

Investigating Sovereign Spread Divergence Due to Inflation and Revaluation Risk versus Unconventional Monetary Policy in Iran: Asymmetric Effects Using the NARDL Model

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Abstract: In the Iranian economy, sovereign spread divergence cannot be explained merely as a technical difference between government bond yields and a benchmark rate; rather, this phenomenon emerges within a multilayered structure characterized by chronic inflation, exchange-rate uncertainty, government fiscal constraints, and unconventional monetary interventions. In such a context, the sovereign spread, rather than primarily reflecting default risk in its conventional sense, mirrors the risk of declining purchasing power, instability of expectations, and uncertainty regarding policymakers' ability to preserve the value of the national currency. In other words, Iran's debt securities market considers not only the government's capacity for nominal repayment but also the real value of repayment at maturity and whether existing monetary and fiscal policies can prevent its erosion. It is precisely in this context that inflation and revaluation risk shift from peripheral factors to major determinants of sovereign spread divergence. The present study investigates sovereign spread divergence in the Iranian economy, with an emphasis on the effects of inflation and revaluation risk in the presence of unconventional monetary policies. The study aims to analyze the asymmetric effects of positive and negative shocks to macroeconomic variables on the sovereign spread using the nonlinear autoregressive distributed lag (NARDL) approach. For this purpose, the Economic Policy Uncertainty Index was constructed using principal component analysis (PCA) by combining fluctuations in monetary, fiscal, and exchange-rate policies and economic growth. The data were collected from the World Bank database for the selected period, and the research model was estimated by decomposing economic policy uncertainty and inflation into partial sums of positive and negative changes. The estimation results indicate that positive and negative shocks to the shadow economy, credit risk, liquidity risk, debt ratio, and non-performing loan ratio exert asymmetric effects on the sovereign spread index. The findings indicate that all model variables are statistically significant in the long run, while the error-correction component was estimated at -0.72 , reflecting the speed at which short-run deviations adjust toward long-run equilibrium. Diagnostic tests, including the LM and White tests, confirm the adequacy of the model, while structural stability tests (CUSUM and CUSUMQ) demonstrate the stability of the model coefficients over time. Accordingly, management of economic policy uncertainty and inflation control are identified as central mechanisms for reducing risk and stabilizing the sovereign spread in the Iranian economy.

Keywords: sovereign spread, economic policy uncertainty, inflation, Iranian economy, monetary policy

1. Introduction

Sovereign spreads constitute an important market-based indicator of the risk premium associated with government-related financing conditions, but their interpretation becomes substantially more complex in economies characterized by chronic inflation, recurrent exchange-rate depreciation, fiscal dominance, financial repression, and extensive monetary intervention. In conventional settings, sovereign spreads are often interpreted primarily as compensation for default risk, liquidity risk, term risk, and shifts in global risk appetite. In economies such as Iran, however, the observed spread may also embody concerns about the erosion of the real value of financial claims, persistent inflation expectations, exchange-rate instability, banking-sector fragility, and uncertainty regarding the credibility and effectiveness of monetary and fiscal policy. Under such circumstances, the distinction between nominal solvency and real repayment capacity becomes economically meaningful because creditors may expect nominal obligations to be honored while simultaneously facing substantial losses in purchasing power. This creates an environment in which sovereign spread divergence reflects not only conventional credit fundamentals but also the interaction among inflation, currency revaluation risk, policy uncertainty, banking vulnerabilities, and unconventional monetary interventions. Recent research on monetary policy in Iran has similarly emphasized that monetary impulses may be transmitted through multiple and potentially nonlinear channels, affecting financial markets, investment behavior, output, and broader macroeconomic stability [1-3]. These considerations make the Iranian case particularly suitable for an asymmetric investigation of sovereign spread dynamics.

The macroeconomic environment in which sovereign risk is priced is critically influenced by inflation. Inflation affects sovereign spreads through several channels, including the erosion of real returns, the adjustment of nominal interest rates, the weakening of policy credibility, the deterioration of bank balance sheets, and the repricing of domestic and external risk. Persistent inflation can raise uncertainty about future monetary conditions and increase the compensation required by investors for holding nominal assets. When inflation is accompanied by depreciation of the national currency, the associated revaluation risk becomes more pronounced because both domestic purchasing power and the external value of financial claims are exposed to deterioration. Evidence from the Iranian economy indicates that inflation may have nonlinear effects on economic performance when currency depreciation is taken into account, suggesting that the macroeconomic consequences of inflation cannot necessarily be represented by a constant linear coefficient [4]. Inflation and exchange-rate volatility may also worsen the quality of banking assets by increasing non-performing claims, with potentially different responses across regimes and economic conditions [5]. Such effects are relevant to sovereign spread formation because banking-sector weakness may ultimately increase contingent fiscal liabilities, tighten liquidity conditions, and reinforce perceptions of macro-financial risk. Accordingly, an increase in inflation may have a stronger effect on sovereign spreads than an equivalent decrease, particularly when inflationary shocks interact with exchange-rate depreciation and weak policy credibility.

The importance of inflation in Iran is inseparable from fiscal dominance and the institutional relationship between the government, banking system, and central bank. When fiscal imbalances are financed directly or indirectly through the banking network or monetary expansion, monetary policy may become subordinated to government financing requirements. In this environment, the expansion of central bank liquidity can stabilize short-run funding conditions while simultaneously increasing inflationary expectations and weakening the credibility of nominal anchors. Research on Iran has highlighted the inflationary consequences of government off-

budget operations and the role of fiscal dominance within the banking network, demonstrating that fiscal activities outside the conventional budgetary framework may generate monetary and inflationary pressures [6]. Earlier DSGE-based evidence has likewise shown that monetary and fiscal shocks can materially affect the real sector of the Iranian economy, confirming the importance of policy interactions rather than the isolated operation of individual instruments [7]. This interaction is especially relevant to sovereign spreads because the government's perceived fiscal position is partly endogenous to monetary conditions. Policies that temporarily reduce financing costs may therefore have the opposite effect over a longer horizon if they intensify inflation, increase currency depreciation, or undermine expectations of macroeconomic stability.

These concerns become even more important when conventional monetary instruments are constrained or ineffective. Unconventional monetary policy broadly refers to interventions that operate beyond standard adjustments in short-term policy rates and may include large-scale asset purchases, directed credit, liquidity support, balance-sheet expansion, yield management, forward guidance, and other measures designed to influence financial conditions. International evidence shows that unconventional monetary policies can compress yield curves and reshape the pricing of future interest rates, demonstrating their capacity to alter the term structure of financing costs beyond conventional policy channels [8]. At the same time, prolonged periods of exceptionally accommodative monetary policy may generate adverse side effects, including risk mispricing, excessive leverage, financial fragility, and weakened incentives for balance-sheet adjustment [9]. These findings are particularly relevant to an economy in which direct and indirect monetary interventions are frequently implemented under conditions of structural inflation and constrained financial markets. Contemporary monetary policy research therefore increasingly emphasizes that unconventional instruments should not be evaluated solely according to their immediate effect on liquidity or interest rates, but also according to their longer-term implications for financial stability, risk-taking, expectations, and asset pricing [10, 11].

In the Iranian context, the effectiveness and consequences of unconventional monetary policies may be even more heterogeneous because the transmission mechanism operates through a banking-dominated financial system, administratively influenced interest rates, fiscal pressures, exchange-rate disturbances, and segmented asset markets. Evidence suggests that conventional and unconventional monetary policies may exert different effects on economic growth under uncertainty, particularly when the economy is modeled within a dynamic stochastic general equilibrium framework [12]. More recent studies have also emphasized the implications of unconventional monetary policies for financial market stability in Iran and the possibility that monetary interventions may reduce instability in some circumstances while amplifying imbalances in others [3]. Similarly, monetary policies can affect the investment decisions of firms by changing financing costs, liquidity conditions, expected returns, and the relative attractiveness of alternative assets [2]. These transmission channels imply that the sovereign spread may respond not only to the level of monetary accommodation but also to the way markets interpret the sustainability, credibility, and inflationary consequences of that accommodation. A policy-induced reduction in nominal rates may initially compress spreads, whereas its subsequent effects on inflation expectations and currency valuation may eventually widen them.

The asymmetry of these responses is therefore a central analytical issue. Traditional linear models implicitly assume that positive and negative shocks of equal magnitude produce equal but opposite effects. Such an assumption may be restrictive in economies exposed to persistent inflation, exchange-rate instability, balance-sheet vulnerabilities, and nonlinear policy reactions. A rise in inflation, for example, may produce a much larger increase in sovereign spreads than the decline generated by an equivalent reduction in inflation, because adverse shocks

may trigger precautionary behavior, capital reallocation, higher liquidity preference, and a nonlinear reassessment of policy credibility. Likewise, tightening and easing monetary policies may affect macroeconomic variables differently across short-run and long-run horizons. Evidence from Iran using the nonlinear autoregressive distributed lag framework indicates that monetary-policy shocks may have asymmetric short- and long-run effects on macroeconomic variables, supporting the use of nonlinear specifications rather than models imposing symmetry by construction [13]. This methodological perspective is especially relevant when examining sovereign spreads because both financial-market responses and macroeconomic adjustment processes tend to be state dependent.

Another important dimension of sovereign spread dynamics concerns banking-sector risks. Sovereign and banking risks can reinforce one another through multiple channels, particularly when banks hold government-related claims, rely on central bank liquidity, or operate under implicit public guarantees. Credit risk and liquidity risk can increase the probability that financial distress is transferred to the public sector, while deteriorating sovereign conditions can weaken bank balance sheets and funding access. Inflation and exchange-rate volatility may further intensify this nexus by reducing borrowers' repayment capacity and increasing the burden of financial obligations. The observed relationship between inflation, exchange-rate volatility, and non-performing claims in Iran underscores the importance of incorporating banking-sector fragility into the analysis of broader risk pricing [5]. In addition, changes in monetary conditions can influence capital-market volatility and alter the transmission of risk between financial markets and the real economy [1]. Consequently, sovereign spread divergence should be understood within an integrated macro-financial framework in which monetary policy, inflation, credit quality, liquidity conditions, and public-sector risk are mutually connected.

The shadow economy constitutes another potentially important source of sovereign risk in economies where informal activity is substantial. A large underground economy can weaken the tax base, reduce fiscal transparency, complicate macroeconomic measurement, and constrain the government's capacity to mobilize stable non-inflationary revenues. These consequences may intensify fiscal dependence on banking resources or monetary financing and thereby strengthen the interaction between fiscal deficits, inflation, and sovereign risk. From this perspective, the shadow economy is not merely an institutional or tax-compliance issue; it can influence sovereign spreads through its effect on fiscal sustainability, policy credibility, and the informational environment in which investors assess macroeconomic fundamentals. When combined with persistent inflation and unconventional monetary interventions, a large informal sector may further obscure the distinction between temporary liquidity support and structurally inflationary financing. This mechanism strengthens the rationale for including shadow-economy indicators alongside conventional measures of credit, liquidity, debt, and monetary conditions in empirical models of sovereign spread behavior.

The broader international literature also demonstrates that sovereign risk has become increasingly sensitive to nontraditional sources of uncertainty. Climate-related shocks, natural disasters, environmental vulnerabilities, and transition risks are now recognized as relevant determinants of financial stability and sovereign credit conditions. Cross-country evidence shows that climate risk can affect financial stability through banking, asset-pricing, and macroeconomic channels [14], while climate uncertainty may generate spillovers in sovereign credit risk across countries [15]. More recent evidence further indicates that natural disasters can increase sovereign credit risk and produce contagion effects across markets, illustrating how large exogenous shocks can be incorporated into sovereign risk premia [16]. Although the present study focuses on inflation, revaluation risk, and monetary policy rather than climate shocks, this literature is relevant because it demonstrates a broader principle: sovereign spreads increasingly reflect multidimensional uncertainty rather than conventional default fundamentals alone. Financial

markets respond to the capacity of governments and institutions to absorb shocks, stabilize expectations, and preserve the real value of financial claims.

Related studies on climate-related financial risk also highlight how changing perceptions of risk can alter the pricing of government and corporate securities. Sovereign and quasi-sovereign securities may exhibit pricing differentials associated with environmental characteristics, as reflected in evidence on the sovereign “greenium” [17]. Green bond pricing has also been shown to incorporate climate-change risk, indicating that bond-market spreads can reflect risk factors that extend well beyond standard default measures [18]. Research has similarly examined whether government-related securities are systematically perceived as greener or less environmentally risky than private alternatives, demonstrating that public-sector status does not automatically eliminate risk-pricing differentials [19]. These studies are conceptually important for the present research because they reinforce the notion that spreads are composite prices of uncertainty. If environmental risk can be capitalized into sovereign and bond-market pricing, inflation risk, currency revaluation risk, policy uncertainty, and unconventional monetary interventions can likewise generate distinct and potentially asymmetric risk premia.

The interaction between central banking and broader forms of financial risk has also become increasingly important. Central banks are no longer viewed solely as institutions setting short-term policy rates; their balance sheets, credit-allocation policies, market interventions, and regulatory signals can influence the composition of financing and the distribution of risk throughout the financial system. Research on green central banking demonstrates both the potential and the limits of using credit policy to redirect financial flows, emphasizing that monetary and credit interventions may produce incomplete outcomes when structural constraints remain unaddressed [20]. Similarly, global evidence suggests that central bank policies can affect green bond issuance and broader financing conditions, confirming that monetary-policy frameworks shape not only aggregate liquidity but also the composition of capital-market activity [21]. This insight has direct relevance for sovereign spread analysis: unconventional interventions may reduce immediate financing stress, but their ultimate effect depends on whether they improve underlying macroeconomic credibility or merely postpone adjustment while increasing future inflation and valuation risks.

Systemic shocks can further modify the transmission of these policies. Recent monetary-policy research increasingly emphasizes the need to understand policy under digital transformation, financial innovation, and systemic disturbances rather than under stable conventional conditions [10, 11]. Financial systems exposed to major shocks often display nonlinear adjustment because liquidity demand, risk aversion, collateral values, and policy responses change simultaneously. Natural disasters, for example, can alter firms’ investment behavior and stimulate changes in green innovation strategies [22], while disaster losses may be distributed unevenly according to local vulnerability and institutional capacity [23]. Although these studies address different shocks, their findings illustrate why macro-financial relationships should not automatically be assumed to be linear or symmetric. In the Iranian economy, inflationary shocks, currency depreciation, policy interventions, and banking-sector stress can similarly interact in ways that generate regime-dependent and asymmetric sovereign spread responses.

At the global level, interconnected financial systems also mean that risks originating in one domain may propagate through multiple markets and jurisdictions. Climate-related financial risk mapping shows that exposures can be transmitted through financial networks and balance-sheet interdependencies, producing systemic effects that cannot be understood by examining isolated assets or institutions [24]. Sovereign credit-risk spillovers under climate uncertainty provide additional evidence that changes in perceived risk may transmit across sovereign markets rather than remaining country specific [15]. These findings support a network-oriented

interpretation of sovereign risk in which fiscal, monetary, banking, and market risks interact. For Iran, this perspective is especially important because domestic macroeconomic imbalances may be amplified by external financing constraints, exchange-rate pressures, and shifts in global financial conditions. Consequently, sovereign spreads may respond strongly to domestic uncertainty even when conventional measures of default probability change only modestly.

From a methodological perspective, these theoretical considerations imply that the empirical analysis of sovereign spreads should distinguish between positive and negative shocks and between short-run and long-run effects. The nonlinear autoregressive distributed lag model provides an appropriate framework because it allows explanatory variables to be decomposed into positive and negative partial sums while simultaneously estimating dynamic short-run adjustments and long-run equilibrium relationships. This property is particularly useful in the Iranian economy, where inflation increases and decreases are unlikely to be perceived symmetrically and where unconventional monetary interventions may generate effects that differ according to the direction of the shock and the time horizon considered. The existing Iranian evidence on asymmetric monetary-policy transmission reinforces the appropriateness of this framework [13]. Moreover, the broader literature on financial instability, unconventional monetary policy, and emerging systemic risks shows that nonlinearities are economically plausible rather than merely statistical artifacts [8, 9, 14].

Despite the growing literature on monetary policy, inflation, financial instability, and sovereign credit risk, an important research gap remains regarding the joint analysis of sovereign spread divergence, inflation-related revaluation risk, policy uncertainty, and unconventional monetary policy in Iran. Existing Iranian studies have examined the effects of monetary and fiscal shocks on the real economy [7], conventional and unconventional monetary policy under uncertainty [12], fiscal dominance and inflation [6], the asymmetric effects of monetary policy on macroeconomic variables [13], investment responses to central bank policy [2], financial-market stability under unconventional monetary interventions [3], and the implications of inflation and currency depreciation for economic performance and banking-sector distress [4, 5]. International research, meanwhile, has broadened the analysis of sovereign and financial risk to encompass unconventional monetary policy, climate uncertainty, natural disasters, green-bond pricing, and systemic risk transmission [16–19, 24]. However, comparatively limited attention has been devoted to whether positive and negative shocks to inflation, policy uncertainty, banking risks, debt conditions, and the shadow economy produce asymmetric changes in sovereign spreads within an inflationary and policy-constrained emerging economy.

Accordingly, the aim of the present study is to investigate the asymmetric short-run and long-run effects of inflation, economic policy uncertainty, shadow-economy fluctuations, credit risk, liquidity risk, debt conditions, and non-performing loans on sovereign spread divergence in Iran under unconventional monetary policy using the nonlinear autoregressive distributed lag (NARDL) approach.

2. Methods and Materials

The present study is applied in terms of purpose and descriptive-analytical in terms of nature and falls within the category of *ex post facto* research. Following the studies of David Case et al. (2024), Sadeghi et al. (2023), and Golpe et al. (2023), the present study examines sovereign spread divergence arising from inflation and revaluation risk in the context of unconventional monetary policy in Iran, with particular emphasis on the asymmetric effects estimated using the NARDL model. Macroeconomic data for Iran—including the exchange rate, inflation, monetary

policy, the underground economy, economic policy uncertainty, and other variables—were extracted and compiled from the World Bank's World Development Indicators (WDI) database. The research model is specified as follows:

$$TRSP_{it} = \beta_0 + \beta_1 VIX_{it} + \beta_2 CDS_{it} + \beta_3 LR_{it} + \beta_4 Debt_{it} + \beta_5 CPI_{it} + \beta_6 Shr_{it} + \beta_7 ESM_{it} + \varepsilon_{it}$$

By rewriting the above equation in error-correction form and decomposing the variables of economic policy uncertainty (*VIX*) and inflation targeting (*CPI*) into the partial sums of positive and negative changes, namely positive changes in economic policy uncertainty,

$$POSE = \sum_{j=1}^t \Delta LVIX^+$$

positive changes in inflation targeting,

$$POSS = \sum_{j=1}^t \Delta CPI^+$$

negative changes in economic policy uncertainty,

$$NEGE = \sum_{j=1}^t \Delta LVIX^-$$

and negative changes in inflation targeting,

$$NEGS = \sum_{j=1}^t \Delta CPI^-,$$

the nonlinear (asymmetric) short-run and long-run effects of positive and negative shocks to economic policy uncertainty and inflation targeting can be estimated as follows:

$$\begin{aligned} SP = & \mu + \sum_{i=1}^{n_1} \beta_i \Delta Shr_{t-i} + \sum_{i=1}^{n_2} \beta_i \Delta qCDS_{t-i} + \sum_{i=1}^{n_3} \beta_i \Delta LR_{t-i} \\ & + \sum_{i=1}^{n_4} \beta_i \Delta Debt_{t-i} + \sum_{i=1}^{n_4} \beta_i \Delta ESM_{t-i} \\ & + \sum_{i=1}^{n_5} \beta_{1,i} \Delta \ln POSE_{t-i} + \sum_{i=1}^{n_6} \beta_{2,i} \Delta \ln NEGE_{t-i} \\ & + \sum_{i=1}^{n_7} \beta_{3,i} \Delta \ln POSS_{t-i} + \sum_{i=1}^{n_8} \beta_{2,i} \Delta \ln NEGS_{t-i} + \varepsilon_t \end{aligned}$$

SP: The sovereign spread index, defined as the interest-rate spread, namely the difference between the lending interest rate and the interest rate paid on deposits (Sovereign Premium). This rate is measured as the difference between the average interest rate on loans granted and the average interest rate paid on deposits.

VIX: Economic policy uncertainty. Given the prevailing economic conditions in Iran and other developing countries—which, because of government budgetary requirements, frequently experience exchange-rate changes, growth in the money supply, inadequate economic growth, and increasing dependence on oil revenues, thereby

reducing attention to tax revenues—it is necessary to develop an index specifically suited to such economies that captures economic policy uncertainty. In this regard, following Loni et al. (2021), fluctuations in three policy dimensions—fiscal, monetary, and exchange-rate policies—are first calculated using the Hodrick–Prescott filter. A composite index is then estimated from the fluctuations in these variables and introduced as an economic uncertainty index. To measure this uncertainty index, fluctuations in monetary policy variables (monetary base and interest rate), fiscal policy variables (budget deficit and taxation), exchange-rate policy (exchange rate), as well as inflation and economic growth are first calculated using the Hodrick–Prescott filter. Subsequently, principal component analysis (PCA) is applied to construct the economic uncertainty index. PCA reduces the dimensionality of observations through a composite index and classification of similar observations. Under this method, correlated variables in a multidimensional space are summarized into a set of uncorrelated components, each of which is a linear combination of the original variables. The resulting uncorrelated components are known as principal components (PCs) and are derived from the eigenvectors of the covariance matrix or correlation matrix of the original variables. Reducing the number of variables and identifying the relational structure among variables are among the most important applications of PCA. Its principal advantage in econometrics is the mitigation of multicollinearity caused by the large number of explanatory variables included in a model.

Table 1. Economic Policy Uncertainty Index

Policy	Abbreviation	Operational Definition
Monetary policy: The set of factors and mechanisms through which the central bank regulates the quantity of money in the economy to achieve price stability and sustainable economic growth. Monetary policy instruments generally include open market operations, changes in banks' reserve requirements, and changes in the rediscount rate.	IR	Interest rate: The interest rate is the cost of borrowing money that must be paid by the borrower. In other words, it represents the opportunity cost borne by the lender, who could otherwise have earned a return by employing the funds rather than lending them.
	M	Monetary base: The total amount of highly liquid money created by the central bank. More specifically, it includes all currency in circulation, comprising banknotes and coins held by the public, together with certain deposits held by the public in banks and reserves held by commercial banks at the central bank.
Fiscal policy: Fiscal policy refers to the set of policies that alter government expenditure or revenue and thereby affect output. Fiscal policy was initially emphasized by Keynes, and its primary instruments are government expenditure and tax rates.	BD	Budget deficit: A budget deficit occurs when government expenditure exceeds government revenue. Such a policy is generally employed when the economy is in recession.
	TAX	Taxation: Taxation constitutes the transfer of a portion of society's income to the government or the government's share of profits generated through economic activity. Tax revenues are generally generated through two mechanisms. The first involves changes in tax rates, which are reduced by the government during recessions and increased during economic expansions. The second concerns expansion of the tax base, namely imposing taxes on goods or economic entities that had previously not been subject to taxation.
Exchange-rate policy: The mechanism through which the exchange rate is determined in an economy is referred to as the exchange-rate regime or exchange-rate policy. Two predominant regimes are generally identified: a fixed exchange-rate regime, in which the central bank intervenes in the market to support an officially determined exchange rate, and a floating exchange-rate regime, in which no	ER	Exchange rate: The exchange rate is the rate at which the domestic currency is valued relative to a foreign currency and at which exchange occurs. In other words, it is the amount of domestic currency required to purchase one unit of another country's currency.

predetermined official exchange rate is imposed and the market mechanism determines the exchange rate.		
Inflation rate: Inflation is the persistent average increase in the general price level. Under this definition, temporary increases in prices are not considered inflation.	INF	To calculate inflation, the price index of 386 goods in Iran is first calculated, and the average growth rate of the price index is regarded as inflation. Two principal price indices are commonly employed: the Producer Price Index and the Consumer Price Index.
Economic growth: Economic growth is defined as the growth of a country's gross domestic product over a specified period.	GDP	Gross domestic product: GDP is the total monetary value of final goods and services produced by resident economic units within a country during a specified period, generally annually or quarterly. GDP is therefore a flow variable, and its value represents the monetary flow of goods and services within a specified period, typically one year.

CDS: Credit risk, defined as:

$$CDS = \frac{\text{Outstanding bank loans} + \text{Outstanding bank commitments}}{\text{Total assets}}$$

LR: Liquidity risk, defined as:

LR

$$= \frac{(\text{Government banks' liabilities to the central bank} + \text{demand, savings, and short-term deposits}) - (\text{cash} + \text{coins} + \text{net balance})}{\text{Total assets}}$$

Debt ratio: The debt (borrowing) ratio is a financial ratio measuring the degree of leverage of banks in the economy. It is expressed as the ratio of banks' total liabilities to banks' total assets.

CPI: Consumer Price Index (inflation rate).

Shr: Shadow economy index. Because data on the size of the underground economy (shadow economy) are often unavailable, numerous researchers have used different methods to estimate it, some measuring tax evasion directly and others indirectly. In the present study, the size of the underground economy is estimated using Tanzi's currency-demand approach, as specified below. The basic assumption of Tanzi's monetary approach is that all transactions in the informal economy are carried out in cash to remain concealed. Therefore, to estimate the size of the informal economy using this approach and within an autoregressive distributed lag framework, Equation (1) estimates the ratio of currency in circulation to broad money while accounting for the tax-burden variable:

$$\left(\frac{C}{M_2}\right)_t = \beta_0 + \beta_1 TaxBar_t + \beta_2 RGDPG_t + \epsilon_t \quad (1)$$

where (C/M_2) represents the ratio of currency in circulation to broad money and is considered the dependent variable. $TaxBar$ represents the tax burden and is measured as the ratio of total tax revenues to gross domestic product. $RGDPG$ denotes economic growth. After estimating the currency-to-liquidity ratio,

$$\left(\frac{C}{M_2}\right)_t,$$

the coefficient of the tax-burden variable is set equal to zero without changing the other regression coefficients, and the currency-to-liquidity ratio in the absence of a tax burden,

$$\left(\frac{C}{M_2}\right)_{wt},$$

is calculated. In the subsequent stage, illegal money (IM) is obtained by multiplying the difference between the two aforementioned ratios by the money supply, according to Equation (2):

$$IM = \left[\left(\frac{C}{M_2} \right)_{TAX} - \left(\frac{C}{M_2} \right)_{WTAX} \right] \times M_2 \quad (2)$$

Next, legal money (LM) is obtained by subtracting the resulting amount from the money stock (M_2), and the velocity of money is calculated using the quantity equation of money, as shown in Equation (3):

$$V = \frac{GNP}{LM} \quad (3)$$

Assuming that the velocity of money is identical in the formal and informal sectors of the economy, the size of the underground economy (*Shadow*) is subsequently estimated by multiplying the money stock in the informal economy (illegal money) by the velocity of money, as specified in Equation (4):

$$Shadow = IM \times V \quad (4)$$

ESM: Ratio of non-performing loans to total loans.

3. Findings and Results

Principal component analysis has extensive applications across various scientific disciplines. One of its applications in economics is index construction, and in the present study this method is employed to construct the Economic Policy Uncertainty Index. In examining many economic phenomena, various factors with different dimensions and originating from different domains may influence the emergence of a phenomenon. Incorporating all such factors into econometric analyses not only produces technical difficulties such as multicollinearity but also increases model complexity and the likelihood of statistical error. Accordingly, techniques such as PCA provide a means of reducing such errors and generating more reliable results. Multivariate data analysis has a fundamental role in information analysis. Multivariate datasets include numerous states or variables for each observation. If a dataset contains n variables, each variable may possess several dimensions. Because understanding and intuitively interpreting multidimensional space is generally difficult, principal component analysis reduces the dimensionality of all observations through a composite index and the grouping of similar observations. PCA represents one of the most valuable applications of linear algebra and has been extensively employed in analytical fields ranging from neural networks to computer graphics because it provides a nonparametric method for extracting relevant information from complex datasets. Under this method, correlated variables in a multidimensional space are summarized into a set of uncorrelated components, each of which constitutes a linear combination of the original variables. The resulting uncorrelated components are referred to as principal components (PCs) and are derived from the eigenvectors of either the covariance matrix or the correlation matrix of the original variables. Generally, the two major applications of PCA are reducing the number of variables and identifying the relational structure among variables, effectively resulting in the classification of variables. Its principal econometric advantage is the elimination or reduction of multicollinearity arising from the large number of influential variables included in the model. Scatterplot 1 illustrates points along the two coordinate axes X_1 and X_2 . In the present study, the following variables are used to construct the composite Economic Policy Uncertainty Index.

IR: Interest rate; M: Monetary base; BD: Budget deficit; TAX: Taxation; ER: Exchange rate; INF: Inflation; GDP: Gross domestic product.

To determine the general orientation of the observations, an ellipse is drawn to demonstrate the correlation between the variables.

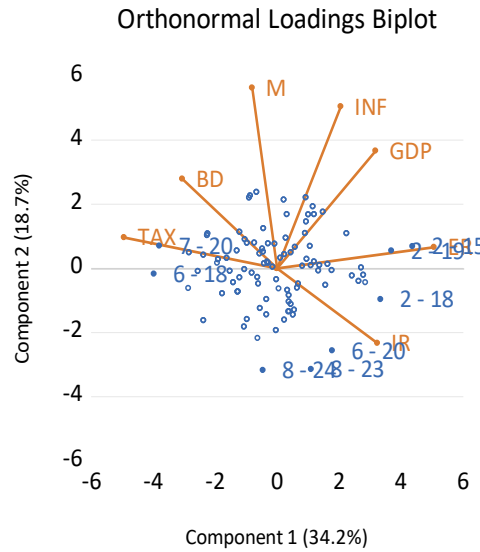


Figure 1. Transformation of Data into Principal Components

The principal direction of dispersion of the observations lies neither along X_1 , nor along X_2 , nor along any of the other original variables; rather, it lies between them, predominantly along the major axis of the ellipse. This axis is referred to as PC_1 , representing the first principal component of the variation in X_1 , X_2 , and the other variables. The second component (PC_2) lies along the minor axis of the ellipse, exactly perpendicular to PC_1 , and explains the remaining variation in X_1 , X_2 , and the other variables. PC_1 , PC_2 , and the remaining principal components constitute new axes for explaining X_1 , X_2 , and the other variables. Accordingly, X_1 , X_2 , and the remaining variables up to X_7 can be expressed as linear combinations of $PC_1, PC_2, \dots, PC_7$:

$$X_i = a_{i1}PC_1 + a_{i2}PC_2 + a_{i3}PC_3 + a_{i4}PC_4 + a_{i5}PC_5 + a_{i6}PC_6 + a_{i7}PC_7$$

The scores of the principal components can also be obtained using the following equations:

$$\begin{aligned} PC_1 &= W_1X_1 + W_2X_2 + W_3X_3 + W_4X_4 + W_5X_5 + W_6X_6 + W_7X_7 \\ PC_2 &= W_8X_8 + W_9X_9 + W_{10}X_{10} + W_{11}X_{11} + W_{12}X_{12} + W_{13}X_{13} + W_{14}X_{14} \\ &\vdots \\ PC_7 &= W_{43}X_{43} + W_{44}X_{44} + W_{45}X_{45} + W_{46}X_{46} + W_{47}X_{47} + W_{48}X_{48} + W_{49}X_{49} \end{aligned}$$

where W_i denotes the regression coefficient of the principal components on the variables.

Principal components may be calculated directly from the original dataset or, when the original data are unavailable, from the covariance matrix or correlation matrix. The correlation matrix is generally employed when variables have different measurement units or substantially different variances. When the correlation matrix is used, the analysis is effectively based on standardized variables with a mean of zero and a standard deviation of one.

The first extracted principal component accounts for the largest proportion of total variation in the dataset. This means that the first component is correlated with at least several of the variables. The second extracted component has two important characteristics. First, it accounts for the largest proportion of dataset variance that was not

captured by the first component; that is, the second component is correlated with several observed variables that are not highly correlated with the first component. Second, the second component is uncorrelated with the first component, meaning that the correlation between the two components is zero. The other components extracted through this method also possess these two characteristics.

The number of components extracted in each model is equal to the number of variables under examination. However, only a selected number of these components may be retained. Usually, the first two or three components account for a substantial proportion of total data variation. Accordingly, selecting the first two or three components is generally sufficient for subsequent analysis. In some circumstances, however, additional criteria must be considered when determining the number of components to retain. These criteria are as follows.

Criterion 1 (Scree Test): Plotting the eigenvalues against the corresponding principal components produces a scree plot. This plot demonstrates changes in the relative importance of the eigenvalues associated with each principal component. Figure 2 presents a hypothetical scree plot. As shown, the eigenvalue of the first vector is approximately 2.4, representing the variance explained by the first vector; the eigenvalue of the second vector is approximately 1.3; and the eigenvalue of the seventh vector is below 0.4. Initially, the decline in importance is rapid, after which the rate of decline levels off. The break point indicates the maximum number of principal components that should be retained. A number of PCs smaller than the value indicated by the break point may also be appropriate. Accordingly, Figure 2 suggests that either the first component alone or the first three components may be selected.

Criterion 2 (Eigenvalue): Components with eigenvalues greater than one are retained, while the remaining components are excluded.

Criterion 3 (Variance): Components explaining a larger proportion of total variation are considered sufficient for subsequent analysis. Typically, the first component accounts for the largest proportion of variance.

Scree Plot (Ordered Eigenvalues)

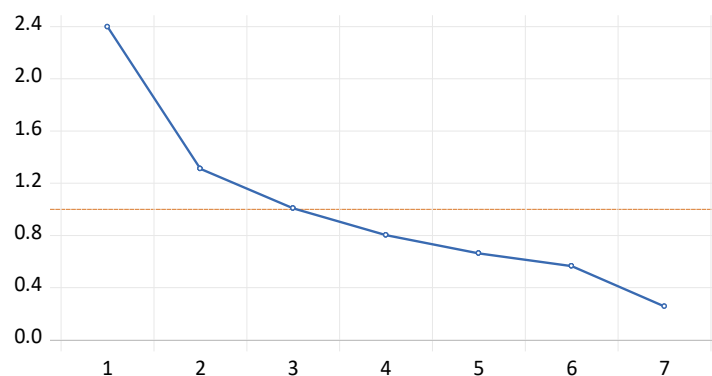


Figure 2. Eigenvalues of the Principal Components

Seven variables—interest rate, monetary base, budget deficit, taxation, exchange rate, inflation, and gross domestic product—were used to construct the Economic Policy Uncertainty Index.

Table 2. Correlation Matrix of Economic Policy Uncertainty Criteria

	IR	M	BD	TAX	ER	INF	GDP
IR	1.000000						
M	-0.066561	1.000000					
BD	-0.157517	0.199898	1.000000				
TAX	-0.344988	0.205616	0.312675	1.000000			
ER	0.432166	0.060198	-0.282109	-0.663387	1.000000		

INF	0.011065	0.121488	-0.009326	-0.225045	0.176880	1.000000	
GDP	0.021601	0.051334	-0.189983	-0.263352	0.383599	0.283143	1.000000

Table 1 indicates relatively strong correlations among the aforementioned criteria. Consequently, by reducing the dimensionality of the variables using principal component analysis, the Economic Policy Uncertainty Index can be extracted.

Table 2 indicates that the eigenvalue of the first component is larger than those of the remaining components and that this component accounts for approximately 25% of the total dispersion of the dataset. Therefore, this component is selected as the most appropriate component for constructing the index.

Table 3. Estimation Results for the Composite Economic Policy Uncertainty Index Using PCA

Principal Components Analysis

Eigenvalues: Sum = 7, Average = 1

Number	Value	Difference	Proportion	Cumulative Value	Cumulative Proportion
1	2.396286	1.085454	0.3423	2.396286	0.3423
2	1.310833	0.303116	0.1873	3.707119	0.5296
3	1.007717	0.206004	0.1440	4.714836	0.6735
4	0.801713	0.139636	0.1145	5.516549	0.7881
5	0.662077	0.096797	0.0946	6.178625	0.8827
6	0.565279	0.309184	0.0808	6.743905	0.9634
7	0.256095	—	0.0366	7.000000	1.0000

Eigenvectors (Loadings)

Variable	PC1	PC2	PC3	PC4	PC5	PC6	PC7
IR	0.351392	-0.253062	0.605601	0.131235	-0.123174	0.625273	-0.149568
M	-0.087394	0.610223	0.477813	-0.479860	-0.262111	-0.176898	-0.247838
BD	-0.331760	0.301830	0.389201	0.506716	0.622904	-0.050419	0.006613
TAX	-0.536287	0.104042	-0.009941	-0.223811	-0.111877	0.522700	0.604689
ER	0.549868	0.071204	0.243049	-0.128570	0.180626	-0.283190	0.710035
INF	0.223425	0.547003	-0.220610	0.586852	-0.477044	0.121988	0.123909
GDP	0.344883	0.396711	-0.381608	-0.291090	0.505374	0.454870	-0.176093

The relationship between the observed variables and the principal components can be expressed using factor loadings, while estimates of the principal components can be obtained using factor scores as follows:

$$PC_1 = 0.35IR - 0.08M - 0.33BD - 0.53TAX + 0.54ER + 0.22INF + 0.34GDP$$

Accordingly, the above expression represents the linear combination of the first principal component (PC_1) used to construct the Economic Policy Uncertainty Index.

Finally, the weights assigned to each sub-index within the composite Economic Policy Uncertainty Index are presented in the following table.

Table 4. Relative Importance of Variables in the Composite Economic Policy Uncertainty Index

Variable	Relative Importance
Interest rate	0.3423
Monetary base	0.1873
Budget deficit	0.1439
Taxation	0.1146
Exchange rate	0.0946
Inflation	0.0807

Gross domestic product	0.0366
Total	100

Table 4 presents the relative importance of the components of the overall Economic Policy Uncertainty Index, disaggregated by the selected variables under the rotated-components regression method. Based on the estimated weights, the interest rate has the largest contribution to the composite Economic Policy Uncertainty Index. Finally, using the PCA generation command, the composite Economic Policy Uncertainty Index is extracted and subsequently incorporated into the final research model.

The stationarity of the research variables is examined next. If time series exhibit coordinated movements over time, such co-movement may indicate the existence of a potential long-run equilibrium relationship. In other words, when two time-series variables move together, it may be possible to specify a long-run equilibrium relationship between them; such variables are said to be cointegrated. In simple terms, cointegration occurs when two time series move approximately along the same long-run trajectory. Economic analysis generally assumes that the variables specified within an economic theory have a long-run equilibrium relationship. In applied econometric analysis, estimation of long-run relationships between variables often assumes that their means and variances remain constant over time and are independent of the time dimension, thereby implicitly assuming behavioral stability. However, empirical research has demonstrated that behavioral stability is frequently absent in time-series variables. Consequently, conventional F - and t -tests obtained from estimation techniques in which stationarity is violated are invalid and may generate misleading results. This problem is known as spurious regression. Therefore, to ensure the validity of empirical results, researchers have revised conventional estimation procedures and systematically examine the stationarity of variables and the existence of cointegration among them.

Cointegration analysis has been regarded as one of the most revolutionary developments in econometrics since the mid-1980s. In simple terms, cointegration analysis examines the co-movement and coordinated behavior of variables even when those variables themselves are nonstationary and may trend upward or downward over time. Their joint movement can nevertheless produce a stable linear relationship in the long run and generate an equilibrium relationship among them. Accordingly, if no such linear relationship exists in the long run, the variables are considered not to be cointegrated. More generally, cointegration analysis provides a framework for estimating long-run equilibrium parameters in relationships involving nonstationary variables. It therefore represents a modern method for specifying, estimating, and testing dynamic models and can be used to evaluate the empirical validity of economic theories. Furthermore, cointegration analysis can also be employed to estimate short-run or disequilibrium parameters because the long-run parameters obtained from cointegration analysis can be incorporated into the estimation of short-run dynamics. Such specifications are known as error-correction models. Error-correction models connect short-run fluctuations in variables to their long-run equilibrium values. By definition, a time series is said to be strictly or strongly stationary if its joint probability density function, for any given value of k , is not a function of t . Under this condition, the distribution and consequently its moments, including the mean and variance, are independent of time. However, examining strict stationarity in practice is complicated, and strong stationarity is generally not required in time-series analysis. Instead, it is usually sufficient to examine stationarity in the weak sense. A time series is weakly stationary when its mean is constant and its autocovariance function for any given k does not depend on t .

As can be observed, the mean of a weakly stationary time series must be constant, while its autocovariance function for any given k must be independent of t . For this reason, a weakly stationary time series is also described

as covariance-stationary. The variance of such a series is likewise time invariant. Throughout the remainder of this study, the term stationarity refers to weak stationarity. The economic importance of stationarity lies in the fact that a stationary series does not permanently retain the effect of a shock. If a shock affects such a series, its impact is temporary and ultimately dissipates. This property is particularly important in cointegration analysis. Nonstationary time series may arise through two principal mechanisms. In the first case, the series x_t is a function of the time variable t , while the disturbance term is stationary. If the deterministic component is a first-degree polynomial in time, the mean of x_t becomes a function of time and the series is consequently nonstationary. In the second case, x_t is generated by a random-walk process, in which the disturbance term is a stationary series with zero mean and constant variance. Through recursive substitution and assuming a known initial value x_0 , it can be shown that the variance of x_t is not independent of time.

Both types of models exhibit a linear trend. In the deterministic-trend specification, the disturbance constitutes a temporary shock, whereas in the random-walk specification it represents a permanent shock. If an economic variable is affected by permanent shocks, the random-walk specification is appropriate. The two models are conventionally referred to as deterministic-trend and stochastic-trend models, respectively. In both cases, x_t may increase at a particular rate; if x_t is expressed logarithmically, this coefficient can be interpreted as its growth rate. However, shocks are temporary in the deterministic-trend model and permanent in the stochastic-trend model. The constant drift term in the latter is referred to as the drift or trend component. Unlike stochastic trends, deterministic trends do not generate fundamental problems for valid statistical inference or cointegration analysis; their effects can be controlled by explicitly including the trend term t in the model. Consequently, cointegration analysis is primarily concerned with stochastic trends. For some purposes, it may be desirable to detrend time series. In a deterministic-trend model, this can be achieved by regressing x_t on time, in which case the resulting residuals are trend-free. Stationarity can also be obtained through differencing:

$$\Delta x_t = x_t - x_{t-1} \quad (7)$$

Likewise, in the corresponding stochastic-trend model, first differencing can produce a stationary series:

$$\Delta x_t = \text{constant} + u_t \quad (8)$$

Equation (7) represents a trend-stationary process, whereas Equation (8) represents a difference-stationary process. This terminology was introduced by Nelson and Plosser.

Cointegration analysis is primarily concerned with difference-stationary processes (DSPs), that is, processes that become stationary after differencing.

For purposes of stationarity testing, the null hypothesis H_0 represents nonstationarity, whereas the alternative hypothesis H_1 represents stationarity.

When the relevant autoregressive coefficient indicates the presence of a unit root, the time series is nonstationary. Conversely, rejection of the unit-root condition implies that the series is stationary and integrated of order zero, denoted $I(0)$.

If a series is nonstationary, differencing can be applied to make it stationary. If it becomes stationary after first differencing, it is said to be integrated of order one, $I(1)$.

More generally, if a series becomes stationary after d differences, it is described as integrated of order d , or $I(d)$.

If k time series, $x_{1t}, x_{2t}, \dots, x_{kt}$, are all integrated of order d and a linear combination of these variables is integrated of order $d - b$, subject to $b > 0$, the variables are said to be cointegrated of order (d, b) .

If several time series are integrated to different orders, their linear combination will generally be integrated to the highest order involved. In the present study, the conventional Phillips–Perron unit-root test is employed.

Table 5. Phillips–Perron Test Results for the Model Variables

Variable	Test Statistic	Probability	Integration Status
CDS	0.706538	0.8635	—
D(CDS)	-10.89521	0.0000	I(1)
CPI	-0.743894	0.3871	—
D(CPI)	-9.166059	0.0000	I(1)
DEBT	0.665645	0.9531	—
D(DEBT)	-12.51469	0.0000	I(1)
ESM	-2.623637	0.2728	—
D(ESM)	-5.916729	0.0001	I(1)
LR	-2.980257	0.1515	—
D(LR)	-7.838104	0.0000	I(1)
SHR	1.757753	0.9953	—
D(SHR)	-11.46529	0.0000	I(1)
TRSP	-0.800990	0.3620	—
D(TRSP)	-9.512460	0.0000	I(1)
VIX	-1.457232	0.1332	—
D(VIX)	-11.19993	0.0000	I(1)

Based on the unit-root test results for the model variables, all variables become stationary after first differencing and are therefore integrated of order one, $I(1)$.

Two principal techniques are commonly used to examine cointegration: the Engle–Granger method and the Johansen method. Under the Engle–Granger approach, the existence of only one long-run relationship among two or more variables is investigated. When the number of variables exceeds two, this approach may encounter difficulties in identifying multiple long-run relationships because only one residual component is incorporated into the equation (Nowferesti, 2008). Furthermore, differencing may result in the loss of some information. The Johansen approach has several advantages over alternative cointegration tests. One advantage is that it does not require differencing variables solely to achieve stationarity, because differencing can eliminate important long-run equilibrium properties among variables. Instead, the Johansen methodology calculates an error-correction component and incorporates it into equations formulated in differences, thereby preserving long-run equilibrium properties. Under this method, the cointegrating vectors—that is, the coefficients representing long-run equilibrium relationships among variables—are identified and estimated. The relationship between the model and cointegration makes it possible to derive the cointegrating vectors directly from the coefficients of the vector autoregressive model (Nowferesti, 1999). In the Johansen–Juselius cointegration procedure employed in the present study to examine long-run relationships among model variables, determining the cointegration rank constitutes the first essential step (Johansen & Juselius, 1991). In the second stage, following determination of the cointegration rank, the existence of cointegration among the model variables is assessed.

According to the procedure proposed by Johansen, the trace test (λ_{Trace}) and maximum-eigenvalue test (λ_{max}) are employed to determine the number of cointegrating vectors. As shown in the following table, under the trace test, the null hypothesis that no cointegrating vector exists is rejected, while the alternative hypothesis indicating the existence of more than one cointegrating vector is supported. The principal distinction between the λ_{max} and λ_{Trace} statistics is that the λ_{max} test specifies a more explicit alternative hypothesis. The results of the two tests may

occasionally be inconsistent. Because the $\lambda_{\max}$ test is associated with a more specific alternative hypothesis, if contradictory findings arise, selecting the smaller number of cointegrating vectors is generally preferred (Enders, 2007). The results are presented below.

Table 6. λ_{Trace} and $\lambda_{\max}$ Tests

Unrestricted Cointegration Rank Test (Trace)				
Hypothesized Number	Eigenvalue	Trace Statistic	0.05 Critical Value	Probability
None *	0.832893	206.6339	159.5297	0.0000
At most 1 *	0.761192	144.0147	125.6154	0.0023
At most 2	0.660660	93.89145	95.75366	0.0667
At most 3	0.474198	56.06512	69.81889	0.3754
At most 4	0.383901	33.56602	47.85613	0.5257
At most 5	0.217278	16.61384	29.79707	0.6684
At most 6	0.159240	8.039623	15.49471	0.4612
At most 7	0.054701	1.968902	3.841466	0.1606

Trace test indicates two cointegrating equations at the 0.05 level.

Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized Number	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Probability
None *	0.832893	62.61914	52.36261	0.0033
At most 1 *	0.761192	50.12327	46.23142	0.0183
At most 2	0.660660	37.82633	40.07757	0.0878
At most 3	0.474198	22.49910	33.87687	0.5690
At most 4	0.383901	16.95219	27.58434	0.5841
At most 5	0.217278	8.574212	21.13162	0.8652
At most 6	0.159240	6.070721	14.26460	0.6042
At most 7	0.054701	1.968902	3.841466	0.1606

Under both tests, the hypothesis that there are at least two cointegrating vectors is supported.

Engle and Granger (1987) argued that if the Dickey–Fuller test is applied to the model residuals and the residual time series is found to be stationary, this provides evidence of cointegration. Nevertheless, caution must be exercised when applying this method because the conventional critical values are not fully appropriate; instead, the critical values developed by Engle and Granger should be used. Under this procedure, stationarity and nonstationarity are examined through the Dickey–Fuller unit-root test, with the following hypotheses:

H_0 : No cointegration

H_1 : Cointegration

The test results are reported in the following table.

Table 7. Cointegration Test

Engle–Granger Test Statistic	Probability
-5.730189	0.0000

To estimate the sovereign spread index model, the dynamic coefficients of the model and the ECM specification are estimated to determine the speed of adjustment of deviations from long-run equilibrium. The NARDL model was selected to estimate the long-run coefficients as well as the error-correction coefficient and short-run coefficients associated with the equation. The dynamic coefficient estimates obtained from this model are presented in Table 8.

Table 8. Estimation Results for the Sovereign Spread Index Model

Variable	Coefficient	Standard Error	<i>t</i> -Statistic	Probability
TRSP(-1)	0.196692	0.095788	2.053411	0.0549
DSHR_POS	0.112346	0.033569	3.346760	0.0036
DSHR_NEGE	-0.474573	0.124633	-3.807772	0.0066
DCDS_POS	0.203325	0.054194	3.751771	0.0045
DCDS_NEGE	-0.116226	0.039950	-2.909323	0.0094
DLR_POS	0.850933	0.336984	2.525142	0.0212
DLR_NEGE	-0.137959	0.041832	-3.297954	0.0040
DDEBT_POS	0.513400	0.157400	3.262762	0.0098
DDEBT_NEGE	-0.211651	0.068448	-3.092118	0.0047
DESM_POS	0.242560	0.040182	6.036585	0.0002
DESM_NEGE	-0.228524	0.054462	-4.196024	0.0005
DVIX	0.326966	0.146003	2.239450	0.0339
DCPI	0.123968	0.028846	4.297628	0.0004
ECM(-1)	-0.298586	0.126089	-2.368049	0.0256

Diagnostic Tests

Normality Test	Wald Test	Heteroskedasticity Test	LM Test
87.9865 (0.0000)	45.15060 (0.0000)	0.2684 (0.9586)	0.45263 (0.6123)

The *t*-statistic is used to assess the statistical significance of the coefficients of the independent variables. The null hypothesis of the *t*-test is specified as:

$$H_0: \beta_i = 0$$

and is evaluated using the conventional *t*-statistic:

$$t = \frac{\hat{\beta}_i}{SE(\hat{\beta}_i)}$$

To determine whether the null hypothesis should be accepted or rejected, the calculated *T*-statistic is compared with the tabulated *t*-value at $N - K$ degrees of freedom and a 95% confidence level. If the absolute value of the calculated *T*-statistic exceeds the critical *t*-value, the test statistic falls within the rejection region and the null hypothesis is rejected. In this case, the relevant coefficient is statistically significant at the 95% confidence level, indicating the existence of a relationship between the corresponding independent variable and the dependent variable.

The Wald test results reported at the bottom of the preceding table assess the symmetry or asymmetry of positive and negative shocks to the shadow economy, credit risk, liquidity risk, debt ratio, and ratio of non-performing loans to total loans in the short run. The results indicate that the null hypothesis of equality between the coefficients associated with positive and negative shocks to the shadow economy, credit risk, liquidity risk, debt ratio, and non-performing-loan ratio is rejected, while the alternative hypothesis is accepted. Therefore, the asymmetric effects of positive and negative shocks to the shadow economy, credit risk, liquidity risk, debt ratio, and non-performing-loan ratio on the sovereign spread index are confirmed.

The estimated results indicate that the model variables are statistically significant. The findings demonstrate that negative fluctuations in the shadow economy have a negative effect, whereas positive fluctuations in the shadow economy have a positive effect on the sovereign spread index in the short run. Thus, in this model, the shadow economy directly affects the sovereign spread. Decomposition of shadow-economy fluctuations into positive and

negative components indicates that the sovereign spread index responds differently to each type of fluctuation, thereby demonstrating an asymmetric effect. To examine the validity of this asymmetry more precisely, the *t*-test was applied. Because the null hypothesis of equal coefficients is rejected, coefficient inequality is established, thereby supporting the asymmetric effect of shadow-economy fluctuations on the sovereign spread.

The findings demonstrate that negative fluctuations in credit risk exert a negative effect, whereas positive fluctuations in credit risk exert a positive effect on the sovereign spread index in the short run. Accordingly, credit risk directly affects the sovereign spread in this model. By decomposing credit-risk fluctuations into positive and negative components, the estimated coefficients show that the sovereign spread index responds differently to each type of fluctuation, indicating asymmetric effects. To investigate the validity of this asymmetry more precisely, the *t*-test was employed. Because the null hypothesis of coefficient equality is rejected, coefficient inequality is established, thereby confirming the asymmetric effect of credit-risk fluctuations on the sovereign spread.

The findings demonstrate that negative fluctuations in liquidity risk exert a negative effect, whereas positive fluctuations in liquidity risk exert a positive effect on the sovereign spread index in the short run. Accordingly, liquidity risk directly affects the sovereign spread in this model. By decomposing liquidity-risk fluctuations into positive and negative components, the estimated coefficients indicate that the sovereign spread index responds differently to each type of fluctuation, thereby demonstrating asymmetric effects. To investigate the validity of these asymmetric effects more precisely, the *t*-test was employed. Because the null hypothesis of equal coefficients is rejected, coefficient inequality is established, confirming the asymmetric effects of liquidity-risk fluctuations on the sovereign spread.

The findings demonstrate that negative fluctuations in the debt ratio exert a negative effect, whereas positive fluctuations in the debt ratio exert a positive effect on the sovereign spread index in the short run. Accordingly, the debt ratio directly affects the sovereign spread in this model. By decomposing debt-ratio fluctuations into positive and negative components, the corresponding coefficients indicate that the sovereign spread index responds differently to each type of fluctuation, thereby demonstrating asymmetric effects. To evaluate this asymmetry more precisely, the *t*-test was employed. Because the null hypothesis of coefficient equality is rejected, inequality between the coefficients is established, thereby confirming the asymmetric effect of debt-ratio fluctuations on the sovereign spread.

The findings indicate that negative fluctuations in the ratio of non-performing loans to total loans have a negative effect, whereas positive fluctuations in this ratio have a positive effect on the sovereign spread index in the short run. Accordingly, the ratio of non-performing loans to total loans directly affects the sovereign spread in this model. By decomposing fluctuations in the non-performing-loan ratio into positive and negative components, the corresponding coefficients indicate that the sovereign spread index responds differently to each fluctuation, thereby demonstrating an asymmetric effect. To assess the validity of this asymmetry more precisely, the *t*-test was employed. Because the null hypothesis of coefficient equality is rejected, inequality of the coefficients is established, thereby confirming the asymmetric effect of fluctuations in the non-performing-loan ratio on the sovereign spread.

The short-run coefficients extracted from the asymmetric error-correction equation indicate that the coefficients of economic policy uncertainty and the Consumer Price Index are positive and equal to 0.326966 and 0.123968, respectively. Given that their calculated probabilities are below 0.05, both coefficients are statistically significant at the 95% confidence level.

Finally, the value of the ECM is reported as -0.72 , which indicates two important points. First, because the ECM coefficient is negative, convergence from short-run disequilibrium toward long-run equilibrium is confirmed.

Second, the magnitude of the ECM coefficient represents the speed at which short-run disequilibrium adjusts toward long-run equilibrium.

To ensure the validity and reliability of the results obtained from the estimated model, diagnostic tests for serial correlation and heteroskedasticity were conducted on the residuals of the NARDL specification.

Table 9. Diagnostic Test Results for the Model

Diagnostic Test	Result
LM serial correlation test	Prob. = 0.4852
Breusch–Pagan–Godfrey heteroskedasticity test	Prob. = 0.2268
Ramsey RESET specification-error test	Prob. = 0.3690

As indicated by the diagnostic test results presented in the table above, the estimated model does not exhibit heteroskedasticity, specification error, or serial correlation. These findings support the validity of the estimated model results.

To calculate the model's long-run statistic using the Banerjee–Dolado approach, the same dynamic model is employed. The long-run coefficients associated with the X variables are obtained from the relevant dynamic-model relationship.

To determine whether the long-run relationship generated by this method is spurious, the hypothesis of no long-run relationship is tested.

The null hypothesis represents the absence of cointegration or a long-run relationship because a necessary condition for short-run dynamics to converge toward long-run equilibrium is that the sum of the relevant coefficients be less than one. The alternative hypothesis represents the existence of cointegration or a long-run relationship.

To conduct this test, one is subtracted from the sum of the coefficients of the lagged dependent variable, and the result is divided by the corresponding standard error.

If the absolute value of the resulting t -statistic exceeds the absolute value of the critical values reported by Banerjee, Dolado, and Mestre, the null hypothesis is rejected and the existence of cointegration, or a long-run relationship, is accepted.

The calculated statistic is equal to -8.38 . Since its absolute value is larger than the Banerjee–Dolado–Mestre critical value of -3.27 , the null hypothesis of no long-run relationship is rejected. Accordingly, the model variables are cointegrated.

The calculated statistic is obtained as follows:

$$\frac{0.196692 - 1}{0.095788} = -8.386311$$

The long-run NARDL estimation results are presented in Table 10.

Table 10. Long-Run NARDL Estimation

Variable	Coefficient	Standard Error	t -Statistic	Probability
SHR	0.237409	0.102221	2.322508	0.0321
CDS	0.123066	0.026111	4.713159	0.0002
LR	0.225637	0.062077	3.634768	0.0019
DEBT	0.251691	0.097504	2.581332	0.0188
ESM	0.063825	0.023099	2.763094	0.0145
VIX	0.070174	0.026734	2.624946	0.0191

CPI	0.437643	0.207664	2.107460	0.0494
C	0.376683	0.176713	2.131606	0.0471

The results indicate that all variables included in the model are statistically significant in the long run at the 5% significance level.

Structural stability is examined using the cumulative sum of recursive residuals (CUSUM) and the cumulative sum of squared recursive residuals (CUSUMQ), originally proposed by Brown et al.

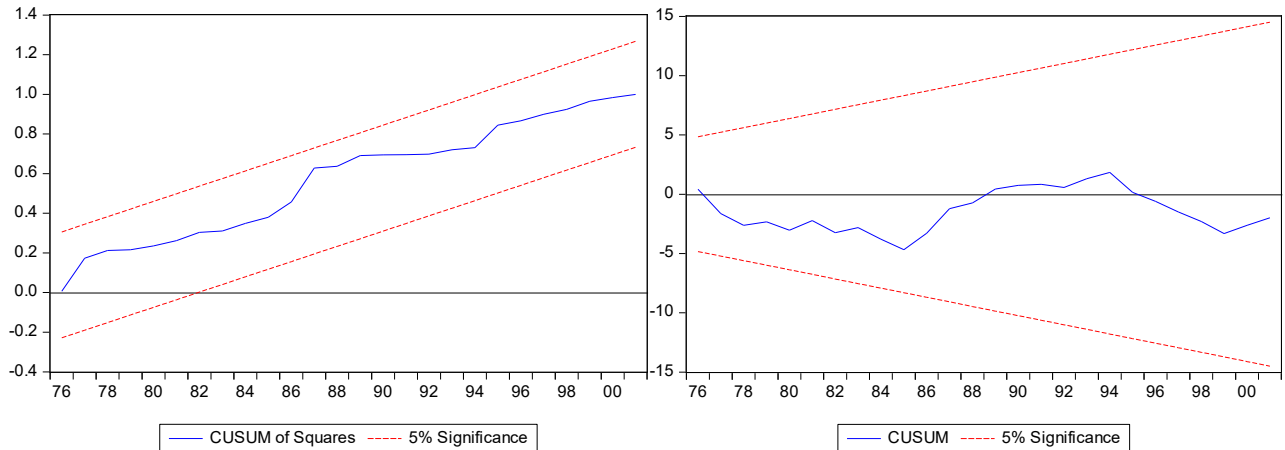


Figure 3. CUSUM and CUSUMQ Stability Tests

As shown, the plots of the cumulative residuals and the cumulative squared residuals remain between the two straight boundary lines representing the 95% confidence interval.

If the plotted statistics remain within the confidence interval, the null hypothesis of no structural break is accepted. Conversely, if either plot moves outside the confidence interval or crosses one of the confidence boundaries, the null hypothesis of no structural break is rejected, indicating the presence of structural instability. The CUSUM statistic is primarily used to identify systematic changes in regression coefficients, whereas the CUSUMQ statistic is particularly useful when deviations from coefficient stability are sudden and irregular.

4. Discussion and Conclusion

The present study examined the asymmetric effects of macroeconomic and financial shocks on sovereign spread divergence in Iran within a nonlinear autoregressive distributed lag framework. The findings showed that positive and negative shocks to the shadow economy, credit risk, liquidity risk, the debt ratio, and the ratio of non-performing loans to total loans exerted statistically significant and asymmetric effects on the sovereign spread in the short run. Specifically, positive shocks to the shadow economy, credit risk, liquidity risk, debt ratio, and non-performing loans increased the sovereign spread, whereas negative shocks to these variables reduced it. The Wald test further confirmed that the coefficients associated with positive and negative shocks were not equal, indicating that the sovereign spread reacts differently depending on the direction of the shock. In addition, economic policy uncertainty and inflation exerted positive and statistically significant short-run effects on the sovereign spread. In the long run, all explanatory variables, including the shadow economy, credit risk, liquidity risk, debt ratio, non-performing loans, economic policy uncertainty, and inflation, were statistically significant. The negative and significant error-correction coefficient confirmed convergence toward long-run equilibrium after short-run disturbances, while the diagnostic tests indicated the absence of serial correlation, heteroskedasticity, and major

specification error. The stability tests also confirmed the structural stability of the estimated coefficients over the sample period. Taken together, the findings indicate that sovereign spread divergence in Iran is generated by an interconnected set of monetary, fiscal, banking, and macroeconomic risks rather than by a single dimension of sovereign default risk.

One of the central findings was the positive effect of economic policy uncertainty on the sovereign spread in both the short and long run. This result implies that increasing uncertainty regarding monetary, fiscal, and exchange-rate policies raises the risk premium demanded in the financial system. In an economy characterized by recurrent policy changes, inflationary financing pressures, and exchange-rate instability, uncertainty affects expectations concerning both nominal returns and the future real value of financial claims. The result is consistent with the broader argument that financial markets respond not only to implemented policy but also to uncertainty regarding the persistence, timing, and credibility of policy interventions. The finding is aligned with studies emphasizing the changing transmission mechanisms of monetary policy under systemic shocks and nontraditional policy environments [10, 11]. It is also consistent with evidence showing that unconventional monetary policy can materially affect financial-market stability in Iran [3]. When policy actions are difficult to anticipate or are perceived as temporary responses to deeper structural imbalances, their stabilizing effects may be offset by higher uncertainty premiums. This mechanism provides a plausible explanation for the positive association between policy uncertainty and sovereign spread divergence observed in the present study.

The positive and significant effect of inflation on the sovereign spread represents another major result. In the short run, the coefficient of the Consumer Price Index was positive and statistically significant, while the long-run coefficient was also positive and significant. This indicates that inflationary pressures raise sovereign financing risk in a persistent manner. The finding can be interpreted through several mechanisms. First, higher inflation lowers the expected real return on nominal financial assets and therefore increases the compensation required by investors. Second, persistent inflation may weaken confidence in monetary policy and intensify expectations of additional depreciation of the national currency. Third, inflation affects banking-sector balance sheets and debt-servicing capacity by changing real income, financing costs, and repayment conditions. The result is compatible with evidence that inflation and exchange-rate volatility contribute to the deterioration of non-performing bank claims in Iran [5]. It also aligns with findings indicating that inflation can exert nonlinear effects on Iranian economic growth when national-currency depreciation is considered [4]. These studies support the interpretation that inflation in Iran operates through interconnected macroeconomic and financial channels and therefore has consequences that extend beyond the general price level.

The positive long-run relationship between inflation and the sovereign spread is also consistent with the role of fiscal dominance in the Iranian economy. Persistent fiscal imbalances may create incentives for monetary accommodation, indirect deficit financing through the banking system, or liquidity expansion, thereby strengthening inflationary pressures. Under such conditions, market participants may interpret higher inflation as a signal of fiscal-monetary imbalance rather than as an isolated price-level phenomenon. The finding is therefore consistent with evidence demonstrating the importance of government off-budget operations and fiscal dominance in generating inflationary pressures within Iran's banking network [6]. Earlier evidence also shows that monetary and fiscal shocks significantly affect Iran's real economy and that policy interactions should be considered jointly rather than independently [7]. Consequently, the positive inflation coefficient obtained in this study can be interpreted as reflecting not only direct inflation risk but also the market's assessment of underlying fiscal and monetary credibility.

The results further demonstrated asymmetric effects of credit risk on sovereign spreads. Positive shocks to credit risk increased the sovereign spread, whereas negative shocks reduced it, and the magnitudes of these effects differed. This finding implies that deterioration in bank credit quality is priced more strongly than improvement in credit conditions. Such asymmetry is economically plausible because worsening credit risk may immediately increase concerns regarding bank solvency, financial-sector support requirements, and contingent fiscal liabilities, whereas reductions in credit risk may improve confidence more gradually. The result is consistent with the relationship between macroeconomic instability and deteriorating banking claims documented for Iran [5]. It is also compatible with the broader observation that monetary policy affects the financial system through credit conditions, investment financing, and risk-taking behavior [2]. When credit risk rises, investors may anticipate greater intervention by the government or central bank, thereby transferring part of the perceived banking-sector risk to the sovereign balance sheet.

A similar asymmetric pattern was identified for liquidity risk. Positive liquidity-risk shocks had a strong positive effect on the sovereign spread, whereas negative shocks reduced it. The magnitude of the positive coefficient suggests that deterioration in liquidity conditions can rapidly increase sovereign risk pricing. Liquidity shortages may raise banks' dependence on central bank facilities, increase funding costs, and intensify pressure for monetary accommodation. In a banking-oriented economy, such conditions can transmit quickly from financial institutions to broader sovereign risk perceptions. This interpretation accords with studies highlighting the importance of unconventional monetary policy in stabilizing financial conditions when conventional policy instruments are constrained [3, 12]. At the same time, evidence from international research warns that prolonged or excessively accommodative monetary intervention may generate adverse side effects when liquidity provision persists for extended periods [9]. Therefore, the positive liquidity-risk coefficient may reflect both immediate financial stress and expectations regarding the future monetary consequences of liquidity support.

The asymmetric effect of the debt ratio further reinforces the importance of leverage and balance-sheet fragility in sovereign spread dynamics. Positive shocks to the debt ratio increased the sovereign spread, while negative shocks reduced it. Higher leverage raises the vulnerability of financial institutions to adverse changes in interest rates, asset quality, and liquidity conditions. In such circumstances, even when the debt formally belongs to banks rather than directly to the government, market participants may anticipate implicit guarantees or policy intervention if systemic distress emerges. This can create a sovereign-bank nexus in which increasing private or banking-sector leverage contributes to sovereign risk pricing. The result is consistent with studies showing that central bank policies affect investment and financing decisions through changes in funding conditions and financial leverage [2]. It is also aligned with research emphasizing that unconventional monetary policies may alter the distribution of risk within financial markets rather than merely reducing aggregate financing costs [8]. Thus, the debt ratio acts not only as a balance-sheet indicator but also as a channel through which financial-sector fragility can be transmitted to the sovereign risk premium.

The ratio of non-performing loans to total loans also showed significant asymmetric effects. Positive shocks increased the sovereign spread, whereas negative shocks reduced it. The result is important because non-performing loans represent accumulated credit deterioration and reduced banking-sector efficiency. An increase in non-performing loans constrains the ability of banks to extend new credit, reduces profitability, increases provisioning requirements, and may ultimately necessitate regulatory or fiscal intervention. The finding is consistent with evidence that inflation and exchange-rate volatility significantly influence non-performing banking claims in Iran [5]. In periods of inflation and currency depreciation, borrower repayment capacity may deteriorate

unevenly across sectors, while the real value of bank assets and liabilities changes in different directions. Consequently, rising non-performing loans may be interpreted as evidence of systemic vulnerability and may therefore increase sovereign spreads. The asymmetry identified in the present study additionally suggests that deterioration in asset quality produces a stronger and faster repricing of risk than improvement in asset quality.

The shadow economy was another significant determinant of sovereign spread behavior. Positive shocks to the shadow economy increased the sovereign spread, whereas negative shocks reduced it, with different magnitudes across the two directions. This finding can be explained by the fiscal consequences of informal economic activity. A larger shadow economy narrows the effective tax base, reduces transparency, weakens the reliability of fiscal statistics, and can increase dependence on alternative sources of financing. In an economy where government financing and banking-sector liquidity are closely connected, such conditions may increase the probability of inflationary financing and reinforce uncertainty regarding public-sector solvency in real terms. The result therefore complements the findings concerning fiscal dominance and inflation [6]. It also helps explain why sovereign spread divergence in Iran may respond to structural variables that are not directly represented in conventional sovereign-risk models. The shadow economy affects fiscal capacity and policy credibility, both of which influence the perceived sustainability of sovereign obligations.

The findings regarding asymmetric transmission are strongly consistent with previous evidence obtained using nonlinear macroeconomic models for Iran. Ramehdoost Bonab et al. showed that positive and negative monetary-policy shocks can exert different effects on macroeconomic variables in the short and long run, demonstrating the importance of allowing for nonlinear adjustment [13]. The current study extends this logic to sovereign spreads by showing that changes in financial and macroeconomic risk variables are also direction dependent. Such asymmetry may arise from loss aversion, nonlinear expectations, regime changes, balance-sheet constraints, and differences in the speed with which adverse and favorable information is incorporated into financial prices. From a policy perspective, this means that reversing an adverse shock may not automatically restore the sovereign spread to its previous level. If risk perceptions adjust more rapidly to deterioration than to improvement, policy credibility must be rebuilt over time rather than assumed to recover immediately after macroeconomic stabilization.

The estimated error-correction mechanism also provides important evidence regarding dynamic adjustment. The negative and statistically significant error-correction coefficient confirms that deviations from long-run equilibrium are corrected over time. This indicates that although sovereign spreads respond to short-run disturbances in inflation, economic policy uncertainty, and banking-sector conditions, they remain tied to a long-run equilibrium relationship with these fundamentals. Such a result is consistent with the idea that monetary and financial shocks may generate temporary deviations, but persistent macroeconomic relationships continue to determine long-run pricing. Studies examining conventional and unconventional monetary policy under uncertainty have similarly emphasized the coexistence of short-run stabilization effects and longer-run macroeconomic adjustments [12]. Likewise, evidence on capital-market fluctuations suggests that the effects of monetary policy unfold dynamically rather than instantaneously [1]. The existence of a stable long-run relationship therefore reinforces the necessity of combining short-run policy intervention with structural measures addressing inflation, fiscal imbalance, banking fragility, and policy uncertainty.

The international literature on financial risk provides additional support for the broader interpretation of the findings. Recent research demonstrates that sovereign credit risk increasingly incorporates multiple forms of uncertainty beyond conventional default fundamentals. Climate uncertainty, for example, has been shown to generate spillovers across sovereign credit markets [15], while natural disasters can increase sovereign credit risk

and produce contagion effects across countries [16]. Cross-country evidence also shows that climate-related risk can undermine financial stability through macroeconomic and financial channels [14]. Although these studies concern environmental rather than inflationary or monetary shocks, they confirm that sovereign risk premia are sensitive to multidimensional sources of uncertainty. This supports the present study's interpretation that Iran's sovereign spread reflects not merely default probability but a broader set of risks associated with purchasing-power erosion, liquidity stress, policy credibility, banking-sector vulnerabilities, and macroeconomic instability.

The broader bond-pricing literature also reinforces the argument that spreads incorporate information about nontraditional risks. Evidence on the sovereign greenium shows that government securities can display yield differentials related to environmental characteristics [17], while climate risk has been shown to affect green bond pricing [18]. Furthermore, research questioning whether government issuers are systematically "greener" demonstrates that sovereign status does not eliminate differential risk pricing [19]. These findings are conceptually relevant because they illustrate that bond spreads incorporate risks extending beyond conventional credit metrics. In the present context, inflation uncertainty, exchange-rate depreciation, unconventional monetary intervention, and fiscal-banking linkages may similarly be capitalized into the sovereign spread.

The role of unconventional monetary policy should therefore be interpreted cautiously. Such policies can reduce short-term financing stress and compress yields, but their longer-run effectiveness depends on whether they stabilize expectations and address the structural sources of risk. Evidence from the euro area shows that forward guidance and unconventional monetary policy can compress the yield curve and influence forward rates [8]. However, prolonged ultra-accommodative policy may also create financial distortions and encourage excessive risk-taking [9]. In addition, contemporary central banking research emphasizes that credit policy and balance-sheet interventions face important structural limitations when broader institutional problems remain unresolved [20]. Global evidence that central bank policies can affect bond issuance and the composition of financing further demonstrates that monetary interventions influence not only aggregate liquidity but also market structure and asset pricing [21]. In Iran, the implication is that unconventional monetary support may temporarily moderate financing pressures, but if it increases inflation expectations or signals fiscal dependence, its eventual effect on sovereign spreads may become adverse.

Finally, the study's results fit within a broader understanding of systemic risk as an interconnected phenomenon. Modern financial systems transmit shocks through banks, public-sector balance sheets, financial markets, and expectations. Climate-related financial-risk mapping demonstrates how shocks can propagate through interconnected balance sheets and generate systemic effects [24], while natural-disaster research shows that vulnerability and institutional conditions shape the distribution and magnitude of economic losses [23]. Corporate responses to major shocks may also involve changes in investment and innovation behavior [22]. Although the nature of the shocks differs from those examined here, these studies reinforce the general principle that financial risk should be analyzed as a system of interacting exposures rather than through isolated variables. The present findings similarly indicate that sovereign spread divergence in Iran emerges from the interaction of inflation, policy uncertainty, credit deterioration, liquidity stress, leverage, non-performing loans, and the shadow economy. The significant long-run coefficients and stable model diagnostics therefore support an integrated interpretation in which sovereign spread behavior reflects the cumulative macro-financial consequences of persistent structural imbalances.

The present study has several limitations that should be considered when interpreting its findings. First, the analysis relies on macroeconomic time-series data, and the accuracy of the results depends on the quality,

consistency, and temporal comparability of the underlying data. Some variables, particularly the shadow economy and economic policy uncertainty, are not directly observable and therefore require indirect estimation or index construction, which may introduce measurement error. Second, the Economic Policy Uncertainty Index was constructed using principal component analysis from a selected set of monetary, fiscal, exchange-rate, inflation, and output variables; alternative variable selections or weighting procedures could produce somewhat different estimates. Third, the NARDL framework captures nonlinear asymmetry between positive and negative shocks but does not explicitly model structural regime changes, abrupt political-economic disruptions, or time-varying coefficients. Fourth, sovereign spread measurement in the Iranian context is constrained by the structure and depth of domestic financial markets, and the selected proxy may not capture every dimension of sovereign risk. Finally, the study focuses on aggregate national data and therefore cannot identify heterogeneous responses across individual banks, financial institutions, investor groups, or sectors of the economy.

Future studies could extend the present analysis by employing alternative nonlinear and time-varying approaches, such as threshold vector error-correction models, Markov-switching models, quantile ARDL models, or time-varying parameter specifications, to determine whether the estimated relationships differ across inflationary, recessionary, or financial-stress regimes. Researchers could also construct alternative measures of economic policy uncertainty using textual, news-based, or expectation-based information and compare their explanatory power with the PCA-based index used in this study. Further studies could investigate the interaction between sovereign spreads and exchange-rate expectations, sanctions intensity, oil-price shocks, capital-market volatility, fiscal deficits, and central bank balance-sheet expansion. Bank-level panel data could also be employed to examine whether credit risk, liquidity risk, leverage, and non-performing loans affect sovereign risk differently across public and private banks. Comparative studies involving Iran and other inflation-prone emerging economies would additionally help determine whether the asymmetric relationships observed in this study are country specific or represent a broader pattern in economies characterized by fiscal dominance and unconventional monetary intervention.

Policymakers should prioritize coordinated inflation control, reduction of economic policy uncertainty, and improvement of fiscal and monetary credibility rather than relying predominantly on short-term liquidity interventions. Clearer policy communication, greater consistency between monetary and fiscal objectives, and reduced reliance on banking-system financing of fiscal imbalances can help moderate uncertainty premiums. Banking-sector supervision should place particular emphasis on controlling credit risk, liquidity mismatches, excessive leverage, and the accumulation of non-performing loans because deterioration in these variables appears to translate directly into higher sovereign risk. Measures aimed at expanding the formal economy and strengthening the tax base may also reduce fiscal dependence on inflationary financing and improve the transparency of macroeconomic conditions. Because adverse shocks were found to have asymmetric effects, preventive policy is likely to be more effective than attempting to reverse risk after deterioration has occurred. A comprehensive strategy combining price stability, stronger bank balance sheets, transparent fiscal management, and predictable monetary policy can therefore contribute to reducing sovereign spread volatility and improving long-run financial stability.

Authors' Contributions

Authors equally contributed to this article.

Ethical Considerations

All procedures performed in this study were under the ethical standards.

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Conflict of Interest

The authors report no conflict of interest.

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