et al., 2017). Green bonds are expected to benefit countries of the world by way of reduced carbon emissions, improved ecological values and overall sustainable development, but emerging countries have shown unimpressive records of

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green finance growth. Emerging economies are often referred to as latecomers to the green bond market despite their environmental challenges. Data from the International Monetary Fund (IMF) showed that none of the emerging economies has achieved more than 1% of their GDP. Only Mauritius in 2021 recorded the highest proportion of green bond-to-GDP ratio of 0.52%. However, the growth of green bonds is based on financial institutions such as banks, capital markets, and development finance institutions. To support green bonds in emerging economies, these financial institutions must be efficient enough to serve savers and investors. Thus, it highlights the significance of institutional quality in driving green bond growth in emerging economies.

How well a country's institutions can formulate and follow policies, enforce laws, protect property rights, maintain legal order, keep the government accountable, and implement contracts determines its institutional quality (Abaidoo & Agyapong, 2022; Ellahi et al., 2021; García et al., 2023; Kanamura, 2020; Mertzanis, 2023; Smaoui et al., 2017). It includes factors such as handling corruption, effective regulation, political stability, government effectiveness, the rule of law, and citizen engagement in politics. According to institutional theory, organisations function according to established rules and norms and cultural expectations to receive stakeholder validation from governments and regulators, as well as markets and society (Bae et al., 2022; Comyns, 2016). By extension, organisations and investors under institutional pressure to improve environmental and climate performance choose to invest in green finance initiatives while complying with financial regulations and advancing the ecological mitigation agenda. In this regard, institutional quality focuses less on efficiency-seeking behaviour and more on the social, ecological and economic goals of business organisations (Debrah et al., 2024).

Some prominent institutional challenges in emerging economies, like poor regulation, rule of law and voice and accountability, can be understood as systemic issues that manifest in heavy default culture, operational self-insufficiency, absence of adequate accountability and transparency and insufficient regulatory framework (Zheng et al., 2021). When this happens, institutional quality tends to fail to support green bond market development in emerging economies, as it does in the financial development of developed countries (Lisbinski & Burnquist, 2024; Mbulawa & Chingoiro, 2024). The reason is that it sets the framework for financial development, and without this framework, green bonds may face challenges of growing.

Various studies found that economic development and the issuance of green bonds are mainly affected by three factors: firms' practices, the state of governance in society, and the country's economic and institutional context. As stated by Wahyuni et al. (2024), firm influences include reasons for acting, a drive to create something new, building a good sustainability reputation, and transparency through reporting, all of which fall under accountability frameworks. Following García et al. (2023), Russo, Mariani, and Caragnano (2021), Mertzanis (2023), and Kanamura (2020), research on institutional and environmental governance has highlighted the influence of corporate sustainability factors on the issuance and performance of green bonds. Macroeconomic and institutional factors were highlighted by several authors, including Bhattacharyay (2013), Smaoui et al. (2017), and Tolliver, Keeley, and Managi (2020), who found that openness to global trade, the direction of investment, and the size of economic sectors promoted green bonds. However, interest rate movements and macroeconomic factors had a greater impact than institutional quality.

On the other hand, the quality of institutions plays a similar role in green finance development as it does in general financial development. Studies reveal that there is a joint effort among institutional quality, governance structures, and financial development to achieve the main goals of green finance, mainly the reduction of carbon emissions (Agustin et al., 2025). Good institutions and financial innovation facilitate financial development, thereby improving the delivery of green finance services (Abaidoo & Agyapong, 2022; Alawi et al., 2022; Aman et al., 2023; Khan et al., 2022; Lisbinski & Burnquist, 2024). Combining them enables green finance to play a significant role in achieving sustainable development. Dosso (2023) showed that strong institutions mitigate the downside of relying on natural resources in terms of financial development, thereby helping achieve long-lasting environmental and economic ambitions. Nguyen and Ha

(2021) argue that institutional quality, along with GDP per capita, inflation, and human development, contributes to greater financial inclusion in ASEAN nations. On the other hand, Vo (2024) noted that although institutional quality helps increase financial inclusion in high- and middle-income countries, it does not appear to make a difference in lower-income economies. In addition, Khan et al. (2023) found that better institutions make financial systems work more efficiently, thus strengthening the key position of institutions in helping green finance progress via enhanced infrastructure.

The empirical literature relevant to this study is implied, ranging from determinants of green bonds to the influence of institutional quality on financial development. However, these studies fail to account for how institutional quality could explain green finance growth in emerging economies, as it has in the case of financial development (Abaidoo & Agyapong, 2022; Alawi et al., 2022). The main argument here is that green bonds differ from conventional corporate and government bonds because of their environmental and sustainability focus (Tolliver et al., 2020). The factors that determine them should be considered from the policy point of view. Based on this stance, this study identifies the dearth of institutional quality on green bonds in emerging economies. In the first instance, there is a dearth of empirical studies on the influence of institutional quality on green bonds, especially in emerging economies. This gap exists in understanding the impact of institutional quality on green bond development in emerging economies, as it does for financial growth in other regional countries. This study explores the individual roles of three main institutional qualities—regulatory quality, voice and accountability, and the rule of law—and the influence of their composite value. This approach presents a more comprehensive examination of institutional quality.

Existing empirical studies have highlighted the significance of many factors that can determine green finance, including firm-level factors, environmental factors, and the macroeconomic outlook. While these factors are essential to green finance, they are not devoid of institutional influences in emerging economies. Institutional effects on the green bond market are expected to be as vital as they have been for financial development. However, despite the relevance of institutional factors to economic growth, previous studies have not adequately incorporated them into the green bond market in emerging economies. Moreover, there is a distinct effect of some institutional factors, including regulatory quality, rule of law, and voice and accountability, in emerging economies, which are fundamental concerns in the field of institutional economics.

The associated research question is: What is the effect of institutional quality on green bond development in emerging economies, and how do their individual influences differ from their collective effect? The institutional qualities to be considered in this study are regulatory quality, rule of law, and voice and accountability, as they are the most prominent institutional challenges in emerging economies that bear on the financial market (Abaidoo & Agyapong, 2022). By examining the effects of institutional quality on green finance development in emerging economies, our study makes a significant contribution to knowledge for the following reasons. While many studies have discussed the essential factors for green bonds, such as firm, environmental, and macroeconomic factors (including transparency, innovation, sustainability governance, and economic stability), little is known about how the quality of institutions in emerging economies contributes to green bond growth. In addition, although strong institutions support financial development and inclusion, especially in richer countries, the impact of institutions on green bonds remains unclear. As a result, this study gives a new and detailed understanding by examining the distinct and collective effects of regulatory quality, voice and accountability, and the rule of law on green bond development in emerging economies.

Methods

Data Sources

Our study employs data from 21 emerging economies from 2010 to 2023. The composition of the sampled countries and the chosen period were primarily determined by the availability of relevant data, which the study seeks to investigate. The countries included are Brazil, Thailand, Chile, Türkiye, China, South Africa, Colombia, the Russian Federation, Hungary, Nigeria, India, Indonesia, Vietnam, Malaysia, Mexico, Peru, the Philippines, Poland, Argentina, the United Arab Emirates (UAE), and Mauritius. The databases that serve of the data sources for our study include the World Development Indicators (WDI) and the World Governance Indicators (WGI).

Table 1. Variables and their sources

Variable	Acronym	Description	Source
Green bond	GRB	Fixed income securities that support environmental projects, the ratio of green bonds to GDP ratio	IMF climate dashboard (https://climatedata.imf.org/)
(Bhattacharyay, 2013; Smaoui et al., 2017) Regulatory Quality	REGQ	Ability of government institutions to formulate and implement policies	WDI (https://data.worldbank.org/indicator/RQ.EST)
Rule of Law	RUL	Ability of the legal system to uphold justice for all	WDI (https://data.worldbank.org/indicator/RL.EST)
Voice and accountability	VOACC	Freedom of participation, expression and transparency of processes	WDI (https://data.worldbank.org/indicator/VA.EST)
Institutional quality	INSQ	Principal Component of Regulatory quality, accountability and Rule of Law	PCA of the above
Financial Development	FIND	Credit to the private sector as a percentage of GDP	WDI (https://data.worldbank.org/indicator/FS.AST.PRVT.GD.ZS)
Inflation	IFL	Sustained increase in general prices	WDI (https://data.worldbank.org/indicator/FP.CPI.TOTL.ZG)
Savings	SAV	Income portion not consumed	WDI (https://data.worldbank.org/indicator/NY.GNS.ICTR.ZS)
Investment	INV	Financial resources for fixed or capital assets	WDI (https://data.worldbank.org/indicator/NE.GDI.TOTL.ZS)
Exchange rates	EXR	Official rate of local currency to the US dollar	WDI (https://data.worldbank.org/indicator/PA.NUS.FCRF)
Trade openness	TRADE	Ratio of imports plus exports to GDP	WDI (https://data.worldbank.org/indicator/NE.TRD.GNFS.ZS)

The three principal components of institutional quality — accountability, rule of law, and regulatory quality — explain about 64.5%, 32.4%, and 3.1% of the variation in the data, respectively. Given the significance of each of the institutional quality indicators, all will be included in the derivation of the final composite indices. Estimates of these indices range from approximately -2.5 (lowest) to 2.5 (highest) performances.

All the selected variables are relevant for explaining the growth of green bonds, as reflected in green bond valuation relative to GDP. Table 1 provides definitions of variables and their sources. Given the sensitivity of the green bond market to the macroeconomic environment, there is a need to control for some essential macroeconomic variables (Bhattacharyay, 2013; Smaoui et al., 2017).

The Model

The approach of this research is quantitative, focusing on the description of institutional quality, green finance, and governance indicators. Further, the quantitative analysis of this study applies appropriate econometric methods to examine how institutional quality and governance determine green finance development. Following the tested empirical models of closely related literature

(Ellahi et al., 2021; Khan et al., 2023; Lisbinski & Burnquist, 2024; Muhammed et al., 2024; Tolliver et al., 2020; Vo, 2024), the standard form of the econometric model of this study is:

$$GRB_{it} = \varphi_{it} + \pi_{1,it} INSQ_{it} + \pi_{1,it} INSV_{it} + \pi_{j,it} C_{it} + \varepsilon_{it} \quad (1)$$

Where, $\mathrm{GRB_{it}}$ green bond issuance to GDP, φ_{it} is the model intercept, $\mathrm{INSQ_{it}}$ is the composite index of institutional quality, $\mathrm{INSV_{it}}$ is the vector of institutional quality components, namely regulatory quality (REGQ), rule of law (RUL), and voice and accountability (VOACC). C_{it} represents the control variables made up of Financial Development (FIND), inflation (IFL), savings (SAV), investment (INV), exchange rates (EXR), and trade openness (TRADE). The reason for including financial development, inflation, savings, investment, exchange rates, and trade openness in the green bond development model is that each plays a significant role in shaping green finance. Evidence from Bhattacharyay (2013), Smaoui et al. (2017), and Tolliver et al. (2020) suggests that trade growth, increased foreign investment, and the size of major sectors are major drivers of green bond issuance. However, inflation and currency exchange rate fluctuations tend to discourage it. The progress of financial development strengthens institutions and drives the development of new financial methods that support green finance (Khan et al., 2022; Lisbinski & Burnquist, 2024). In the same way, mobilising savings and investing them helps direct the economy and is fundamental for financing green projects (Dosso, 2023). All of these factors, together, adjust for changes in the economy and the quality of regulatory frameworks that are important to green bonds (Nguyen & Ha, 2021).

The study relied on the panel Fully Modified Ordinary Least Squares (PFMOLS) to estimate long-run panel cointegration relationships between institutional quality and green bond development. The PFMOLS was designed to address the issues of inefficient and biased estimators and inherent endogeneity issues (Özdemir & Kayhan, 2021). The PFMOLS was developed by Phillips and Hansen (1990). It also addresses the issue of serial correlation problems using the Generalised Least Squares (GLS) method. With the benefit of heterogeneous cointegration, Hamit-Haggar (2012) availed that the FMOLS technique is ideal for panel analysis. For a panel FMOLS estimator, the coefficient β of the model in equation 1 was specified by Pedroni (1996) and Khan et al. (2019) to be:

$$\beta^*_{NT} - \beta = \left(\sum_{i=1}^{N} L_{22i}^{-2} \sum_{i=1}^{T} (\chi_{it} - \bar{\chi}_i)^2\right)^{-1} \sum_{i=1}^{N} L_{11i}^{-1} L_{22i}^{-1} \left(\sum_{i=1}^{T} (\chi_{it} - \bar{\chi}_i)\phi^*_{it} + T\hat{\gamma}^*_i\right) \quad (2)$$

Where, $\phi^*_{it} = \phi_{it} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}} \Delta\chi_{it}, \hat{\gamma}^*_i = \hat{\Gamma}_{21i}\hat{\Omega}^0_{21i} - \frac{\hat{L}_{21i}}{\hat{L}_{22i}}\left(\hat{\Gamma}_{22i} - \hat{\Omega}^0_{22i}\right)$ and $\hat{L}_i$ was the lower triangulation of $\hat{\Omega}_i$. The PFMOLS will be conducted on four models of interest, in which we will consider stepwise independent individual effects of institutional quality indicators (regulatory quality, rule of law, voice and accountability, and the composite index of the three coded INSQ indicators).

Results and Discussion

The descriptive statistics in Table 2 provide valuable insights into the patterns and variability of the variables studied: green bonds (GRB), institutional indicators, and macroeconomic variables. From the perspective of the mean, GRB and institutional qualities are at average levels of 0.01% of GDP and 0.05 index, respectively, indicating that green bond activity and institutional qualities are approaching their baseline values. A situation that, on average, depicts the low levels of green bond and institutional quality in emerging economies. Their distributions do not vary much, with standard deviations of not more than 1.18 for INSTQ. Though GRB and institutional qualities tend to show minimal variation, their recorded maximum values of 0.53% and 2.98% indicate that these metrics can reach high outliers.

Among macroeconomic control variables, trade openness has the highest mean value, expressed as of percentage. In contrast, the high values of exchange rates are mainly due to Vietnam, which is struggling with low currency values relative to the panel. The highest levels of financial development and inflation were observed in China and Argentina, respectively.

Table 2. Descriptive analysis

	Obs.	Mean	SD	Max.	Min.
GRB	287	0.01	0.04	0.53	0.00
INSQ	287	0.05	1.18	2.98	-2.06
REGQ	287	0.19	0.57	1.54	-1.02
RUL	287	-0.06	0.57	1.35	-1.18
VOACC	287	-0.08	0.75	1.11	-1.68
FIND	287	66.23	44.06	194.67	0.00
IFL	287	6.51	11.56	133.49	-2.08
SAV	287	26.02	10.15	59.22	8.78
INV	287	24.88	6.93	46.66	12.35
EXR	287	1899.34	5344.68	23787.32	1.50
TRADE	287	78.28	46.29	202.33	16.35

Source: Researchers' computation of data from WDI

Table 3. Correlation analysis

	GRB	INSQ	REGQ	RUL	VOACC	FIND	IFL	SAV	INV	EXR	TRADE
GRB	1										
INSQ	0.21***	1									
REGQ	0.20***	0.93***	1								
RUL	0.19***	0.89***	0.84***	1							
VOACC	0.13**	0.71***	0.48***	0.40***	1						
FIND	0.06	0.16***	0.20***	0.37***	0.20***	1					
IFL	0.03	0.19***	0.30***	0.22***	0.06	0.28***	1				
SAV	0.16***	0.13**	0.02***	0.19***	0.60***	0.27***	0.17***	1			
INV	0.08	0.28***	0.20***	0.05	0.51***	0.30***	0.10	0.67***	1		
EXR	0.04	0.27***	0.26***	0.13***	0.30***	0.02	0.05	0.21***	0.33***	1	
TRADE	0.06	0.34***	0.43***	0.55***	0.19***	0.21***	0.22***	0.40***	0.03	0.21***	1

Note: ***, **, and * indicate significant at 1%, 5% & 10%, respectively

Correlation analysis, besides showing the nature of the association between two variables, is also a tool for detecting multicollinearity and determining the appropriateness of including certain variables. To this end, a correlation of 0.80 is considered high enough to signal multicollinearity problems. The correlation results between some of the variables clearly violates this rule. For instance, there was a powerful positive relationship between INSQ and each of REGQ, RUL and VOACC (Table 3). Given the potential for our results to be compromised, a panel stepwise FMOLS will be most appropriate to avoid modelling them together.

The correlation analysis highlights the vital contributions of institutional quality indicators in fostering green bond development and, in turn, influencing other variables. Table 3 shows that institutional quality indicators exhibit mostly significant positive relationships with green bonds and macroeconomic variables. Stronger institutions ensure the effective implementation of environmental policies, while they offer regulatory oversight and enhanced credibility that will attract green investments (Alawi et al., 2022; Khan et al., 2022; Khan et al., 2023). By reducing risks and signalling policy stability, this institution enhances investor confidence in green bonds and broader sustainability initiatives.

The high correlation between GRB and institutional quality indicators underscores how, together, they provide a firm foundation for green bond development. It reduces transaction costs, ensures regulatory compliance, enhances credibility towards instruments of green bonds, and, therefore, this combination is much more effective.

The unit roots of the variables are checked using three panel tests: the LLC (Levin, Lin, and Chu) test and the IM tests. The rule for deciding based on the test results is that both tests' results must agree for a decision to be made. For instance, GRB, INSTQ, REGQ, RUL, VOACC, FIND, SAV, and TRADE were stationary at levels in the LLC results but not in the IM results, and thus do not provide overwhelming evidence of stationarity at levels. However, they became stationary after the first difference. From our panel stationarity tests in Table 4, it can be observed that all variables become overwhelmingly stationary after first differencing, prompting us to conclude that all series are I(1). Exchange rates (EXR) and inflation (IFL) were unarguably I(I) series because of their nonstationary outcome at levels, except after first differencing. So, our data are all I(I) series, and the chosen model supports this characteristic.

Table 4. Panel stationarity tests

Variables	Levels		First difference		Order of Integration
	LLC	IM	LLC	IM	
GRB	-1.90**	-1.48*	-18.40***	-18.30***	*I*(1)
	(0.0288)	(0.0698)	(0.0000)	(0.0000)	
INSQ	-2.81***	0.75	-11.63***	-8.64***	*I*(1)
	(0.0024)	(0.7726)	(0.0000)	(0.0000)	
REGQ	-1.65**	0.45	-13.69***	-10.20***	*I*(1)
	(0.0499)	(0.6746)	(0.0000)	(0.0000)	
RUL	-2.32***	0.82	-13.67***	-9.94***	*I*(1)
	(0.0100)	(0.7926)	(0.0000)	(0.0000)	
VOACC	-2.28**	0.71	-9.37***	-6.18***	*I*(1)
	(0.0114)	(0.7614)	(0.0000)	(0.0000)	
FIND	-3.94***	-1.26	-8.02***	-4.68***	*I*(1)
	(0.0000)	(0.1046)	(0.0000)	(0.0000)	
IFL	-1.32*	-0.190	-13.92***	-9.59***	*I*(1)
	(0.0933)	(0.4260)	(0.0000)	(0.0000)	
SAV	-3.23***	-1.45*	-12.80***	-10.31***	*I*(1)
	(0.0006)	(0.0735).	(0.0000)	(0.0000)	
INV	-1.160	-1.260	-10.02***	-8.87***	*I*(1)
	(0.1237)	(0.1032)	(0.0000)	(0.0000)	
EXR	0.65	4.94	-10.60***	-6.43***	*I*(1)
	(0.7434)	(1.0000)	(0.0000)	(0.0000)	
TRADE	-3.78***	-1.520	-12.11***	-10.22***	*I*(1)
	(0.0001)	(0.0647)	(0.0000)	(0.0000)	

Note: ***, **, and * indicate significant at 1%, 5% & 10%, respectively

Table 5. Panel cointegration tests

		Model 1	Model 2	Model 3	Model 4
Kao Tests for Cointegration	Stat.	-3.88***	-3.82***	-4.06***	-3.82***
	prob.	(0.0001)	(0.0001)	(0.0000)	(0.0001)

Note: ***, ** & * are significant at 1%, 5% & 10%, respectively

The results from the Kao cointegration tests provide helpful information on whether the variables are in a long-run equilibrium relationship. In Table 5, the Kao statistic is -3.88, -3.82, -4.06, and -3.82, while their respective p-values were never more than 0.0001 for all four models of the study. This strongly indicates a significant cointegration relationship between institutional qualities and green bonds, hence supporting the hypothesis of long-run equilibrium.

In the first instance, the results of the panel FMOLS in Table 6 provide overwhelming evidence that institutional qualities in regulatory quality, rule of law, and voice and accountability are individually significant to green bond development. From these results, the rule of law (RUL) and regulatory quality (REGQ) components of institutional quality significantly supported the long-run growth of green bonds in emerging economies. This finding is an affirmation of the earlier positions of Lisbinski and Burnquist (2024), Alawi et al. (2022), Abaidoo and Agyapong (2022), Khan et al. (2022) and Aman et al. (2023) that innovation and good institutional quality are relevant for promoting the effectiveness of green bonds. Green bond issuance in emerging economies is highly dependent on improved regulation and the rule of law to fund environmental projects.

A study by Mbulawa and Chingoiro (2024) affirms the significance of institutional quality in developing the financial sector and, by extension, the economy as a whole. It emphasises that institutional quality is the long-term foundation for financial progress. Lisbinski and Burnquist

(2024) argue that institutional quality is not just relevant for financial development in low-income countries but also the reason advanced countries have made significant progress in their financial sectors. These studies have shown that institutional quality is essential, thereby underscoring its indispensable role in green bond financial development (Abaidoo & Agyapong, 2022; Alawi et al., 2022; Aman et al., 2023; Khan et al., 2022).

Table 6. Panel fully modified ordinary least squares estimates

	Model 1	Model 2	Model 3	Model 4
FIND	-0.000**	-0.000***	-0.000***	-0.000***
	(0.010)	(0.007)	(0.008)	-0.003
IFL	0.0003**	0.0001	-0.000	0.000
	(0.029)	(0.333)	(0.815)	(0.560)
SAV	-0.001***	-0.0014***	-0.002***	-0.002***
	(0.000)	(0.000)	(0.000)	(0.000)
INV	-0.0001	0.0002	0.0002	0.0001
	(0.785)	(0.425)	(0.608)	(0.637)
EXR	0.0000003	0.000002*	0.000002	0.000002
	(0.817)	(0.056)	(0.140)	(0.096)
TRADE	-0.0003***	-0.0003***	-0.0003***	-0.0003***
	(0.003)	(0.000)	(0.001)	(0.000)
REGQ	0.0105**			
	(0.048)			
RUL		0.0188***		
		(0.000)		
VOACC			-0.0377***	
			(0.000)	
INSTQ				-0.0086***
				(0.002)
Obs.	243	243	243	243
R-Square.	0.28	0.28	0.29	0.28
Adj. R-Square.	0.19	0.20	0.21	0.20

Note: ***, **, and * indicate significant at 1%, 5% & 10%, respectively

The results of model three showed that voice and accountability had a detrimental effect on green bond development in emerging economies. Further, in model four, the results revealed that the composite index of institutional quality also negatively influences green bonds in emerging economies. These findings are in variance with Nguyen and Nguyen and Ha (2021), Khan et al. (2023), and Vo (2024), who found that institutional quality improved financial inclusion and efficiency. Despite the supportive individual roles of regulation and the rule of law, the weight of the detrimental effects of voice and accountability was enough to make the overall effect of institutional quality problematic for green bond development. The uniqueness of our findings is relevant to the recommendation that partial institutional quality will not benefit the development of the green bond market in emerging economies. These findings imply that institutional reforms that are not holistic and comprehensive will be detrimental to green bond issuance. Finally, the effectiveness of institutional quality in promoting the green bond market may need to be complemented by macroeconomic support, such as proper regulation for redirecting savings to green bonds, promoting interment culture, trade openness policies that promote environmental conservation and promoting more financial depth (Bhattacharyay, 2013; Smaoui et al., 2017; Tolliver et al., 2020).

Conclusion

The underdevelopment of the green bond market in emerging economies led the study to consider the influences of institutional quality in explaining green bond development in twenty-one (21) emerging economies, from 2010 to 2023. Given data availability and the nature of the data, the most suitable technique is the panel Fully Modified Ordinary Least Squares (PFMOLS) estimator. The proxy for institutional quality data is regulatory quality, voice and accountability, and the rule of law. The influence of institutional attributes in this study led to three distinct relationships with the green bond. First, efforts in emerging economies that independently prioritise regulation and rule of law will benefit green bond development in those countries. Second, voice and accountability are individually detrimental to the development of green bonds in emerging economies. Finally, the overall composite index of institutional quality is sensitive to the shortcomings of a single indicator, and the detrimental effect of voice and accountability weighs on the overall quality of institutions, thereby harming green bonds.

Based on the study's findings, we offer the following recommendations. It is recommended that emerging economies strengthen their existing institutional structures. Rather than isolated reforms in voice and accountability, enforcement should also be improved in promoting regulatory quality and the rule of law. This is because the detrimental effect of a troubled voice and accountability can weigh on the overall challenge of the entire institutional structure. Encouraging inclusive institutional quality can build legitimacy and green investors' confidence, thereby supporting long-run green finance outcomes.

The main challenge in the study stemmed from the data, which was available only in panel form. Therefore, the best approach was restricted to the panel Fully Modified Ordinary Least Squares estimator. Additionally, the complex interaction between institutional policies and macroeconomic factors shaping green bond markets suggests that other unobserved factors or economic conditions may also influence outcomes, which the current study could not comprehensively address. For further research, it would be valuable to incorporate additional supporting economic variables, such as financial stability, in greater detail, and to proffer a comparative analysis between developed and developing economies when more data are appreciably available. Again, further studies can be carried out on how macroeconomic and external (global) shocks affect the green bond market in emerging and developing economies.

Use of AI tools declaration

The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest

The authors declare no conflicts of interest.

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Asymmetric impacts of exchange rate and petroleum pump price on economic welfare in Nigeria

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Abstract

Purpose — This study examines the asymmetric effects of exchange rate fluctuations and petroleum pump prices on economic welfare in Nigeria. While previous research examined these shocks in isolation, this study jointly evaluates their short-run and long-run effects, thereby addressing a key gap in the literature.

Methods — The study employs the Nonlinear Autoregressive Distributed Lag (NARDL) model to analyse time-series data from 1970 to 2023.

Findings — Exchange rate depreciations and petroleum price increases have larger and lasting welfare losses than the short-run benefits of appreciations and price declines. In the long run, these shocks can be turned into potential welfare benefits through structural changes and redistribution of the budget. Inflation, unemployment, subsidies and international oil prices further mediate outcomes.

Implications — Policymakers should strike a balance between short-term household protection and longer-term structural changes. With the complete removal of petroleum subsidies in Nigeria in May 2023, the focus should shift to special transfers, social security, and compensation to mitigate welfare losses. Exchange rate stability, fiscal discipline and diversification are equally essential for enhancing long-term welfare.

Originality — This study advances understanding of welfare by concurrently examining the asymmetries of exchange rates and petroleum pump prices, thereby moving beyond the single-shock approach.

Keywords — Exchange rate, oil price, economic welfare, subsidies, unemployment

Introduction

There is no gainsaying that the Nigerian economy is heavily influenced by the petroleum sector, which has served as a primary source of government revenue and a critical input to various economic activities since the 1970s. As Africa's largest oil producer and one of the world's top oil producers, Nigeria is profoundly affected by changes in global oil prices (Adams & Olamide Bello, 2022; Ojumu & Osho, 2023). Oil exports account for a significant share of the country's gross domestic product (GDP) and foreign exchange earnings (Okon et al., 2020), making the economy highly susceptible to fluctuations in global oil prices and exchange rates. This dependence has created a situation of which the nation's well-being is closely tied to the health of the petroleum sector, making it a vital area of focus for economic stability and welfare.

In recent years, exchange rate volatility has been a persistent issue in Nigeria, driven by various factors, including global oil market dynamics, domestic economic policies, and fluctuations

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in foreign exchange reserves. Based on data from the Central Bank of Nigeria (CBN), the naira has experienced significant depreciation, with the average exchange rate rising from approximately N197 per USD in 2015 to about N365 per USD in 2019 and N460 per USD in mid-2023, before depreciating drastically to over N1000 per USD after the adoption of floating exchange rate system in the third quarter of 2023. Such depreciation has a direct impact on the cost of imported goods and services, including petroleum products, since Nigeria imports a substantial share of its fuel. Exchange rate movements can cause significant shifts in the price of petroleum, affecting consumers and businesses alike.

Petroleum products play a fundamental role in the Nigerian economy, powering transportation, manufacturing, and various other sectors. The Nigerian National Petroleum Corporation (NNPC) reported that in 2020, Nigeria imported about 20 billion litres of petrol. As a result of this heavy reliance on imported fuel, the cost of petroleum prices is susceptible to exchange rate fluctuations. For instance, when the naira weakens, the cost of importing fuel rises, leading to higher domestic pump prices. A depreciation of 10% in the naira, for example, could lead to a corresponding increase in petrol import costs by roughly the same percentage, or more, ultimately driving up consumer prices. As petroleum products become more expensive, transportation and production costs rise, contributing to overall inflation and eroding the purchasing power of Nigerian consumers.

The impact of fuel price hikes is particularly severe for low-income households, who spend a substantial portion of their income on necessities, including transportation. According to the CBN, annual inflation rates in Nigeria increased from 9% in 2015 to 15.7% in the following year, and to about 22% in 2023, with a notable portion of the rise attributed to higher fuel prices. This inflationary trend erodes real incomes, exacerbates poverty, and widens the gap between wealthy and vulnerable segments of the population. The World Bank, for instance, estimates that as of 2019, 40.1% of Nigerians were living below the poverty line, surviving on less than $1.90 per day. Rising fuel prices further strain these households by limiting their access to essential goods and services and diminishing their overall quality of life.

In addition to its impact on the cost of living, exchange rate, and petroleum price volatility can also destabilise the broader economy. Rising input costs can discourage investment and reduce business profitability, with adverse effects on employment levels. Nigerian unemployment rate, as reported by the NBS, surged to 33.3% in the fourth quarter of 2020 and to 35% in 2023, underscoring the country's broader economic challenges. A high unemployment rate, coupled with increasing inflation and fluctuating fuel prices, poses a substantial risk to economic welfare and social stability.

Meanwhile, the Nigerian government has attempted to mitigate these issues through various policy interventions, including subsidies and price controls. In 2020, for example, Nigeria allocated over N2 trillion to fuel subsidies, which increased to about N3 trillion in 2023, aiming to shield consumers from the full impact of rising petroleum prices. While subsidies may provide temporary relief, they also place a heavy burden on the national budget and may not address the basic factors driving price volatility (Jayne & Rashid, 2013). Furthermore, the Nigerian government's periodic adjustments to fuel prices often lead to public dissatisfaction and uncertainty, eroding consumer confidence and complicating long-term economic planning. Hence, the current administration in Nigeria announced the total removal of petroleum subsidies on May 29, 2023, citing a paucity of funds and the need of diversify the economy as the bases.

Despite these contributions, the literature remains fragmented. Most studies investigate oil prices, exchange rates, or subsidies separately, focusing on outcomes such as trade balance (Areghan et al., 2022), fiscal costs (Aruofor & Ogbeide, 2023; Okwanya et al., 2015; Orlu, 2018), and inflation (Adams et al., 2024; Maku et al., 2018; Sakanko et al., 2021). Some concentrate on fluctuations in exchange rates without explicitly linking them with welfare outcomes (Aimer & Lusta, 2021). However, little empirical attention has been given to the combined and asymmetric effects of exchange rate fluctuations and petroleum pump price changes on economic welfare in an oil-reliant country such as Nigeria. Moreover, the distinction between short-run and long-run effects is commonly neglected. This paper fills existing gaps by applying the NARDL model to analyse the asymmetric effects of exchange rate variation and petroleum pump price on economic welfare in Nigeria between 1970 and 2023. This method generates new perspectives on the interaction of essential macroeconomic factors that influence welfare outcomes and provides empirical direction for policy reforms.

The Mundell-Fleming model extends the IS-LM framework to open economies, illustrating how exchange rates, interest rates and output interact under different exchange rate regimes (Bofinger et al., 2009). It suggests that under a fixed exchange rate, fiscal policy is more effective while monetary policy is constrained (Serrano & Summa, 2015). Conversely, under a flexible exchange rate, monetary policy dominates, influencing capital flows, interest rates, and trade balances (Darius, 2010). Nigeria's alternating exchange rate regimes affect imports, inflation, and economic growth, predominantly by influencing petroleum prices and overall welfare. Similarly, the purchasing power parity theory posits that exchange rates should adjust to equalise the price levels of goods across countries (Nakorji et al., 2021). However, Nigerian exchange rate misalignments and market imperfections create price disparities in petroleum products, affecting the cost of living and welfare.

Also, the exchange rate pass-through theory investigates the extent to which exchange rate changes influence domestic prices (Aron et al., 2014; Zubair et al., 2013). High pass-through rates indicate that exchange rate fluctuations significantly affect inflation, whereas low pass-through rates suggest a moderate impact. In Nigeria, a high pass-through rate, due to import reliance and petroleum price sensitivity, contributes to economic instability. In addition, welfare economics provides a framework for assessing the impact of monetary policies on individual well-being (Engelbrecht, 2009). Fluctuations in exchange rates and petroleum prices affect real income, consumption patterns, and inequality. Changes in petroleum pump prices alter transportation costs and household expenditures, influencing welfare. Meanwhile, the resource curse theory suggests that countries rich in natural resources, such as oil, may experience slower economic growth due to challenges including rent-seeking behaviour, corruption, and financial volatility (Henri, 2019; Mlambo, 2022). Nigeria exemplifies the resource curse as its dependence on oil has led to economic instability, exchange rate volatility and obstacles to sustainable development.

Areghan et al. (2022) show that exchange rate stability is critical for maximising the contributions of the oil sector to the balance of payments. Also, Abere (2023) reaches the same conclusion that naira undervaluation enhances balance of trade while overvaluation worsens performance, and shows that foreign trade is vulnerable to fluctuations in the exchange rate. Adagunodo (2013) maintains that whereas the removal of PMS subsidies could yield significant fiscal savings, the welfare justification of kerosene subsidies is far-fetched. Using an error-correction model, Okwanya et al. (2015) and Orlu (2018) show that PMS price increases had decreased growth and created short-term inflationary pressures. Aruofor & Ogbeide (2023) cautiously note that the removal of subsidies might reduce production, suppress demand, and worsen the risk of inflation. Similarly, Amaechi (2024) notes that the withdrawal of fuel subsidies has an adverse though insignificant effect on the standard of living in Nigeria. These studies portray trade-offs between fiscal sustainability and welfare protection.

However, Sakanko et al. (2021) demonstrate that a petroleum price shock creates lasting inflationary pressure, while Maku et al. (2018) show that increases in pump prices and inflation have a net detrimental effect on welfare by disrupting the economy. Adams et al. (2024) state that an increase in the price of petrol significantly increases the consumer price index in the short run, but the effect decreases over time. Similarly, Orekoya et al. (2024) expand on this evidence by showing asymmetric impacts of domestic prices of oil price shocks on domestic prices. In the short term, positive shocks have a greater effect than their negative counterparts, but in the long term, adverse shocks have a greater impact. These suggest that a long-term price stabilisation could be achieved by removing subsidies and investing in refining capacity.

Olayungbo (2019) argues that oil revenue has significantly contributed to economic growth, while also creating structural inadequacies, such as poor human capital and restrictive trade policies. Okon et al. (2020) also report a positive effect of oil revenue on GDP, but caution against overdependence on oil and suggest the diversification of the economy with agriculture. Adams & Olamide Bello (2022) emphasise the importance of downstream production and private sector involvement in sustaining growth. Ojumu & Osho (2023) shift the focus to declining U.S. demand for Nigerian oil and its economic implications, particularly in terms of trade surpluses and the depletion of foreign reserves, thereby exposing Nigeria's vulnerability to global oil market fluctuations. Also, Somoye et al. (2024) indicate that the use of fossil fuels and renewable energy sources helps increase life expectancy in both the short and long term. As Wasurum (2025) shows, oscillations in oil prices and exchange rates support economic growth in Nigeria, but with uneven impacts through the channels of life expectancy and literacy.

Aimer & Lusta (2021) demonstrates that oil price shocks have a more pronounced impact on the exchange rate than U.S. economic policy uncertainties, leading to adverse currency effects. Ojeyinka & Aliemhe (2023) demonstrate that demand-driven oil price shocks positively affect aggregate and sectoral stock returns, whereas supply shocks are significant only for the oil and gas sector. El Yadmani (2025) concludes that fluctuations in oil prices negatively impact Morocco due to its high dependence on imported fossil fuels, therefore, compromising its long-term growth and household welfare.

Methods

Model Specification

To determine the effects of exchange rate fluctuations and petroleum pump price on economic welfare, this study adapts the models by Ighosewe et al. (2021) and Areghan et al. (2022). Thus,

$$ECOW_t = f(EXR_t, PPP_t) \tag{1}$$

where $ECOW_t$, EXR_t and PPP_t represent the economic welfare index, the exchange rate and the petroleum pump price, respectively, at time t. While EXR and PPP are exogenous variables, the model also includes some control variables.

$$ECOW_t = f(EXR_t, PPP_t, INFR_t, UEMPR_t, SUB_t, COIP_t) \tag{2}$$

where $INFR_t$, $UEMPR_t$, SUB_t and $COILP_t$ represent the inflation rate, unemployment rate, government expenditure on subsidies and crude oil price, respectively, at time t. The model is explicitly expressed as:

$$ECOW_t = \alpha_0 + \alpha_1 EXR_t + \alpha_2 PPP_t + \alpha_3 INFR_t + \alpha_4 UEMPR_t + \alpha_5 SUB_t + \alpha_6 COILP_t + \mu_t \tag{3}$$

where all variables are as earlier stated, μ_t is the stochastic term and $\alpha_i (i = 0,1,2,3,\ldots,6)$ are the estimated parameters. The data measurement and sources are presented in Table 1.

Table 1. Data Measurement and Sources

Variable	Measurement	Source
Economic welfare (ECOW)	Composite index based on poverty headcount, Gini coefficient, and real per capita income (through PCA)	WDI; Naveed & Gordon (2024); Smiech et al. (2025)
Exchange rate (EXR)	Nominal annual average exchange rate of the naira against the US dollar	CBN
Petroleum pump price (PPP)	Annual mean retail pump price of petrol (litre)	NNPC
Inflation rate (INFR)	Annual percentage change in consumer price index	CBN
Unemployment rate (UNEMPR)	Proportion of the labour force without jobs and in pursuit of employment opportunities	CBN
Government expenditure on subsidies (SUB)	Annual government spending on petroleum subsidies	CBN
International oil prices (COILP)	Annual average Brent crude oil price (USD/barrel)	OPEC

PCA: Principal Component Analysis, CBN: National Bureau of Statistics, Nigerian National Petroleum Corporation (NNPC), OPEC: Organisation of Petroleum Exporting Countries, WDI: World Development Indicators

Estimation Technique

To determine the asymmetric effects of exchange rate fluctuations and petroleum pump prices on economic well-being, this study employs the Nonlinear Autoregressive Distributed Lag (NARDL) technique developed by Shin et al. (2014). This paradigm decomposes the explanatory variables into positive and negative changes, thereby quantifying possible asymmetry in both long- and short-term dynamics.

Decomposition of Shocks

Following Shin et al. (2014), the exchange rate (EXR) and petroleum pump price (PPP) are decomposed into cumulative sums of positive and negative changes.

$$EXR_t^+ = \sum_{j=1}^{t} \Delta EXR_j^+ = \sum_{j=1}^{t} max(\Delta EXR_j, 0) \tag{4}$$

$$EXR_t^- = \sum_{j=1}^{t} \Delta EXR_j^- = \sum_{j=1}^{t} min(\Delta EXR_j, 0) \tag{5}$$

$$PPP_t^+ = \sum_{j=1}^{t} \Delta PPP_j^+ = \sum_{j=1}^{t} max(\Delta PPP_j, 0) \tag{6}$$

$$PPP_t^- = \sum_{j=1}^{t} \Delta PPP_j^- = \sum_{j=1}^{t} max(\Delta PPP_j, 0) \tag{7}$$

Here, EXR_t^+ and PPP_t^+ are cumulative increases while EXR_t^- and PPP_t^- are cumulative decreases. The asymmetric effects are tested through these decompositions (Baharumshah et al., 2017; Widarjono et al., 2023).

Pre-estimation Tests

ADF and PP unit root tests are conducted to determine the order of integration of variables. NARDL accommodates $I(0)$ and $I(1)$ variables but excludes $I(2)$ to ensure the validity of its bounds test (Pesaran et al., 2001). The Akaike Information Criterion determines the optimal lag length (AIC) for efficiency (Baharumshah et al., 2017).

Cointegration and Diagnostic Tests

The existence of a long-run relationship is tested using the bounds test developed by Pesaran et al. (2001). The null hypothesis of no cointegration is rejected, implying a long-run relationship. Diagnostic tests – Breusch-Godfrey (serial correlation), Breusch-Pagan-Godfrey (heteroskedasticity), Jarque-Bera (normality) and Ramsey-RESET (functional form) – are conducted to validate model adequacy.

Asymmetry tests

The Wald test is utilised to verify whether positive and negative changes exert symmetric or asymmetric effects. For instance, $\alpha_1 = \alpha_2$ measures whether exchange rate appreciation and depreciation have equal impacts while $\alpha_3 = \alpha_4$ testing the symmetry of petroleum pump price changes.

Error Correction Representation

After establishing cointegration, the NARDL model is expressed in error correction form:

$$\Delta ECOW_t = \phi ECT_{t-1} + \sum_{i=1}^{p} \beta_i \Delta ECOW_{t-1} + \sum_{j=0}^{q_1} \left(\tau_j^+ \Delta EXR_{t-j}^+ + \tau_j^- \Delta EXR_{t-j}^-\right) + \sum_{j=0}^{q_2} \left(\delta_j^+ \Delta PPP_{t-j}^+ + \delta_j^- \Delta PPP_{t-j}^-\right) + \sum_{k=0}^{q_3} \lambda_k \Delta OV_{t-k} + \mu_t \tag{8}$$

where ECT_{t-1} is the error correction term, and its coefficient ϕ must be negative and significant to confirm convergence to the long-run equilibrium. The index k denotes the lag order of the control variables (OV).

Long-run Model

The NARDL long-term asymmetric relationship is derived as:

$$ECOW_t = \theta_0 + \theta_1^+ EXR_t^+ + \theta_1^- EXR_t^- + \theta_2^+ PPP_t^+ + \theta_2^- PPP_t^- + \sum_{s=1}^{S} \lambda_s OV_{s,t} + \quad (9)$$

where θ_1^+ and θ_1^- capture the long-run effects of exchange rate appreciation and depreciation, θ_2^+ and θ_2^- capture the long-run effects of petroleum pump price increases and decreases, $OV_{s,t}$ denotes the control variables at time t, λ_s represents long-run coefficients of control variables, and S is the total number of control variables.

Results and Discussion

Unit Root Tests

Since most time series data are non-stationary at the level, the literature suggests that stationarity tests should be conducted. Hence, the variables in the estimated model are tested for the presence or absence of a unit root. As shown in Table 2, both the Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests indicate that government expenditure on subsidies (SUB) is stationary at level $I(0)$, while the other variables become stationary after first differencing $I(1)$. Meanwhile, the $NARDL$ technique is suitable for handling a model comprising a mixture of $I(0)I(1)$ variables (Pesaran et al., 2001).

Table 2. Unit Root Test

Variable	ADF Level	ADF First Difference	PP Level	PP First Difference	Remark
ECOW	-1.627	-7.694***	-1.671	-7.694***	I(1)
EXR	0.336	-5.497***	0.935	-5.274***	I(1)
PPP	-2.169	-6.390***	-2.342	-6.384***	I(1)
INF	-3.757**	-	-2.976	-5.694***	I(1)
UNEMPR	-6.290***	-2.076	-2.076	-7.767***	I(1)
SUB	-4.878***	-	-5.208***	—	I(0)
COILP	-2.563	-6.536***	-2.709	-6.593***	I(1)

*** and ** indicate 1% dan 5% levels of significance, respectively

Optimum Lag Selection

The VAR lag-order selection criteria in Table 3 indicate that the Akaike Information Criterion (AIC), Hannan-Quinn Criterion (HQ), Final Prediction Error (FPE), and Sequential Likelihood Ratio (LR)all indicate four delays. At the same time, the Schwarz Criterion (SC) suggests only one lag. This study uses four lags, as indicated by most criteria, especially AIC. Furthermore, the model selection summary $(4,4,4,4,4,3,4,4,4)$ of Figure 1 shows that most variables enter the model with four lags, except subsidy, which has three lags. This structure captures both current and past effects, allowing us $NARDL$to model dynamic and asymmetric responses of economic welfare $(ECOW)$ to changes in the explanatory variables.

Table 3. Lag Order Selection Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-1272.134	NA	3.92e+13	51.165	51.433	51.267
1	-926.827	580.115	2.85e+08	39.313	41.455*	40.129
2	-861.108	92.007	1.64e+08	38.644	42.660	40.173
3	-750.897	123.436	1.94e+07	36.195	42.085	38.438
4	-655.612	80.039*	6.16e+06*	34.344*	42.107	37.301*

* Indicate lag order selected by criterion

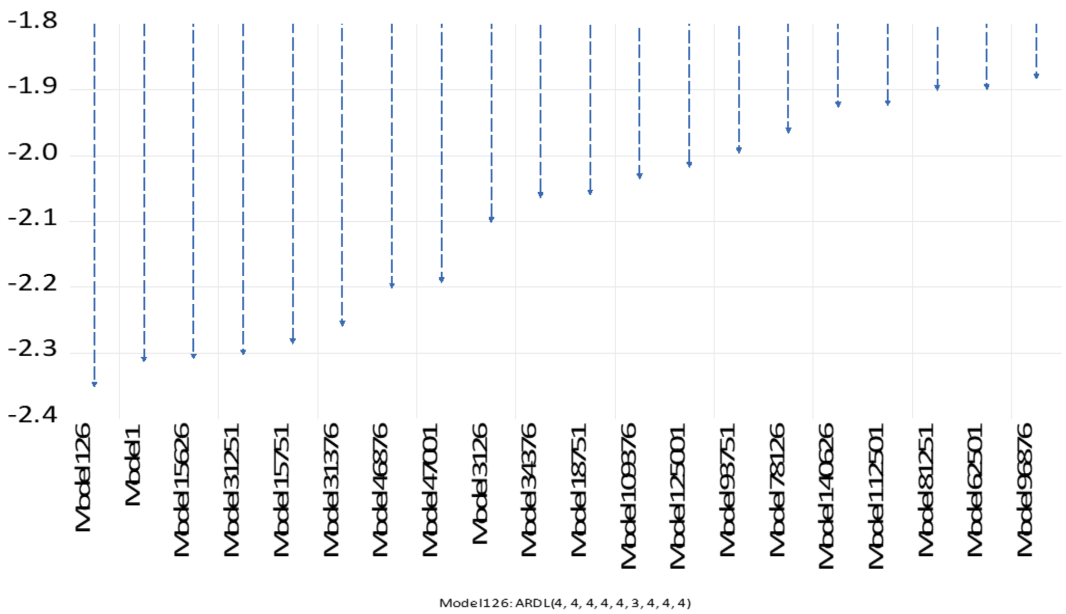


Figure 1. Akaike Information Criterio (top 20 models)

Diagnostic Tests

This study employs several diagnostic tests to guarantee the reliability of its findings (Table 4). The multiple coefficient of determination $(R - squared)$and the adjusted multiple coefficient of determination$(Adjusted\ R - squared)$ help assess the fitness of the estimated $NARDL$model, as $Adjusted\ R - squared$shown by the fact that about 98% of the overall variation in economic welfare is jointly captured by the exogenous variables. The non-significance of $Jarque - Bera$ statistic implies that the variables are normally distributed. Also, the insignificance of $F - statisticSerial\ Correlation$ the test suggests the absence of autocorrelation, indicating that the residuals of the estimated model are uncorrelated. Similarly, the non-significance of $F - statistic$ in the $Heteroskedasticity$ test indicates that the residuals are homoscedastic, implying that they do not differ across all values of the exogenous variables.

Table 4. Diagnostic Test

Test	Statistic	(p-value)
R-squared	0.995	-
Adjusted R-squared	0.984	-
Jarque-Bera	1.051	(0.591)
Breusch-Godfrey Serial Correlation	7.751	(0.065)
Breusch-Pagan-Godfrey Heteroskedasticity	0.604	(0.833)

Cointegration Test

The $F - bounds$statistic $(F = 12.277)$ is substantially above the upper threshold $1\%(3.77)$ (Table 5). Since the computed$F - statistic$ exceeds all critical values $I(0)$ and $I(1)$, the null hypothesis of no level relationship is strongly rejected. This confirms the existence of a co-integrating relationship over time.

Table 5. NARDL F-Bounds Test

Test Statistic	Value	Sig.	I(0)	I(1)
F-statistic	12.277	10%	1.85	2.85
k	8	5%	2.11	3.15
		2.5%	2.33	3.42
		1%	2.62	3.77

Asymmetric Tests

This study investigates the asymmetric effects of the exchange rate and the petroleum pump price on economic welfare using the NARDL model. The $Wald\ test$ results in Table 6 provide significant evidence of asymmetry in both the short- and long-run. The null hypotheses of symmetry are consistently rejected, indicating that positive and negative shocks do not exert equal effects on economic welfare. In the short run, the $Wald\ test$ results show the existence of statistically significant asymmetry in the effects of exchange rate shocks $(t = -3.61, \rho < 0.05; F = 13.02, \rho < 0.05; \chi^2 = 13.02, \rho < 0.01)$. The restriction estimate of -0.378 $(SE = 0.105)$ shows that the cumulative effect of depreciations exceeds that of appreciations in magnitude. In the long run, asymmetry persists, as the $Wald\ test$ data consistently reject the null hypothesis of symmetry $(t = -4.22, \rho < 0.01; F = 17.82, \rho < 0.01; \chi^2 = 17.82, \rho < 0.01)$. The restriction estimate of -0.318 $(SE = 0.075)$ suggests that the negative effects of depreciations on welfare are noticeably higher than the potential benefits of appreciation. This outcome aligns with the findings by Areghan et al. (2022), which emphasises the importance of currency stability for sustaining external balance and welfare.

Table 6. Asymmetric Test (Wald Test)

NARDL Shortrun Asymmetry			
Asymmetric Variable		EXR	PPP
Test Statistic	df	Value (Probability)	Value (Probability)
t-statistic	5	-3.608 (0.015)	-5.350 (0.003)
F-statistic	(1,5)	13.015 (0.015)	28.618 (0.003)
Chi-square	1	13.015 (0.000)	28.618 (0.000)
Normalised Restriction [Std.Error]			
NARDL Longrun Asymmetry			
Asymmetric Variable		EXR	PPP
Test Statistic	df	Value (Probability)	Value (Probability)
t-statistic	5	-4.222 (0.008)	-6.048 (0.002)
F-statistic	(1,5)	17,824 (0.008)	36.576 (0.002)
Chi-square	1	17,824 (0.000)	36.576 (0.000)
Normalised Restriction [Std.Error]		-0.318 (0.075)	-0.585(0.097)

***, **, and * indicate 1%, 5%, and 10% levels of significance respectively

Similarly, the results show considerable evidence of asymmetry in the short-run effects of changes in petroleum pump prices. The $Wald\ test$ data substantially support the rejection of the null hypothesis of symmetry $(t = -5.35, \rho < 0.01; F = 28.62, \rho < 0.01; \chi^2 = 28.62, \rho < 0.01)$. The restriction estimate of -0.850 $(SE = 0.159)$ indicates that price increases have a more substantial negative influence than the positive effect of price decreases. This is consistent with empirical evidence that pump price increases have immediate inflationary and welfare-reducing effects while price reductions provide only transient relief (Aruofor & Ogbeide, 2023; Orlu, 2018). The null hypothesis of symmetry is unequivocally rejected in the long run $(t = -6.05, \rho < 0.01; F = 36.58, \rho < 0.01; \chi^2 = 36.58, \rho < 0.01)$. The restriction estimate of -0.585 $(SE = 0.097)$ confirms that pump price increases exhibit greater persistent and detrimental impacts on economic welfare than decreases. These finding complements Sakanko et al. (2021) and Maku et al. (2018), who emphasise that pump price shocks induce lasting inflationary pressures, while subsidy-driven price reductions offer only limited and uneven welfare benefits.

The findings are consistent in suggesting that both exchange rate and petroleum pump price shocks are asymmetric in both the short-run and long-run. The negative restriction estimates across all specifications indicate that the net adverse impacts of negative shocks – exchange rate depreciation and fuel price increases – are much larger and more persistent than the potential benefits of appreciations or price reductions. Thus, exchange rate depreciation worsens welfare more persistently than the modest benefits from appreciation. At the same time, petroleum price hikes impose immediate welfare losses that outweigh the temporary relief provided by price cuts. These results underscore the need to stabilise the exchange rate in response to ongoing depreciation pressures and to design fuel price reforms alongside compensatory social protection programs. By acknowledging asymmetry, policymakers can better anticipate the disproportionate costs of adverse shocks and avoid the erroneous assumption of symmetric adjustment in macroeconomic management.

Short-run Dynamics

The short-run result in Table 7 shows strong asymmetric effects of exchange rate fluctuations on economic welfare. Positive shocks to the exchange rate (naira depreciation) consistently diminish welfare, while adverse shocks (appreciation) yield mixed results: initially beneficial but, with lags, thereafter detrimental. This result corroborates the findings of studies by Areghan et al. (2022) and Abere (2023), which emphasises the importance of exchange rate stability and undervaluation for trade performance. The results show that prolonged depreciation decreases welfare by increasing import costs and reducing purchasing power, whereas appreciation may only provide temporary relief before adjustment costs materialise.

Table 7. NARDL Error Cointegration Regression

Variable	Coefficient	Std. Error	t-Statistic	Prob.
D(ECOW(-1))	0.114*	0.055	2.074	0.093
D(ECOW(-2))	0.193**	0.067	2.903	0.034
D(ECOW(-3))	-0.307***	0.052	-5.873	0.002
D(EXR_POS)	-0.003	0.001	1.942	0.199
D(EXR_POS(-1))	-0.048***	0.003	-14.525	0.000
D(EXR_POS(-2))	-0.067***	0.004	-16.419	0.000
D(EXR_POS(-3))	-0.029***	0.003	-11.019	0.000
D(EXR_NEG)	0.316**	0.031	10.338	0.000
D(EXR_NEG(-1))	-0.008	0.026	-0.314	0.766
D(EXR_NEG(-2))	-0.158***	0.030	-5.272	0.003
D(EXR_NEG(-3))	-0.164**	0.043	-3.772	0.013
D(PPP_POS)	0.0003	0.001	0.303	0.774
D(PPP_POS(-1))	-0.185***	0.008	-22.075	0.000
D(PPP_POS(-2))	-0.130***	0.009	-14.287	0.000
D(PPP_POS(-3))	-0.066***	0.005	-14.519	0.000
D(PPP_NEG)	0.585***	0.026	22.311	0.000
D(PPP_NEG(-1))	-0.313***	0.017	-18.249	0.000
D(PPP_NEG(-2))	-0.598***	0.031	-19.351	0.000
D(PPP_NEG(-3))	-0.566***	0.038	-14.825	0.000
D(INFR)	-0.003*	0.001	-2.482	0.056
D(INFR(-1))	0.013***	0.001	9.030	0.000
D(INFR(-2))	0.013***	0.001	10.999	0.000
D(UNEMPR)	-0.347***	0.015	-22.503	0.000
D(UNEMPR(-1))	0.123***	0.014	9.004	0.000
D(UNEMPR(-2))	0.358***	0.020	17.700	0.000
D(UNEMPR(-3))	0.195***	0.010	18.616	0.000
D(LNSUB)	1.204***	0.098	12.244	0.000
D(LNSUB(-1))	1.255***	0.075	16.780	0.000
D(LNSUB(-2))	1.448***	0.101	14.294	0.000
D(LNSUB(-3))	0.736***	0.099	7.474	0.001
D(COILP)	-0.020***	0.003	-6.773	0.001
D(COILP(-1))	0.003	0.003	0.965	0.379
D(COILP(-2))	-0.020***	0.003	-7.940	0.001
D(COILP(-3))	-0.019***	0.002	-7.858	0.001
CointEq(-1)*	-1.313***	0.071	-18.541	0.000

***, **, and * indicate 1%, 5%, and 10% levels of significance respectively

Similarly, petroleum pump price adjustments exert significant asymmetric short-run effects. Increases in pump prices reduce welfare through adverse, persistent lagged effects, while decreases yield immediate welfare gains that quickly revert. This reinforces findings of Okwanya et al. (2015) who assert that increases in petroleum prices reduce growth and trigger inflationary pressures, alongside Aruofor & Ogbeide (2023) who note that the withdrawal of subsidies may reduce consumption and further worsen welfare outcomes. The mixed results on adverse price shocks resonate with Amaechi (2024), who found that subsidy removal has an insignificant effect on living standards, highlighting the transitory nature of welfare gains from fuel price reductions.

Inflation and unemployment are also projected to have noticeable impacts. Inflation reduces welfare contemporaneously, but its lagged positive effects reflect delayed nominal adjustments, confirming the results of Adams et al. (2024) and Sakanko et al. (2021), which indicates that petroleum-induced price shocks initially worsen consumer welfare before gradually subsiding. Unemployment exerts negative immediate effects but becomes positive with lags, reflecting structural labour-market adjustments in which informal employment offsets shocks. This corresponds with Maku et al. (2018), who found that inflation and employment-related mechanisms interact with petroleum prices to disrupt short-term welfare. Government expenditure on subsidies demonstrates a robust, consistent positive impact, affirming its protective function in sustaining household wellbeing. This concurs with Aruofor & Ogbeide (2023), who draw attention to the issue of welfare trade-offs connected with subsidy reform. Conversely, an increase in crude oil prices diminishes welfare both immediately and over time, aligning with the findings of Ojumu & Osho (2023) and El Yadmani (2025) that global oil dynamics affect family welfare. The error-correction term is substantial, negative, and significant, indicating rapid adjustment towards the long-run equilibrium. This supports Wasurum's (2025) claim that oil price fluctuations and exchange rates affect welfare in Nigeria through long-run channels.

Long-run Result

The long-run results in Table 8 show that both the exchange rate and the petroleum pump price have a significant asymmetric impact on economic welfare. The effect of positive exchange rate shocks (EXR_POS), denoting depreciation, is positive in the long run; however, the impact of adverse shocks (EXR_NEG), signifying appreciation, is even stronger. This finding reinforces trade-oriented literature such as Areghan et al. (2022) and Abere (2023), which emphasises the importance of exchange rate fluctuations in achieving external equilibrium. Although depreciation may increase the competitiveness of the tradeable sector and indirectly boost welfare, the greater welfare gains from appreciation suggest that stability and reduced import costs yield longer-lasting effects, consistent with the sensitivity of welfare to currency valuation.

Table 8. NARDL Long-run Estimate

Variable	Coefficient	Std. Error	t-Statistic	Prob.
EXR_POS	0.057**	0.018	3.110	0.026
EXR_NEG	0.345**	0.097	3.538	0.017
PPP_POS	0.097***	0.023	4.216	0.008
PPP_NEG	0.744***	0.182	4.083	0.010
INFR	-0.017***	0.003	5.772	0.002
UNEMPR	-0.409**	0.111	-3.691	0.014
LNSUB	1.020***	0.137	7.434	0.001
COILP	-0.026***	0.008	-3.439	0.019
c	-2.358***	0.192	-12.307	0.000

***, **, and * indicate 1%, 5%, and 10% levels of significance respectively

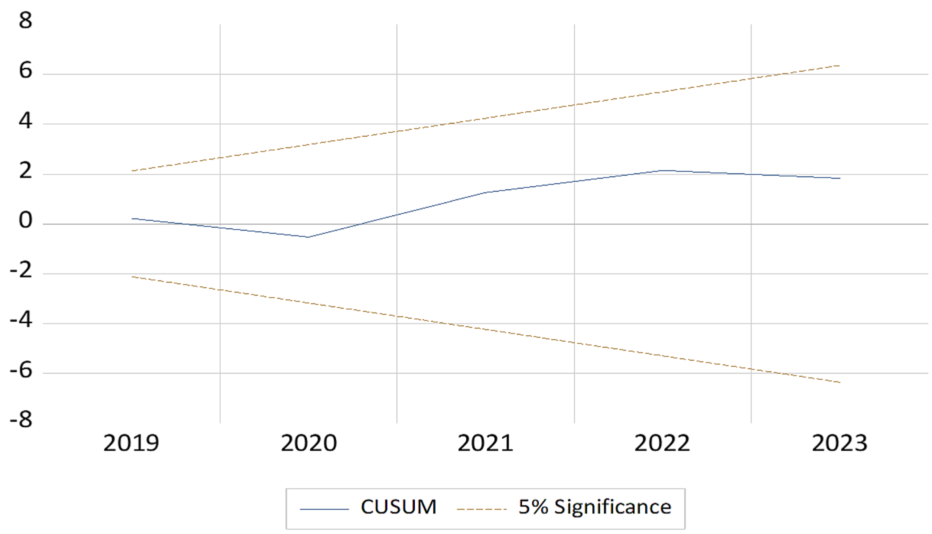
The effects of petroleum pump prices are asymmetrical. Positive shocks (PPP_POS), corresponding to price increases, contribute significantly to better welfare in the long term, while negative shocks (PPP_NEG), denoting price declines, have a greater beneficial effect. This result aligns with Orlu (2018) and Okwanya et al. (2015), who also found that price increases cause short-term distortions. However, it differs from the results of Aruofor & Ogbeide (2023) and Amaechi (2024), who emphasise the adverse welfare effects of subsidy removal. The finding implies that, in the long term, higher pump prices may enhance fiscal capacity and reallocate investment, while price reductions may temporarily alleviate household expenditure burdens. These suggest that welfare responses to petroleum prices are complex and context-dependent.

Controlling variables that significantly reduce welfare include inflation and unemployment, consistent with the results of Sakanko et al. (2021), who focuses on the inflationary transmission of petroleum price shocks, and Maku et al. (2018), who shows that both inflation and unemployment adversely affect welfare. Conversely, government subsidies demonstrate significant benefits, emphasising their stabilising role, as noted by Aruofor & Ogbeide (2023). Crude oil prices adversely affect welfare.

Unlike the short-run results, these long-run results show more stable and positive welfare responses to exchange rate and petroleum price changes. Both depreciation and appreciation positively influence welfare in the long run, with the latter being more significant. Similarly, while initial increases in petroleum prices negatively impact welfare, they are associated with improved long-run performance, possibly because increased fiscal capacity enables more public investment. However, price declines always have a larger impact on welfare. The findings show that the sustainability of long-term welfare depends on effective policy management, particularly in subsidy removal and exchange rate adjustment. These policies may impose short-run costs but can be more sustainable in the long run if accompanied by compensatory measures such as targeted transfers or productive investments. The distinction between short-term and long-term outcomes pinpoints the trade-off between immediate welfare protection and long-term economic adjustment.

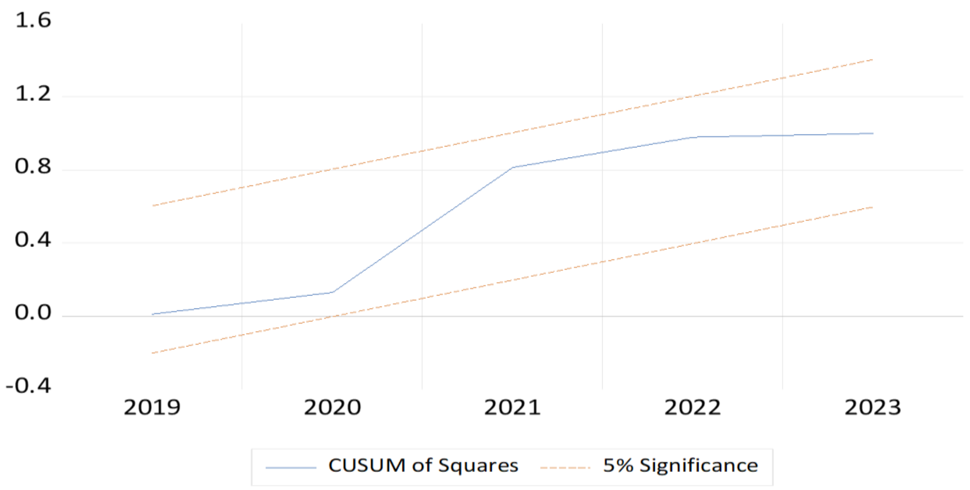
Stability Tests

To ascertain the stability, or otherwise, of the long-run coefficients of the $NARDL$ model, this study performs the cumulative sum $(CUSUM)$ and the cumulative sum of squares $(CUSUMSQ)$ tests. As shown in Figures 2 and 3, the $CUSUM$ and the $CUSUMSQ$ graphs indicate that the long-run coefficients are stable, as their plots remain within the 5% critical bounds. Thus, the graphs buttress the long-run impacts of the exogenous variables on economic welfare in Nigeria. Also, the study utilises the nonlinear dynamic multiplier to verify whether the exchange rate and petroleum pump price have a symmetric (when the broken red line is close to zero) or asymmetric (when the broken red line deviates from or is far from zero) relationship with economic welfare. The dynamic multiplier graphs in Figures 4 and 5 reveal the existence of an asymmetric relationship between the exchange rate and the petroleum pump price with economic welfare in Nigeria.



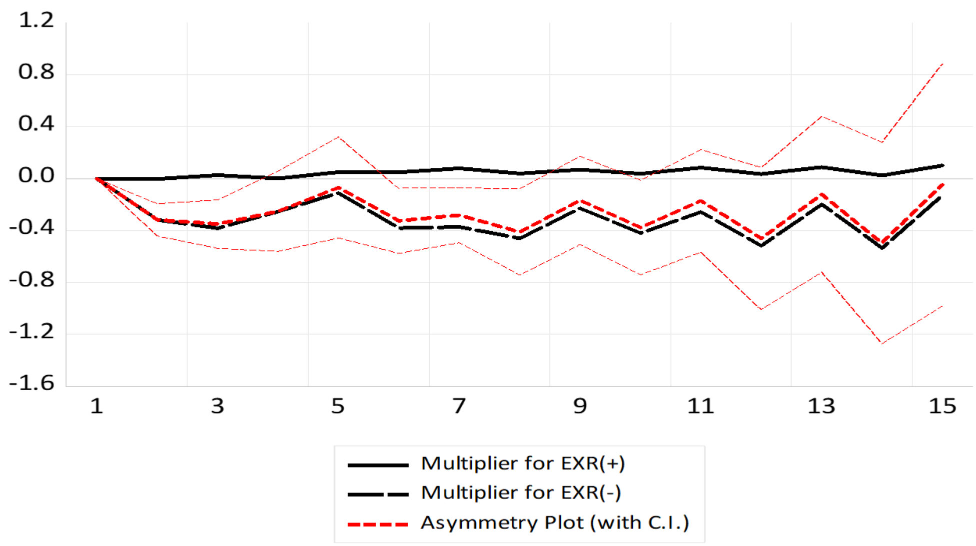
Source: Author’s analysis

Figure 2. Cumulative Sum (CUSUM)



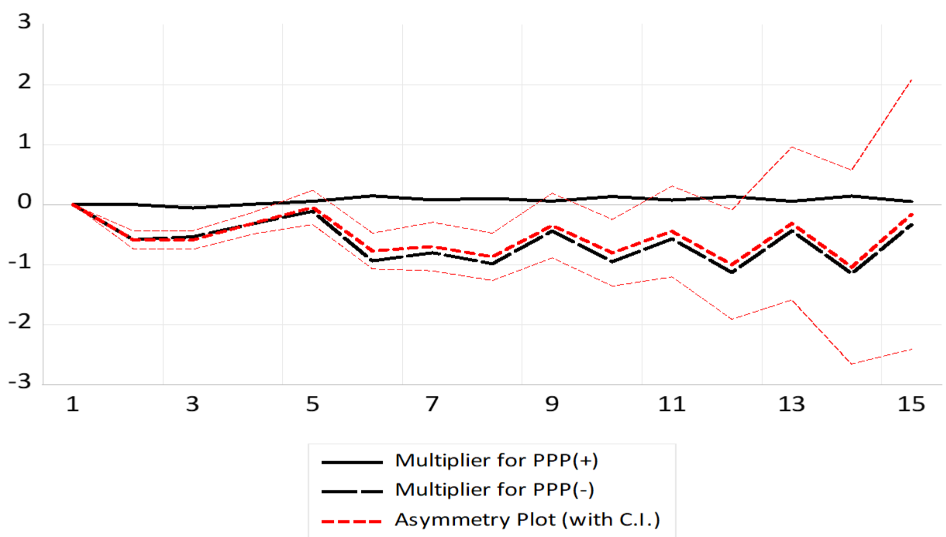
Source: Author's analysis

Figure 3. Cumulative Sum of Square (CUSUMSQ)



Source: Author's analysis

Figure 4. Dynamic Multiplier Graph of Exchange Rate



Source: Author's analysis

Figure 5. Dynamic Multiplier Graph of Petroleum Pump Price

Conclusion

This study examines the asymmetric effects of exchange rate fluctuations and petroleum pump price on economic welfare in Nigeria using the $NARDL$ approach. The results confirm the existence of both short-run and long-run asymmetries: exchange-rate depreciations and increases in the petroleum pump price led to more substantial and more persistent welfare losses than moderate and temporary welfare gains associated with appreciations and price reductions. As implied by the short-run dynamics, there is an immediate loss of welfare due to the impact of inflationary pass-through and the loss of household purchasing power. At the same time, the long-run analysis indicates that structural reforms and fiscal reallocation can offset the shocks to enhance welfare benefits. The findings also demonstrate a critical trade-off: short-term protection requires cushioning households from inflationary and petroleum price pressures. At the same time, long-term welfare depends on sound macroeconomic management, such as fiscal discipline and structural transformation. Sequel to the removal of the petroleum subsidy in Nigeria since May 2023, the findings reinforce the need for targeted social protection and inclusive measures to alleviate the social costs of reforms, while channelling fiscal savings into productive sectors to enhance long-term development.

Based on the findings of this study, the following recommendations are made to improve economic welfare in Nigeria. There is a need to maintain a unified macroeconomic policy to protect welfare in Nigeria. Exchange rate stability should be prioritised as persistent depreciation increases inflation and disproportionately affects consumers. In the aftermath of the removal of the petroleum subsidy in May 2023, compensatory measures such as targeted transfers, subsidised transport, and the availability of alternative energy sources are essential to safeguard vulnerable groups and sustain popular support for reform. Strengthening the social protection system is equally vital, as it will ensure that aid is well-targeted and responsive to exchange rate shocks and fuel price pressures. Similarly, macroeconomic policy should also address inflation and unemployment by investing in infrastructure, agriculture and small-scale enterprises to increase employment opportunities and support the purchasing power. Nigeria needs to reduce its reliance on oil revenues by diversifying the economy and allocating the fiscal savings from subsidy removal towards education, healthcare and infrastructure. Such proactive investments will transform short-term fiscal adjustments into sustainable welfare improvements.

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Use of AI tools declaration

The authors used AI tools (ChatGPT and DeepSeek) for language editing and grammar review of this manuscript. The authors are fully responsible for the content of this publication.

Conflict of interest

The authors declare no conflicts of interest.

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