

# The Effects of Political Influence, Institutional Quality, and Central Bank Intervention on Banking Stability During Recession and Expansion: A Markov-Switching Approach

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**Citation:** Oveisiforodoei, N., Daghighiasli, A., Afsharirad, M., & Aminrashti, N. The Effects of Political Influence, Institutional Quality, and Central Bank Intervention on Banking Stability During Recession and Expansion: A Markov-Switching Approach. *Business, Marketing, and Finance Open*, 3(4), 1-19.

**Abstract:** The present study examines the effects of political influence, institutional quality, and central bank intervention on banking stability during periods of recession and expansion using a Markov-switching model. In terms of purpose, the study is applied research, while in terms of methodological nature, it is descriptive-analytical and falls within the category of ex post facto research. The study covers Iran over the period 1991–2023. The results of the Markov model estimation indicate that most coefficients are statistically significant at the 95% confidence level. According to the estimation results, the intercept coefficient is  $-0.72$  in Regime 1 and  $0.14$  in Regime 2. The regime with a negative intercept represents the recession regime, whereas the regime with a positive intercept represents the expansion regime. Furthermore, the variance of the disturbance term is  $5.99$  in Regime 1 (recession) and  $0.10$  in Regime 2 (expansion). These values indicate that Regime 1, corresponding to the recessionary period, exhibits greater volatility than Regime 2, corresponding to the expansionary period. The model estimation results further show that the coefficients of foreign-exchange revenues derived from oil exports and political stability are positive during both recession and expansion. Moreover, central bank intervention has a negative effect on banking stability during recession and a positive effect on banking stability during expansion, both at the 95% confidence level.

**Keywords:** political influence, institutional quality, central bank intervention, banking stability, Markov-switching regime model

Received: 03 March 2026

Revised: 07 July 2026

Accepted: 15 July 2026

Published: 30 July 2026



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## 1. Introduction

Banking stability is a fundamental prerequisite for the effective functioning of modern economies because banks constitute a central mechanism through which savings are mobilized, credit is allocated, liquidity is created, and monetary policy is transmitted to households and firms. A stable banking system is expected to maintain its intermediation functions under ordinary economic conditions while retaining sufficient resilience to withstand adverse macroeconomic and financial shocks.

Conversely, deterioration in banking stability can intensify credit constraints, increase the probability of bank distress, disrupt the allocation of financial resources, and amplify fluctuations in output and investment. The close interaction between financial systems and economic cycles is therefore of particular importance, as financial-sector conditions can both reflect and reinforce changes in aggregate economic activity [1]. The consequences of banking

instability may become especially severe during recessionary periods, when declining asset quality, reduced profitability, increasing non-performing loans, and heightened uncertainty place simultaneous pressure on banks' balance sheets. The relationship between financial instability and macroeconomic performance is consequently dynamic rather than static, and evidence from Middle Eastern economies confirms that financial instability must be understood within a broader system of interactions among financial, institutional, and macroeconomic variables [2]. This perspective implies that the determinants of banking stability may operate differently across expansionary and recessionary phases and that models assuming constant relationships throughout the economic cycle may fail to capture important nonlinearities.

Business cycles provide an especially relevant framework for examining these nonlinearities. During periods of economic expansion, increasing income, investment, asset prices, and credit demand may improve bank profitability and borrowers' repayment capacity; however, rapid credit expansion may simultaneously encourage risk accumulation that becomes visible only when economic conditions deteriorate. During recessions, declining output and income can weaken debt-servicing capacity, increase credit risk, reduce collateral values, and adversely affect bank balance sheets. Evidence from Iran demonstrates that business-cycle fluctuations are associated with important changes in banking outcomes. Research on Iranian banks has shown that the occurrence of business cycles is related to the behavior of non-performing loans [3], while differences in profitability between public and private banks are also influenced by cyclical economic conditions [4]. More broadly, macroeconomic variables have been found to exert regime-dependent effects on Iranian business cycles when examined through Markov-switching models [5]. These findings reinforce the argument that the determinants of banking stability should not necessarily be expected to have uniform effects across all macroeconomic conditions. Instead, recession and expansion can constitute distinct regimes in which banks, policymakers, firms, and households respond differently to economic incentives and financial constraints.

One major determinant of banking stability is the quality of the institutional environment within which financial institutions operate. Institutional quality encompasses the effectiveness of legal frameworks, protection of property rights, regulatory quality, monetary credibility, market-supporting rules, and the capacity of governance arrangements to reduce uncertainty and constrain opportunistic behavior. Strong institutions can facilitate contract enforcement, improve information quality, reduce transaction costs, strengthen financial supervision, and increase confidence in economic and financial relationships. Recent cross-country evidence from the Asia-Pacific region indicates that institutional quality produces significant long-term consequences for the development and accessibility of financial systems, particularly through its relationship with financial inclusion [6]. Evidence from European Union countries likewise shows that banking-system stability is influenced by a combination of structural, financial, and institutional factors and that the resilience of banking systems cannot be separated from the broader quality of their operating environments [7]. Accordingly, institutional quality should be treated not merely as a background condition but as an active determinant of the capacity of banks to absorb shocks and preserve financial intermediation.

The importance of institutions becomes even more apparent in economies characterized by political uncertainty, regulatory inconsistency, or recurring external and internal shocks. Political institutions influence expectations regarding economic policy, regulatory enforcement, fiscal discipline, property rights, and the predictability of government actions. In this context, political stability represents an important dimension of political influence because it affects investors' expectations and the credibility of economic policymaking. Research examining the Iranian economy has demonstrated that political stability is significantly related to economic growth, although the

magnitude of this relationship may differ across the conditional distribution of growth outcomes [8]. International evidence also indicates that political stability can condition the relationship between political institutions and economic performance, suggesting that macroeconomic outcomes depend not only on formal institutional arrangements but also on the stability within which those arrangements operate [9]. From a political-economy perspective, governance quality and the organization of government policies can further shape economic performance by determining the consistency, accountability, and effectiveness of policy implementation [10]. More generally, the experience of major systemic disruptions has highlighted the importance of governance capacity and institutional coordination in protecting economic and human security under conditions of uncertainty [11]. These considerations suggest that political stability and institutional quality may substantially influence banking resilience, particularly when the economy is exposed to recessionary shocks.

Empirical evidence from the Middle East and North Africa further demonstrates the close relationship between institutional conditions and banking stability. Elfeituri showed that banking stability in MENA economies is associated with institutional quality, market concentration, competition, and political conflict, indicating that banking-system resilience is shaped by both market structure and the broader institutional-political environment [12]. This relationship may also be nonlinear. Ghazouani and Basti found that the interactions among bank stability, competition, and intervention quality in North African countries cannot be adequately characterized as purely linear relationships [13]. Such findings are particularly relevant for economies in which regulatory interventions, political conditions, and financial-market structures change substantially over time. They suggest that the marginal contribution of institutional quality or political stability to banking stability may depend on prevailing economic conditions. A favorable institutional framework may, for example, provide greater stabilizing benefits during periods of economic stress, while weak institutions can magnify uncertainty and financial fragility precisely when banks face the greatest balance-sheet pressures.

Central bank intervention represents another critical dimension of banking stability, particularly in economies exposed to substantial exchange-rate fluctuations and foreign-exchange market pressure. Central banks can influence banking-system conditions through monetary policy, reserve requirements, liquidity facilities, foreign-exchange intervention, and other instruments intended to stabilize financial and macroeconomic conditions. Intervention in the foreign-exchange market can be particularly important where exchange-rate volatility affects inflation expectations, banks' foreign-currency exposures, import costs, liquidity conditions, and the value of financial assets and liabilities. In Iran, central bank intervention has been shown to have significant implications for commercial-bank profitability, and nonlinear approaches suggest that the effects of intervention can depend on prevailing economic conditions [14]. Similarly, evidence concerning the interaction between financial risk, exchange market pressure, central bank policy intervention, and banking stability demonstrates that the effects are nonlinear rather than constant across different states of the economy [15]. These studies provide an important basis for investigating whether central bank intervention exerts asymmetric effects on banking stability during recession and expansion.

The direction of the effect of central bank intervention is theoretically ambiguous. On the one hand, timely intervention can alleviate liquidity shortages, reduce excessive exchange-rate fluctuations, stabilize expectations, and prevent temporary market disruptions from developing into systemic financial distress. Precautionary liquidity provision may be particularly important when banks increase excess reserves in response to uncertainty and when liquidity shocks interact with business-cycle conditions [16]. On the other hand, frequent or poorly calibrated intervention may distort market signals, generate moral hazard, weaken incentives for prudent risk

management, or create uncertainty concerning future policy. These considerations are especially important in a banking system in which capital adequacy and the capacity to absorb losses under crisis conditions are central determinants of financial stability [17]. The net effect of intervention may therefore differ substantially across economic regimes: measures that support stability during an expansion may have weaker or even adverse effects during a recession if they interact with severe balance-sheet constraints, inflationary pressures, exchange-rate instability, or declining confidence.

Liquidity conditions and monetary developments also form an important link between central bank intervention, macroeconomic fluctuations, and banking-system resilience. Financial-market uncertainty can influence real macroeconomic variables through changes in investment, consumption, asset valuation, and financing conditions, creating feedback loops between the financial sector and the real economy [18]. Economic uncertainty can also affect the informational environment of capital markets; evidence from companies listed on the Tehran Stock Exchange indicates that economic uncertainty is associated with changes in stock-return synchronicity [19]. Policy uncertainty is likewise relevant to emerging forms of financial assets and accounting environments, illustrating the broader implications of uncertain policymaking for financial decisions and asset valuation [20]. Although these strands of literature examine different dimensions of financial behavior, collectively they indicate that uncertainty can alter market participants' responses, influence the transmission of monetary policy, and intensify the sensitivity of financial institutions to macroeconomic disturbances. The resulting effects may be particularly pronounced when inflation, exchange-rate pressures, and liquidity growth occur simultaneously.

Inflation represents another important macroeconomic determinant of banking stability. High and volatile inflation can reduce the real value of financial contracts, increase uncertainty, distort interest-rate signals, and weaken the predictability of borrowers' and lenders' cash flows. In banking systems where nominal interest rates are subject to regulatory or institutional constraints, inflation may create additional pressure on profitability and resource allocation. Inflation can also interact with exchange-rate depreciation, monetary expansion, and central bank intervention, making it difficult to isolate its effect through a simple linear specification. The broader literature on macro-financial risk shows that financial conditions contain important information about the distribution of future economic growth, particularly downside growth risks [21]. This approach highlights the asymmetric nature of financial vulnerability: adverse financial conditions may have disproportionately large effects on economic activity during periods of stress. Such asymmetry provides a theoretical rationale for evaluating banking stability across distinct regimes rather than assuming that macroeconomic determinants exert identical effects throughout the entire sample period.

Bank-specific and systemic financial risks may also change considerably across economic regimes. Liquidity risk and credit risk, in particular, can produce threshold effects on banking stability. Evidence from the MENA region demonstrates that the relationship between these risks and bank stability depends on whether risk indicators exceed particular threshold levels [22]. This finding is consistent with the broader argument that financial systems may exhibit nonlinear responses once vulnerabilities accumulate beyond certain levels. In addition, variables that initially appear to belong to separate markets can contribute to macro-financial transmission through wealth, collateral, and expectations channels. For example, research on the determinants of housing prices illustrates the sensitivity of asset markets to macroeconomic and financial variables [23]. Such asset-price movements can influence banking stability indirectly by changing collateral values, household balance sheets, credit demand, and

the quality of banks' loan portfolios. Accordingly, the stability of the banking sector needs to be analyzed within an interconnected macro-financial system rather than as an isolated institutional outcome.

From an empirical perspective, these relationships also create substantial methodological challenges. Banking stability, central bank intervention, institutional conditions, monetary variables, and macroeconomic performance may be simultaneously determined, raising the possibility of endogeneity and reverse causality. Methodological research on financial econometrics emphasizes that ignoring endogeneity can lead to biased estimates and misleading conclusions concerning causal relationships [24]. In addition, when relationships change across recessionary and expansionary states, conventional linear models that impose constant coefficients may obscure economically meaningful differences between regimes. Regime-switching approaches provide an appropriate framework for identifying such changes because they allow model parameters and disturbance variances to vary according to unobserved states. This feature is particularly relevant to Iran, where the economy has experienced substantial fluctuations in growth, inflation, liquidity, exchange rates, political conditions, and external revenues over time. A Markov-switching framework can therefore distinguish between recessionary and expansionary regimes endogenously and estimate whether the effects of explanatory variables differ across those states.

Oil-derived foreign-exchange revenues constitute an additional structural factor that is particularly important in the Iranian context. Variations in oil revenues can affect government expenditure, foreign-exchange availability, monetary conditions, aggregate demand, exchange rates, and ultimately the financial position of banks. Changes in oil revenues may therefore interact with central bank foreign-exchange interventions and domestic liquidity creation, producing different effects during economic expansions and contractions. The financial system's response to these changes is unlikely to be independent of institutional quality and political stability, because stronger institutions may improve the management of resource revenues and reduce policy uncertainty, whereas weak governance may amplify volatility. The literature on financial instability in Middle Eastern countries reinforces the importance of evaluating such interdependencies within dynamic frameworks [2]. At the same time, evidence concerning institutional quality, political conflict, and bank stability in the region emphasizes that macroeconomic resources alone cannot guarantee banking resilience when the surrounding institutional and political environment is fragile [12].

Despite the expanding literature on banking stability, several gaps remain. A considerable number of studies examine individual determinants such as political stability, institutional quality, central bank intervention, liquidity risk, financial uncertainty, or business cycles separately, while relatively few studies integrate these factors within a unified regime-dependent framework. Moreover, studies focusing specifically on Iran have provided valuable evidence concerning business cycles [5], bank profitability [4], non-performing loans [3], banking-system competition and crisis resilience [17], political stability [8], and central bank intervention [14, 15], but the combined influence of political conditions, institutional quality, intervention policy, monetary developments, inflation, and oil-related foreign-exchange revenues on banking stability across recessionary and expansionary regimes remains insufficiently examined. International evidence concerning institutional quality [6, 7], nonlinear banking relationships [13], financial risk thresholds [22], precautionary liquidity behavior [16], uncertainty [19, 20], governance [10, 11], financial-market effects [18, 21], and asset-market determinants [23] further demonstrates the need for an integrated analytical framework capable of accounting for nonlinear dynamics.

Accordingly, analyzing banking stability through a Markov-switching specification can contribute to both the banking-management and macroeconomic-policy literature by distinguishing the effects of explanatory variables according to the prevailing state of the economy. Such an approach permits the identification of whether central

bank intervention stabilizes or destabilizes the banking system under different macroeconomic conditions, whether institutional quality continues to exert a protective influence during downturns, and whether political stability becomes more important when the economy is exposed to heightened uncertainty. It can also clarify how monetary expansion, inflation, and oil-derived foreign-exchange revenues interact with banking stability when the economic system moves between recession and expansion. These distinctions have practical implications for central bank decision-making, banking-sector risk management, financial regulation, and institutional reform because policies that are effective in one regime may be ineffective or counterproductive in another.

Therefore, the aim of the present study is to examine the effects of political influence, institutional quality, central bank intervention, oil-derived foreign-exchange revenues, liquidity, and inflation on banking stability in Iran during recessionary and expansionary regimes using a Markov-switching model.

## 2. Methodology

The central bank intervention index was first extracted and measured.

The Weymark (1995) model was used to calculate the level of Exchange Market Pressure (EMP) and the degree of central bank intervention, as discussed in the theoretical foundations. The level of exchange market pressure and the degree of central bank intervention in the foreign-exchange market can be calculated by evaluating these measures. The relevant values are presented in the second and third series of Figure 1. The degree of central bank participation depends on the type of exchange-rate regime in operation and the country's economic conditions. Under a pure floating exchange-rate system, the central bank does not intervene. Because the intervention rate is zero,  $I_t = 0$ , changes in the central bank's foreign-exchange reserves may likewise be taken as zero,  $\Delta r_t = 0$ . By contrast, under a fixed exchange-rate system, the intervention rate equals one. Since changes in the exchange rate are zero under this system:

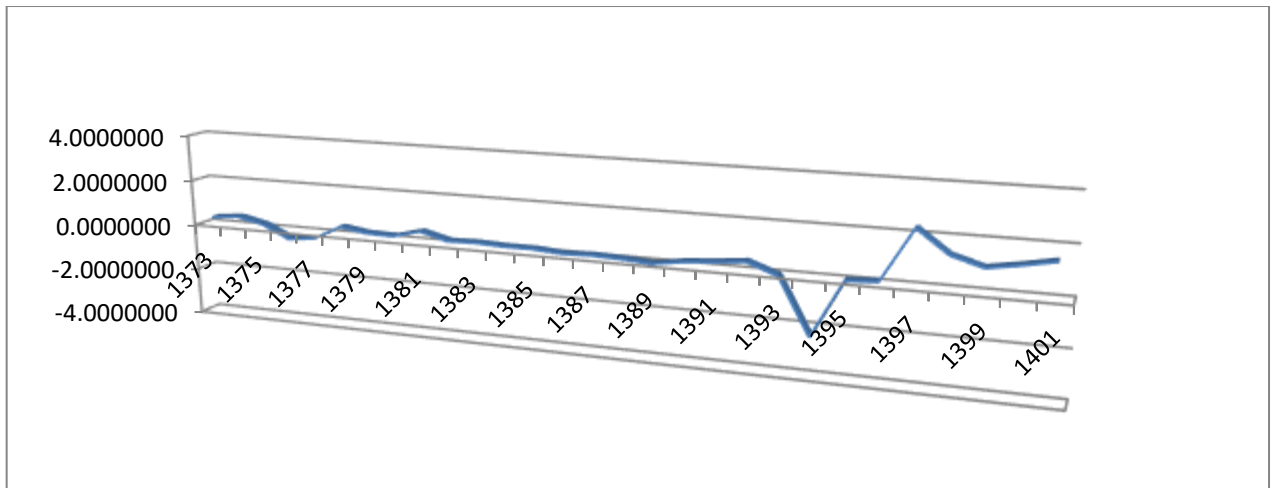
$$\Delta e_t = 0 \Rightarrow EMP_t = \eta \Delta r_t, \quad I_t = \frac{\eta \Delta r_t}{EMP_t} = 1 \quad (1)$$

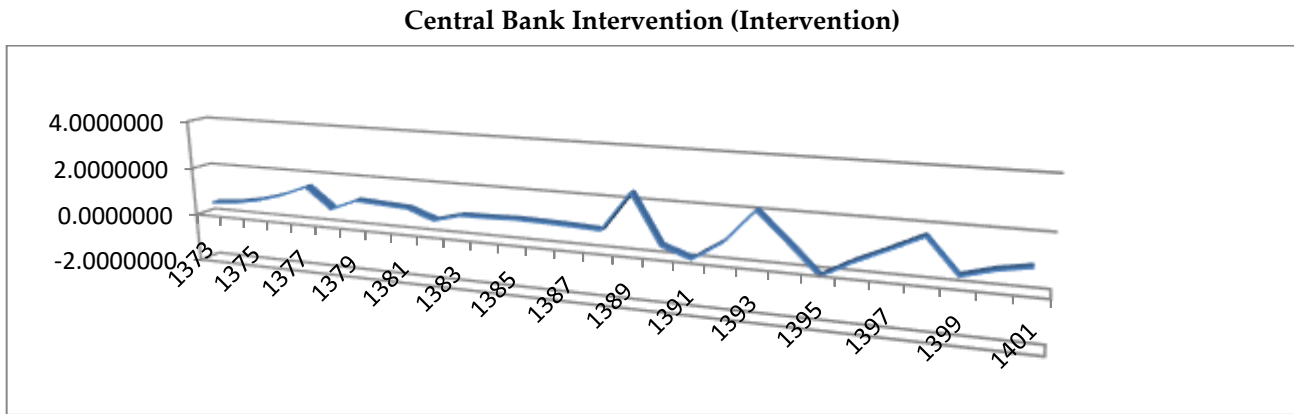
Under a managed floating exchange-rate regime, central bank behavior lies between these two limiting cases, and the degree of intervention can be expressed as follows:

$$0 < I_t < 1 \quad (2)$$

Irrespective of whether  $EMP_t$  is positive or negative, a negative value of  $I_t$  indicates a leaning-with-the-wind intervention policy, whereas a positive value indicates a leaning-against-the-wind intervention policy.

Exchange Market Pressure (EMP)





**Figure 1. Trend of the Central Bank Intervention Index and Exchange Market Pressure**

According to the results presented in Figure 1, EMP was positive in 23 years. In other words, during 23 years, Iran's foreign-exchange market experienced increasing exchange market pressure, or pressure toward depreciation of the Iranian rial, which also contributed to increased inflation. Thus, evidence of speculative attacks can be identified during these 23 years. The degree of central bank intervention is reported in the second series of Figure 1. The average degree of intervention was 0.81. In other words, over the period 1991–2023, central bank intervention activities offset, on average, approximately 19% of exchange market pressure. The estimated values of the degree of central bank intervention indicate that  $I_t > 0$  in 27 years. Therefore, during most of the years under investigation—27 out of 28 years—the central bank implemented a leaning-against-the-wind intervention policy. Moreover, in years when the degree of intervention exceeded one and EMP was negative—five years—policymakers pursued a depreciation of the rial, and changes in foreign reserves exceeded the increase in domestic money demand. In years when the degree of intervention exceeded one and EMP was positive—four years—the central bank pursued a policy aimed at appreciating the rial, and changes in foreign reserves were smaller than the increase in domestic money demand. However, intervention was negative in one year. In that year, the degree of intervention was negative while EMP was positive; therefore, the central bank's policy may be interpreted as pursuing currency depreciation during a period of excess money supply or under similar circumstances. In other words, intervention was leaning with the wind. Figure 1 also shows that the highest degree of central bank intervention occurred during 2010–2011, with a threshold value of 2.36, attributable to the widening gap between official and unofficial exchange rates and the deviation of the exchange rate from its equilibrium level.

The stationarity and cointegration properties of the variables included in the research model were subsequently examined. If time series exhibit coordinated movement, 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 referred to as cointegrated. In simple terms, cointegration occurs when two time series broadly move along the same long-run trajectory. Economic analysis generally assumes that variables included in an economic theory are linked through long-run equilibrium relationships. In applied econometric analysis, estimation of long-run relationships generally assumes that the means and variances of the variables are constant over time and independent of the time factor, thereby implicitly assuming behavioral stability. However, empirical research has shown that, in many cases, time-series variables do not exhibit such behavioral stability.

Consequently, conventional  $F$ - and  $t$ -tests derived from estimation procedures in which behavioral stability or stationarity is not satisfied are invalid and may produce misleading results. This problem is commonly referred to

as **spurious regression**. Therefore, to ensure the validity of empirical findings, researchers have revised conventional estimation procedures and systematically examine the stationarity of variables and cointegration among them.

Cointegration analysis has been recognized 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 the individual variables themselves may be nonstationary and may exhibit upward or downward movements over time. Their joint movement may nevertheless generate a stable linear relationship in the long run and thus establish an equilibrium relationship among them. Conversely, when no such long-run linear relationship exists, the variables are regarded as noncointegrated.

In general, cointegration analysis provides a method for estimating long-run equilibrium parameters in relationships involving nonstationary variables. It therefore constitutes an important methodological framework for specifying, estimating, and testing dynamic models and can consequently be employed to examine the empirical validity of economic theories. Moreover, cointegration analysis can be used to estimate short-run or disequilibrium parameters. The long-run parameters obtained from cointegration analysis can be used to estimate such short-run parameters within **error-correction models**. Error-correction models essentially link the short-run fluctuations of variables to their long-run equilibrium values.

### Nonstationary Time Series

By definition, a time series such as  $X_t$  is said to be **strictly stationary** 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 of  $X_t$ , and consequently its moments—including its mean and variance—are independent of time.

In practice, however, testing this condition is complicated. Moreover, time-series analysis generally does not require the strong-stationarity assumption. Instead, the more relevant concept is **weak stationarity**.

A time series  $X_t$  is said to be weakly stationary if the following conditions hold:

$$E(X_t) = \mu \quad (3)$$

$$\text{Var}(X_t) = \gamma_0 < \infty \quad (4)$$

$$\text{Cov}(X_t, X_{t-k}) = \gamma_k \quad (5)$$

As can be observed, the mean of the time series  $X_t$  must be constant, and its autocovariance function for any given value of  $k$  must not depend on  $t$ . For this reason, a weakly stationary time series is also referred to as **covariance stationary**. It should be noted that  $\gamma_0$ , corresponding to  $k = 0$ , represents the variance of  $X_t$ , which is also independent of time. Hereafter, the term stationarity refers to weak stationarity.

The importance of stationarity in economics arises from the fact that a stationary series does not preserve the effect of a shock indefinitely. If the series is affected by a shock, the effect is not permanent and eventually dissipates.

This property is particularly important in cointegration analysis.

Nonstationary time series may arise in two ways. In the first case, the series  $X_t$  is a function of the time variable  $t$ :

$$X_t = f(t) + u_t \quad (6)$$

where  $u_t$  is a stationary disturbance term. If  $f(t)$  is a first-degree polynomial in time, then:

$$X_t = \alpha + \beta t + u_t \quad (7)$$

The mean of  $X_t$  is a function of time and is therefore nonstationary. In the second case,  $X_t$  is generated according to the following model:

$$X_t = \alpha + X_{t-1} + u_t \quad (8)$$

where  $u_t$  is a stationary series with zero mean and constant variance. The above specification is known as a **random-walk process**. Through successive substitution, Equation (8) may be rewritten, assuming that  $X_0$  is known, as:

$$X_t = X_0 + \alpha t + \sum_{i=1}^t u_i \quad (9)$$

In this case, the variance of  $X_t$  is not independent of time. Nevertheless, both Equations (7) and (8) exhibit a linear trend. In Equation (7),  $u_t$  represents a transitory shock, whereas in Equation (8), it represents a permanent shock.

If an economic variable is affected by permanent shocks, the specification in Equation (8) is appropriate. Equations (7) and (8) are respectively referred to as the **deterministic-trend model** and the **stochastic-trend model**. In both models,  $X_t$  increases at the trend rate  $\beta$ ;  $X_t$  may also be expressed in logarithmic form, in which case  $\beta$  represents its growth rate. In the first model, shocks are transitory, whereas in the second model, they are permanent. The constant term in the second model is commonly referred to as the **drift term**. Unlike stochastic trends, deterministic trends do not create fundamental problems for valid statistical inference and cointegration analysis. Their effects can be controlled by including the time-trend variable  $t$  in the model. Accordingly, cointegration analysis is primarily concerned with stochastic trends.

For certain purposes, researchers may wish to detrend a time series. For Equation (7), this can be achieved by regressing  $X_t$  on time, in which case the resulting residuals are trend-free. Alternatively, differencing can be used to obtain a stationary series:

$$X_t - \beta t = \alpha + u_t \quad (10)$$

Likewise, in Equation (8), first differencing can produce a stationary series with constant mean:

$$\Delta X_t = \alpha + u_t \quad (11)$$

The specifications represented by Equations (10) and (11) are respectively referred to as **trend-stationary** and **difference-stationary** processes. This terminology was introduced by Nelson and Plosser.

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

For this purpose, the following general equations may be considered:

$$X_t = \rho X_{t-1} + u_t \quad (12)$$

and

$$\Delta X_t = (\rho - 1)X_{t-1} + u_t \quad (13)$$

where  $u_t$  is independently and identically distributed.

To examine stationarity, the null hypothesis  $H_0$ , representing nonstationarity, and the alternative hypothesis  $H_1$  are formulated as follows:

$$H_0: \rho = 1 \quad (14)$$

$$H_1: |\rho| < 1 \quad (15)$$

If  $\rho = 1$  in the equations above, the series is nonstationary and is said to contain a **unit root**. If  $|\rho| < 1$ , the hypothesis of nonstationarity is rejected, and the series is said to be integrated of order zero:

$$X_t \sim I(0) \quad (16)$$

If a series is nonstationary, it may be rendered stationary through differencing. If it becomes stationary after first differencing, it is said to be integrated of order one:

$$X_t \sim I(1) \quad (17)$$

More generally, if the series becomes stationary after being differenced  $d$  times, it is said to be integrated of order  $d$ :

$$X_t \sim I(d) \quad (18)$$

If  $k$  time series,  $X_{1t}, X_{2t}, \dots, X_{kt}$ , are all integrated of order  $d$ , and a linear relationship exists among them such that:

$$Z_t = \beta_1 X_{1t} + \beta_2 X_{2t} + \dots + \beta_k X_{kt} \quad (19)$$

and

$$Z_t \sim I(d - b), \quad b > 0, \quad (20)$$

then the variables are said to be cointegrated of order  $(d, b)$ , which may be expressed as:

$$X_t \sim CI(d, b) \quad (21)$$

If several time series are integrated of different orders, their linear combination will generally be integrated of the highest order among them. In the present study, the conventional Phillips–Perron (PP) unit-root test was employed.

**Table 1. Results of the Phillips–Perron Unit-Root Test for the Model Variables**

| Variable | PP Statistic | p-value | Variable Status |
|----------|--------------|---------|-----------------|
| BS       | -2.307357    | 0.1759  | Nonstationary   |
| D(BS)    | -5.992881    | 0.0000  | I(1)            |
| CPI      | -1.071349    | 0.7148  | Nonstationary   |
| D(CPI)   | -5.763331    | 0.0000  | I(1)            |
| EXGO     | -0.914903    | 0.7962  | Nonstationary   |
| D(EXGO)  | -19.91374    | 0.0001  | I(1)            |
| INTE     | -0.281572    | 0.5766  | Nonstationary   |
| D(INTE)  | -7.645239    | 0.0000  | I(1)            |
| M        | -1.985449    | 0.1970  | Nonstationary   |
| D(M)     | -8.451920    | 0.0000  | I(1)            |
| PS       | -2.709227    | 0.2397  | Nonstationary   |
| D(PS)    | -5.917927    | 0.0002  | I(1)            |
| RL       | -1.139346    | 0.2262  | Nonstationary   |
| D(RL)    | -11.39230    | 0.0000  | I(1)            |

Considering the theoretical foundations of stationarity testing, the null hypothesis  $H_0$  in these tests specifies that the variable is nonstationary. Based on the test results, all variables included in the model become stationary after first differencing and are therefore integrated of order one,  $I(1)$ .

### 3. Findings and Results

Because the model variables have the same order of integration,  $I(1)$ , a cointegration test was employed to determine whether a long-run equilibrium relationship exists among them. The Johansen–Juselius procedure was used for this purpose. Application of this procedure requires determining the number of cointegrating vectors. To interpret the cointegration results, an appropriate specification must be selected regarding the presence or absence of a deterministic time trend and intercept in the cointegrating vector. Five specifications are generally considered: Model 1 contains neither an intercept nor a time trend; Model 2 includes a restricted intercept but no time trend; Model 3 includes an unrestricted intercept but no time trend; Model 4 includes an unrestricted intercept and a restricted time trend; and Model 5 includes both an unrestricted intercept and an unrestricted time trend. These five specifications range from the most restrictive, Model 1, to the least restrictive, Model 5. The null hypothesis of no cointegrating vector is initially tested against the alternative hypothesis of one cointegrating vector. Subsequently, the hypothesis of at most one cointegrating vector is tested against the alternative of two

cointegrating vectors. This procedure continues until  $n - 1$  cointegrating vectors are considered, where  $n$  denotes the number of variables. A summary of the trace test ( $\lambda_{\text{Trace}}$ ) and maximum-eigenvalue test ( $\lambda_{\text{Max}}$ ) results for the number of cointegrating vectors across the five specifications is presented in Table 2. As shown, the null hypothesis of no cointegrating vector is rejected against the alternative of at least one cointegrating vector. Therefore, at least one cointegrating relationship exists among the variables under investigation.

**Table 2. Summary of the Number of Cointegrating Vectors**

| Test                    | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|-------------------------|---------|---------|---------|---------|---------|
| Trace test              | 4       | 6       | 4       | 3       | 4       |
| Maximum-eigenvalue test | 2       | 3       | 2       | 1       | 2       |

The model estimation results and the associated cointegration tests are reported in Table 3. According to the trace test, the existence of four cointegrating vectors is supported, while the maximum-eigenvalue test indicates two cointegrating vectors at the 5% significance level. As Johansen notes, when the two tests produce conflicting results, the maximum-eigenvalue test may be given greater weight because it specifies a more restrictive alternative hypothesis. On the basis of the reported analysis, the existence of four cointegrating vectors among the model variables is accepted.

**Table 3. Results of the Cointegration Tests**

| $H_0$      | $H_1$   | Trace Statistic | 95% Critical Value | p-value | Maximum-Eigenvalue Statistic | 95% Critical Value | p-value |
|------------|---------|-----------------|--------------------|---------|------------------------------|--------------------|---------|
| $r = 0$    | $r = 1$ | 190.2576        | 95.75366           | 0.0000  | 79.01611                     | 40.07757           | 0.0000  |
| $r \leq 1$ | $r = 2$ | 111.2415        | 69.81889           | 0.0000  | 60.74473                     | 33.87687           | 0.0000  |
| $r \leq 2$ | $r = 3$ | 50.49676        | 47.85613           | 0.0276  | 20.02921                     | 27.58434           | 0.3391  |
| $r \leq 3$ | $r = 4$ | 30.46755        | 29.79707           | 0.0418  | 18.46238                     | 21.13162           | 0.1135  |
| $r \leq 4$ | $r = 5$ | 12.00517        | 15.49471           | 0.1567  | 10.12059                     | 14.26460           | 0.2041  |
| $r \leq 5$ | $r = 6$ | 1.884585        | 3.841466           | 0.1698  | 1.884585                     | 3.841466           | 0.1698  |
| $r \leq 6$ | $r = 7$ | 3.841466        | 4.139368           | 0.0054  | 14.26460                     | 50.52698           | 0.0644  |

**Table 4. Results of the Likelihood-Ratio (LR) Test**

| Test Statistic | Degrees of Freedom | p-value |
|----------------|--------------------|---------|
| 55.986         | 7                  | 0.000   |

According to the results presented in Table 4, the probability value is below 5%. Therefore, the null hypothesis is rejected and the alternative hypothesis is accepted, supporting the use of the nonlinear Markov-switching approach for model estimation.

**Table 5. Markov-Switching Model Estimation Results for Banking Stability**

| Variable   | Coefficient | Standard Error | t-Statistic | p-value |
|------------|-------------|----------------|-------------|---------|
| $c_1$      | -0.728923   | 0.076055       | -9.584166   | 0.0000  |
| $c_2$      | 0.144886    | 0.021574       | 6.715828    | 0.0000  |
| $\sigma_1$ | 5.990521    | 0.1533015      | 39.080000   | 0.0000  |
| $\sigma_2$ | 0.107799    | 0.020271       | 5.317791    | 0.0000  |
| BS(-1)     | 0.557607    | 0.150120       | 3.714421    | 0.0002  |
| EXGO(1)    | 0.029347    | 0.015478       | 1.896082    | 0.0585  |
| EXGO(2)    | 0.016559    | 0.007091       | 2.335214    | 0.0355  |
| PS(1)      | 0.047530    | 0.024254       | 1.959712    | 0.0506  |
| PS(2)      | 0.027657    | 0.011872       | 2.329532    | 0.0203  |
| RL(1)      | 0.322750    | 0.105102       | 3.070814    | 0.0024  |

|         |           |          |           |        |
|---------|-----------|----------|-----------|--------|
| RL(2)   | 0.076171  | 0.029950 | 2.543238  | 0.0117 |
| INTE(1) | -0.329059 | 0.136048 | -2.418698 | 0.0234 |
| INTE(2) | 0.897400  | 0.289110 | 3.104009  | 0.0041 |
| M(1)    | -0.070881 | 0.021359 | -3.318533 | 0.0010 |
| M(2)    | -0.027478 | 0.009527 | -2.884224 | 0.0184 |
| CPI(1)  | -0.021639 | 0.010429 | -2.074977 | 0.0392 |
| CPI(2)  | -0.631234 | 0.195603 | -3.227118 | 0.0019 |

The Markov-model estimation results indicate that most coefficients are statistically significant at the 95% confidence level. According to the model estimates, the intercept coefficient is  $-0.72$  in Regime 1 and  $0.14$  in Regime 2. The regime characterized by a negative intercept represents the recession regime, whereas the regime characterized by a positive intercept represents the expansion regime. Furthermore, the variance of the disturbance term is  $5.99$  in Regime 1, corresponding to recession, and  $0.10$  in Regime 2, corresponding to expansion. These figures indicate that Regime 1, or the recessionary period, exhibits greater volatility than Regime 2, or the expansionary period. The estimation results also indicate positive coefficients for foreign-exchange revenues derived from oil exports and political stability during both recession and expansion. Moreover, central bank intervention has a negative effect on banking stability during recession and a positive effect on banking stability during expansion at the 95% confidence level.

The probabilities of transition from the recession regime to the expansion regime and vice versa, together with regime persistence, are presented in Table 6. As shown by the transition probability functions of the banking stability model, if the Iranian economy is in the expansion regime at time  $t$ , despite fluctuations in the model variables, it has an approximately  $0.58$  probability of remaining in the same regime and an approximately  $0.42$  probability of switching to the recession regime because of other factors. Conversely, if the economy is in the recession regime at time  $t$ , despite fluctuations in the model variables, it has an approximately  $0.51$  probability of remaining in the recession regime at time  $t + 1$  and an approximately  $0.49$  probability of transitioning to the expansion regime as a result of other factors.

**Table 6. Transition Probabilities Between Regimes for the Banking Stability Model**

| Country | Current Regime | Regime 2 | Regime 1 |
|---------|----------------|----------|----------|
| Iran    | Regime 1       | 0.49973  | 0.50027  |
| Iran    | Regime 2       | 0.58645  | 0.41355  |

As noted in the presentation of the model, the disturbance terms of a Markov-switching model should be normally distributed and free from serial correlation and heteroskedasticity. The results of the diagnostic tests for these properties are reported below.

**Table 7. Diagnostic Test Results for the Banking Stability Model**

| Test                                    | Test Distribution | Test Statistic | p-value |
|-----------------------------------------|-------------------|----------------|---------|
| Serial-correlation test                 | $\chi^2(4)$       | 6.9320         | 0.3652  |
| Normality test                          | $\chi^2(2)$       | 2.9095         | 0.2365  |
| Homoskedasticity test                   | $F(1,10)$         | 0.00483        | 0.8596  |
| Ramsey RESET test (model specification) | $F(2,30)$         | 1.24962        | 0.7425  |

According to the results of the serial-correlation, normality, and homoskedasticity tests, the corresponding probability values exceed 5%. It can therefore be inferred that the disturbance terms do not exhibit serial correlation, departures from normality, or heteroskedasticity. Furthermore, based on the Ramsey RESET specification test, the

model has been correctly specified. Accordingly, the diagnostic results support the validity of the estimated Markov-switching model.

#### 4. Discussion and Conclusion

The findings of the present study demonstrated that banking stability in Iran is characterized by distinct recessionary and expansionary regimes and that the effects of its determinants are not uniform across these regimes. The Markov-switching estimates identified Regime 1 as the recessionary regime, with a negative intercept of approximately (-0.73), and Regime 2 as the expansionary regime, with a positive intercept of approximately (0.14). Moreover, the estimated disturbance variance was substantially higher in the recessionary regime than in the expansionary regime, indicating that banking stability is considerably more volatile during economic downturns. This finding confirms the appropriateness of a nonlinear regime-switching framework for analyzing Iran's banking system because periods of contraction and expansion generate different financial conditions, risk structures, and behavioral responses. The result is consistent with evidence that macroeconomic variables influence Iran's business cycles differently across regimes [5]. It also accords with research showing that financial systems can amplify cyclical fluctuations and that the interaction between the financial sector and the real economy varies according to prevailing macroeconomic conditions [1]. Similarly, studies of financial instability in Middle Eastern economies suggest that financial fragility is inherently dynamic and is affected by interactions among macroeconomic, financial, and institutional factors [2]. The higher volatility observed during recession is therefore plausible because weakening borrowers' balance sheets, reduced profitability, declining asset quality, and increasing uncertainty simultaneously expose banks to multiple sources of vulnerability.

The significant positive coefficient of lagged banking stability indicates persistence in banking-system conditions. In other words, the level of stability achieved in one period influences banking stability in the subsequent period. Such persistence reflects the cumulative nature of banking-sector balance sheets, capital adequacy, loan quality, liquidity positions, and risk-management practices. Banking instability is rarely confined to a single period because deterioration in loan portfolios, profitability, and capital buffers may continue to influence institutions even after the original shock has occurred. This interpretation is consistent with evidence showing that business-cycle conditions affect bank profitability and non-performing loans in Iran [3, 4]. It is also compatible with the finding that liquidity and credit risks can generate threshold effects on bank stability, implying that accumulated vulnerabilities may persist and intensify after critical levels are reached [22]. Furthermore, the importance of capital and loss-absorption capacity under crisis conditions in the Iranian banking system indicates that current financial resilience is partly determined by balance-sheet conditions inherited from previous periods [17].

The findings further showed that foreign-exchange revenues generated by oil exports had a positive coefficient in both economic regimes. The coefficient was positive but only marginally significant in the recessionary regime, whereas it was positive and statistically significant in the expansionary regime. This result can be explained by the structural importance of oil revenues to the Iranian economy. Increased oil-derived foreign-exchange revenues expand the availability of foreign currency, facilitate government financing, support aggregate demand, and can reduce pressures on the foreign-exchange market. These developments may improve firms' and households' repayment capacity and, consequently, strengthen banks' asset quality. However, the weaker statistical effect during recession suggests that increased oil revenues alone may be insufficient to neutralize the broader financial stresses associated with economic contraction. During expansion, by contrast, additional foreign-exchange

revenues may reinforce favorable macroeconomic conditions and therefore produce a clearer stabilizing effect on banks. This interpretation is consistent with studies emphasizing the role of macroeconomic and financial conditions in the emergence of financial instability [2] and with evidence that financial conditions contain important information about future economic growth and downside risk [21]. More generally, the result supports the argument that the effect of a macroeconomic resource depends on the state of the economy rather than operating through a constant linear mechanism.

Political stability also showed a positive relationship with banking stability in both regimes, although its coefficient in the recessionary regime was close to the conventional 5% significance threshold, while the effect in the expansionary regime was statistically significant. The positive relationship indicates that improvements in political stability strengthen the banking environment by reducing uncertainty regarding government policy, investment conditions, property rights, and the continuity of economic decisions. A more predictable political environment encourages investment, improves expectations, reduces risk premiums, and limits precautionary financial behavior, thereby supporting borrowers' financial positions and reducing pressure on banks. This finding agrees with evidence from Iran showing that political stability positively affects economic growth, although the magnitude of its influence may vary across economic conditions [8]. It is also aligned with international evidence indicating that political stability is an important condition through which political and institutional arrangements influence economic performance [9]. From a broader governance perspective, government effectiveness, policy consistency, and institutional accountability contribute to economic resilience [10]. The role of governance becomes especially important when economies are exposed to major shocks, as effective institutional coordination can reduce uncertainty and protect economic and human security [11]. Accordingly, political stability can be interpreted as an important supportive condition for banking resilience.

Institutional quality had a positive and statistically significant effect on banking stability in both recession and expansion, although the estimated coefficient was substantially larger during recession. This finding is particularly important because it suggests that high-quality institutions become more valuable when the banking system is exposed to adverse economic conditions. Strong legal systems, effective protection of property rights, credible monetary arrangements, efficient regulation, and predictable market rules reduce transaction costs and improve contract enforcement. During recessions, when credit risk and uncertainty increase, these institutional characteristics can limit opportunistic behavior, strengthen supervisory effectiveness, and facilitate the restructuring or enforcement of financial contracts. The positive finding is consistent with evidence indicating that institutional quality has important long-term effects on financial development and financial inclusion [6]. It also corresponds with findings from European Union countries showing that banking-system stability is shaped by institutional as well as financial and structural characteristics [7]. Similarly, evidence from MENA economies demonstrates that institutional quality is an important determinant of banking stability alongside competition, market concentration, and political conflict [12]. The stronger coefficient found during recession is therefore theoretically meaningful: when favorable economic conditions weaken, institutional safeguards may become a more important source of systemic resilience.

One of the central findings concerned the asymmetric effect of central bank intervention. Intervention exerted a statistically significant negative effect on banking stability during recession but a significant positive effect during expansion. This asymmetry indicates that the consequences of central bank intervention depend substantially on the prevailing macroeconomic regime. During expansion, intervention may stabilize exchange-rate expectations, facilitate liquidity management, reduce excessive market volatility, and support confidence in the financial system.

In such circumstances, banks generally operate under stronger balance-sheet conditions and may respond more effectively to central bank measures. This positive expansionary effect is compatible with evidence that central bank intervention can influence the profitability and operating conditions of Iranian commercial banks in a nonlinear manner [14]. It also supports the broader finding that the relationship between intervention quality and bank stability can be nonlinear [13].

The negative effect of intervention during recession, however, suggests that intervention under severe macroeconomic stress may be unable to offset underlying financial vulnerabilities and may even intensify instability. During contractions, foreign-exchange intervention can reduce reserve buffers, interact with inflationary pressures, or create uncertainty if market participants perceive interventions as unsustainable. If intervention is undertaken in an environment characterized by excess liquidity, strong depreciation expectations, or declining confidence, its short-term stabilizing benefits may be outweighed by adverse effects on expectations and balance sheets. This explanation closely corresponds with evidence that the effect of financial risk on banking stability in Iran changes in the presence of central bank intervention and foreign-exchange market pressure [15]. It also accords with the idea that precautionary liquidity behavior and excess reserves respond to cyclical conditions and can alter the transmission of monetary policy [16]. Therefore, the negative coefficient should not necessarily be interpreted as evidence that central bank intervention is intrinsically destabilizing; rather, its effectiveness appears conditional on the macroeconomic environment in which it occurs.

Liquidity had a negative and statistically significant effect on banking stability in both recessionary and expansionary regimes. This result suggests that excessive growth in liquidity may weaken banking-system stability by increasing inflationary pressures, stimulating speculative demand, intensifying exchange-rate instability, and reducing the effectiveness of monetary control. In an economy where monetary expansion is not accompanied by proportional increases in productive capacity, excess liquidity can shift toward foreign exchange, real estate, and other assets, increasing price volatility and financial uncertainty. The negative relationship is consistent with studies showing that financial-market uncertainty can affect real macroeconomic variables through complex transmission mechanisms [18]. It is also compatible with evidence that economic uncertainty changes the information environment and the behavior of financial markets in Iran [19], while policy uncertainty can influence financial-asset valuation and accounting behavior [20]. The result further corresponds with the broader literature emphasizing that liquidity risk is an important determinant of banking stability and may generate nonlinear threshold effects [22]. Hence, rapid liquidity growth should not automatically be considered supportive of banking stability, particularly if the additional liquidity contributes to macroeconomic disequilibrium rather than productive credit expansion.

Inflation also exerted a negative and statistically significant effect on banking stability in both regimes, with a considerably larger negative coefficient during expansion. Inflation reduces the real value of financial assets and liabilities, increases uncertainty, weakens the informational content of prices, and complicates the management of interest-rate and maturity mismatches. Persistent inflation can also erode depositors' confidence in domestic financial instruments and encourage portfolio shifts toward foreign currency or physical assets. For banks, such developments can increase funding instability while simultaneously reducing the quality of credit decisions. The finding is consistent with the view that adverse financial conditions contribute disproportionately to downside macroeconomic risk [21]. It also aligns with evidence that asset prices are sensitive to broader macroeconomic and financial conditions [23]. The stronger adverse effect found during expansion may reflect the possibility that

inflationary pressures accompanying rapid liquidity and credit growth undermine otherwise favorable cyclical conditions and create vulnerabilities that subsequently threaten financial stability.

The estimated transition probabilities provide additional evidence regarding the cyclical behavior of banking stability. When the economy was in the expansionary regime, the probability of remaining in that regime in the subsequent period was approximately 0.59, whereas the probability of moving toward recession was approximately 0.41. When the economy was in recession, the probability of persistence was approximately 0.50, with a roughly similar probability of transitioning toward expansion. Thus, neither regime appears permanently dominant, although expansion shows somewhat greater persistence. This relatively frequent possibility of regime transition reinforces the relevance of policy flexibility. Financial regulation and monetary intervention should not be designed on the assumption that economic conditions remain stable for long periods. The findings are consistent with Markov-switching evidence on Iranian business cycles showing meaningful transitions between distinct macroeconomic states [5]. They are also consistent with studies emphasizing nonlinear interactions among bank stability, policy intervention, and competition [13], as well as with research showing that financial instability evolves dynamically in response to changing macroeconomic conditions [2].

Finally, the diagnostic tests supported the adequacy of the estimated Markov-switching model. The tests did not indicate significant serial correlation, non-normality, heteroskedasticity, or specification error. Together with the likelihood-ratio test supporting nonlinear estimation, these results reinforce the appropriateness of distinguishing between recessionary and expansionary regimes. From a methodological standpoint, this is important because relationships among macroeconomic and financial variables may be obscured when a constant-parameter linear model is imposed on data generated under heterogeneous economic conditions. In financial research, inappropriate specification and failure to address complex empirical relationships can lead to biased or misleading conclusions [24]. The present findings therefore contribute to the literature by demonstrating that political stability, institutional quality, central bank intervention, liquidity, inflation, and oil-related foreign-exchange revenues should be interpreted conditionally according to the state of the economy rather than as determinants with fixed effects across time.

The present study has several limitations that should be considered when interpreting its findings. First, the analysis was conducted using annual aggregate data for Iran, and the relatively limited number of time-series observations restricts the ability to capture short-run fluctuations and potentially rapid changes in banking conditions. Second, banking stability was represented by an aggregate index, which may conceal substantial heterogeneity across public, private, specialized, and commercial banks. Third, some institutional and political variables are inherently multidimensional and may not be fully represented by a single composite indicator. Fourth, the Markov-switching framework classifies observations into latent regimes on the basis of statistical properties and therefore may not capture every structural event affecting the Iranian economy. Fifth, the model focused on a selected set of macroeconomic, institutional, and political determinants, while other potentially relevant variables—such as bank capitalization, credit growth, non-performing loans, fiscal deficits, exchange-rate misalignment, international sanctions, and global oil-price volatility—were not simultaneously incorporated. Finally, the observational and ex post facto nature of the study limits strong causal interpretation of the estimated relationships.

Future studies are recommended to extend the analysis using higher-frequency quarterly or monthly data where reliable observations are available, thereby allowing more precise identification of regime transitions and short-term policy effects. Researchers could also estimate bank-level panel Markov-switching models to determine whether the effects identified in the present study differ according to ownership structure, bank size, capitalization,

profitability, liquidity position, or exposure to non-performing loans. Future research may incorporate additional indicators of governance and institutional quality separately rather than relying exclusively on aggregate indices. It would also be valuable to examine the roles of exchange-rate misalignment, international sanctions, oil-price shocks, fiscal dominance, credit growth, and asset-price cycles. Alternative nonlinear methods, including threshold regression, smooth-transition models, time-varying parameter models, and nonlinear autoregressive distributed lag specifications, could be employed to test the robustness of the findings. Comparative studies involving Iran and other oil-exporting or emerging economies could further determine whether the observed regime-dependent relationships are specific to Iran or represent more general characteristics of financially and institutionally constrained economies.

The findings imply that banking and monetary authorities should adopt regime-contingent rather than uniform stabilization policies. Central bank intervention should be calibrated differently during recession and expansion, with particular caution during downturns when intervention may interact adversely with exchange-market pressure, declining confidence, and weakened bank balance sheets. Policymakers should prioritize institutional strengthening, including regulatory consistency, effective supervision, protection of property rights, transparent decision-making, and greater predictability of economic policies, because institutional quality appears especially important during recessionary periods. Monetary authorities should also place greater emphasis on controlling excessive liquidity growth and persistent inflation, as both variables are associated with weaker banking stability across regimes. Oil-derived foreign-exchange revenues should be managed in a manner that reduces procyclical monetary expansion and supports financial buffers rather than generating additional liquidity pressures. At the bank level, managers should strengthen capital and liquidity buffers during expansionary periods, conduct recession-specific stress tests, improve credit-risk monitoring, and prepare contingency plans for abrupt changes in exchange rates, inflation, and policy conditions so that favorable economic periods are used to build resilience before the next downturn.

#### **Authors' Contributions**

Authors equally contributed to this article.

#### **Ethical Considerations**

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

#### **Acknowledgments**

Authors thank all participants who participate in this study.

#### **Conflict of Interest**

The authors report no conflict of interest.

#### **Funding/Financial Support**

According to the authors, this article has no financial support.

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