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# Examining the Impact of Government Expenditure Shocks and Islamic Participation Bonds on the Financial Cycle of Business Cycles during Recession and Expansion in Iran

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## ABSTRACT

The present study examines the impact of government expenditure shocks and Islamic participation bonds on the financial cycle of business cycles during periods of recession and expansion in Iran. From the perspective of research objectives, this study is applied, and in terms of its nature, it is descriptive–analytical. It falls within the category of ex post facto research and relies on historical data covering the period from 1987 to 2024. Based on the results of estimating the Markov-switching model, most coefficients are statistically significant at the 95% confidence level, and their signs are consistent with theoretical foundations. The intercept is 0.32 in the first regime and −0.58 in the second regime. According to Hamilton (1988), a regime with a negative intercept represents a recessionary regime, while a regime with a positive intercept indicates an expansionary regime. Accordingly, in this study, the first regime represents the expansion period, and the second regime represents the recession period. The variance of the disturbance terms is 0.09 in the first regime and 0.43 in the second regime. These values indicate that the first regime (the expansion period) exhibits lower volatility compared to the second regime (the recession period) in the present study. An examination of the Markov model estimation results reported in the table above shows that negative shocks to current and capital government expenditures, consumption, and inflation exert negative effects on the financial cycle in both expansionary and recessionary periods. Shocks to current and capital government expenditures in the first regime (the expansion period) lead to a reduction in the ratio of liquidity to gross domestic product, whereas in the second regime (the recession period), this effect is positive. The production variable also has a negative effect on the ratio of liquidity to gross domestic product in both regimes (recession and expansion). In Iran, due to the fact that financial markets are generally limited and imperfect, investment depends on the ability to secure financing from domestic savings. Consequently, an increase in the rate of return on bonds enhances the feasibility of financing projects.

**Keywords:** Government expenditure, financial cycle, Islamic participation bonds, expenditure shocks, business cycle

## Introduction

Fiscal policy has long been recognized as one of the most influential macroeconomic instruments shaping business cycles, financial cycles, and long-term growth trajectories. Government expenditure, in particular, plays a dual role: it directly affects aggregate demand while simultaneously interacting with financial markets, credit conditions, and expectations. In recent decades, the growing complexity of financial systems, the globalization of


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capital flows, and the increasing prominence of uncertainty and nonlinear dynamics have intensified scholarly interest in understanding how fiscal shocks propagate through the real and financial sectors. This concern is especially salient in emerging and developing economies, where financial markets are often incomplete, fiscal space is constrained, and macroeconomic volatility is more pronounced (1, 2).

A growing body of literature demonstrates that the effects of government spending shocks are not uniform across time or states of the business cycle. Instead, fiscal shocks often generate asymmetric responses depending on whether the economy is in expansion or recession, as well as on the structure of the financial system and the degree of uncertainty prevailing at the time of the shock (3, 4). These asymmetries challenge the traditional linear frameworks that dominated earlier fiscal policy analysis and call for nonlinear and regime-dependent approaches capable of capturing state-contingent dynamics. The recognition of such nonlinearities has been reinforced by advances in econometric modeling, including nonlinear ARDL, Markov regime-switching models, and time-varying parameter frameworks (5, 6).

Parallel to these methodological developments, the concept of the financial cycle has gained prominence as a complementary lens to the traditional business cycle. While business cycles focus primarily on fluctuations in output and employment, financial cycles emphasize the joint dynamics of credit, asset prices, leverage, and liquidity over longer horizons. Empirical evidence suggests that financial cycles tend to be longer and more volatile than business cycles, and that their interaction with fiscal policy can amplify macroeconomic instability if not properly managed (2, 7). Government spending shocks, by altering public borrowing needs, liquidity conditions, and investor expectations, can significantly influence the amplitude and persistence of financial cycles, particularly in economies with shallow capital markets.

In this context, the structure of the financial system becomes a crucial moderating factor. Conventional interest-based banking systems transmit fiscal shocks through channels such as sovereign debt markets, interest rates, and credit spreads. However, in economies where Islamic finance plays a substantial role, the transmission mechanisms may differ in important ways. Islamic financial instruments, including sukuk and participation-based contracts, are grounded in principles of risk sharing and asset backing, which can alter the response of financial variables to fiscal and macroeconomic shocks (8, 9). The rapid global expansion of Islamic banking and finance over the past two decades reflects both demand-side factors, such as ethical considerations, and supply-side factors, such as financial diversification and resilience (9, 10).

Recent empirical studies indicate that Islamic financial markets are increasingly integrated with global financial systems, making them susceptible to global uncertainty, oil price shocks, and cross-border capital flows (10, 11). At the same time, their distinctive contractual structures may dampen or reshape the transmission of certain shocks, particularly those related to leverage and speculative behavior. For instance, evidence from Islamic economies suggests that fiscal shocks may have weaker contractionary effects during downturns when mediated through Islamic banking channels, owing to stronger links between financing and real economic activity (12, 13). These findings underscore the importance of explicitly incorporating Islamic financial instruments into analyses of fiscal policy and financial cycles.

Oil price dynamics further complicate this relationship, especially in resource-dependent economies. Oil price shocks influence government revenues, fiscal balances, and public spending capacity, while simultaneously affecting inflation, exchange rates, and financial conditions. The global literature documents significant spillovers from oil price volatility to trade, investment, and financial markets, with heterogeneous effects across countries and over time (11, 14). In oil-exporting economies, expansionary fiscal policy during oil booms can fuel credit growth and asset price inflation, intensifying financial cycles, whereas fiscal consolidation during oil busts can exacerbate recessions and financial stress. These dynamics highlight the need to jointly analyze government expenditure shocks, financial cycles, and external commodity shocks within an integrated framework.

Another strand of research emphasizes the role of uncertainty and news shocks in shaping macro-financial dynamics. Uncertainty shocks—arising from geopolitical events, policy ambiguity, sanctions, or financial stress—can alter the effectiveness of fiscal policy by affecting private sector expectations and risk-taking behavior (5, 15). Empirical evidence suggests that during periods of elevated uncertainty, fiscal multipliers may be smaller or more volatile, and financial markets may respond more strongly to negative than to positive shocks (7, 16). These findings reinforce the argument for asymmetric and regime-dependent modeling strategies when assessing fiscal policy impacts.

In the context of emerging economies, additional institutional factors further shape fiscal–financial interactions. Government debt sustainability, banking sector dominance, and fiscal–monetary coordination all influence how spending shocks propagate through the economy. Studies focusing on public debt and fiscal costs show that government spending shocks can have lasting implications for debt dynamics and financial stability, particularly when fiscal expansions are debt-financed (17). Moreover, in environments characterized by financial repression or limited access to external financing, domestic banking systems often become the primary conduit for financing government deficits, intensifying the feedback loop between fiscal policy and financial cycles.

Iran represents a particularly relevant case for examining these issues. As an oil-exporting economy with a dual financial system that includes both conventional and Islamic banking elements, Iran exhibits complex interactions between fiscal policy, financial cycles, and external shocks. Government expenditures play a central role in stabilizing economic activity, especially during periods of sanctions, oil revenue volatility, and financial stress (18). At the same time, the increasing use of Islamic participation instruments as a means of financing public and private investment introduces new channels through which fiscal shocks may influence liquidity, credit allocation, and financial stability.

Recent Iranian studies provide evidence of asymmetric macroeconomic responses to fiscal and policy shocks. Nonlinear ARDL and related approaches reveal that variables such as production, taxation, and financial indicators respond differently to positive and negative shocks, as well as across regimes of economic expansion and contraction (6, 19). These findings are consistent with broader international evidence emphasizing the inadequacy of linear models in capturing the true dynamics of fiscal policy effects. However, despite this growing literature, relatively limited attention has been paid to the explicit role of Islamic participation bonds and contracts in shaping the financial cycle response to government spending shocks, particularly within a regime-switching framework.

Furthermore, the interaction between fiscal shocks and financial cycles cannot be fully understood without considering the broader macroeconomic environment, including inflation dynamics, consumption behavior, and output fluctuations. Inflation can mediate the real effects of fiscal expansions by eroding purchasing power and altering real interest rates, while consumption responses reflect household expectations and credit constraints. Empirical studies suggest that these channels often operate asymmetrically, with contractionary effects being stronger during downturns than expansionary effects during booms (3, 4). Incorporating these variables into a unified empirical framework is therefore essential for a comprehensive assessment of fiscal policy effectiveness.

Against this backdrop, the present study contributes to the literature in several important ways. First, it integrates the analysis of government expenditure shocks and Islamic participation instruments within the concept of the financial cycle, rather than limiting attention to output-based business cycle measures. Second, it adopts a nonlinear, regime-dependent econometric framework to capture asymmetric responses across expansionary and recessionary phases, consistent with recent advances in macro-financial modeling (5, 7). Third, by focusing on Iran, the study provides novel evidence from an economy where fiscal policy, oil dependence, and Islamic finance intersect in particularly salient ways, complementing existing evidence from other Islamic and emerging economies (12, 13).

Moreover, this research speaks directly to policy debates concerning the design of countercyclical fiscal strategies and the use of Islamic financial instruments as stabilizing tools. Understanding whether participation-based financing amplifies or dampens financial cycles in response to government spending shocks has important implications for public debt management, financial stability, and sustainable growth. In an era marked by heightened global uncertainty, geopolitical tensions, and volatile commodity markets, such insights are increasingly valuable for policymakers seeking to balance fiscal support with financial resilience (14, 16).

In summary, the existing literature underscores the importance of asymmetric, nonlinear, and regime-dependent approaches to analyzing fiscal policy shocks, particularly in economies with distinctive financial structures and exposure to external shocks. By building on these insights and addressing identified gaps, this study aims to deepen our understanding of how government expenditure shocks and Islamic participation bonds jointly shape financial cycles across different phases of economic activity.

The aim of this study is to examine the asymmetric effects of government expenditure shocks and Islamic participation bonds on the financial cycle across expansionary and recessionary regimes in Iran using a nonlinear regime-switching framework.

## Methods and Materials

The present study is applied in terms of its objective and descriptive–analytical in terms of its nature, and it falls within the category of ex post facto research. The purpose of this article is to examine the asymmetric effects of government expenditure shocks on the financial cycle of business cycles, with an emphasis on Islamic participation bonds, using the NARDL model. The statistical population of the study covers the period from 1987 to 2024. In line with previous studies—Pragidis et al. (2017), Kim (2019), Gadenne (2019), and Pragidis et al. (2017)—this article investigates the asymmetric consequences of government expenditure shocks throughout the financial cycle during periods of recession and expansion by applying the Markov Regime Switching model. Accordingly, by employing the Markov regime-switching framework, the asymmetric effects of government expenditure shocks over the financial cycle during recessionary and expansionary periods in the country are examined.

The Markov-switching model was first introduced by Quandt (1972) and Quandt and Goldfeld (1973), and later developed by Hamilton (1989) to extract financial cycles. Unlike other nonlinear approaches such as STAR and artificial neural networks (ANN), in which transitions from one regime to another occur gradually, regime transitions in the Markov-switching model take place abruptly. In this framework, it is assumed that the regime prevailing at time $t$ is unobservable and depends on an unobserved process ($s_t$). In a two-regime model, $s_t$ can simply take the values 1 and 2. A two-regime AR(1) model can be expressed as follows (Mehregan et al., 2013):

$$y_t = \begin{cases} \varphi_{0,1} + \varphi_{1,1} y_{t-1} + \varepsilon_t \text{if } s_t = 1 \\ \varphi_{0,2} + \varphi_{1,2} y_{t-1} + \varepsilon_t \text{if } s_t = 2 \end{cases}$$

Or, more concisely:

$$y_t = \varphi_{0,s_t} + \varphi_{1,s_t} y_{t-1} + \varepsilon_t$$

To complete the model, the properties of the process $s_t$ must be specified. In the Markov-switching framework, $s_t$ is assumed to follow a first-order Markov process, implying that $s_t$ depends solely on the regime in the previous period ($s_{t-1}$). The model is completed by defining the transition probabilities from one state to another as follows:

$$P(s_t = 1 \mid s_{t-1} = 1) = p_{11}$$
$$P(s_t = 2 \mid s_{t-1} = 1) = p_{12}$$
$$P(s_t = 1 \mid s_{t-1} = 2) = p_{21}$$
$$P(s_t = 2 \mid s_{t-1} = 2) = p_{22}$$

In the above expressions, $p_{ij}$ represents the probability of the Markov chain transitioning from state *i* at time *t* − *1* to state *j* at time *t*. These probabilities are always non-negative and satisfy the following conditions:

$$p_{11} + p_{12} = 1$$
$$p_{21} + p_{22} = 1$$

**Variable Definitions:**

Y: Financial cycles identified using the Hodrick–Prescott filter and positive and negative shocks generated within financial cycles. The ratio of liquidity to gross domestic product is used as the indicator of financial cycles.

GDP: Iran's gross domestic product at constant 2016 prices.

NFCLS: Shock to current government expenditures.

CFNAI: Shock to capital (developmental) government expenditures.

PPIC: Inflation rate.

Interest: Rate of return on participation-based credit facilities.

INV: Islamic participation bonds.

Consumption: Aggregate consumption.

In this study, positive and negative shocks to current and capital government expenditures are derived using the EGARCH model proposed by Nelson (1991). The time span of the research covers the period from 1987 to 2024 in Iran.

## Findings and Results

First, the results obtained from estimating the EGARCH model are used to compute shocks to current and capital government expenditures. In line with the existing literature on volatility models, to calculate shocks to current and capital expenditures, the ratio of government debt to gross domestic product must first be modeled using ARMA models, and the relevant lags for current and capital expenditures must be identified. For this purpose, the Box–Jenkins methodology is employed. The results of modeling current and capital expenditures are presented in the following tables.

**Table 1. Estimation Results of the Current Expenditure Model**

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| AR(1) | 1.014930 | 0.011936 | 85.03441 | 0.0000 |
| MA(1) | -0.275550 | 0.037883 | -7.273665 | 0.0000 |

**Table 2. Estimation Results of the Capital Expenditure Model**

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| AR(1) | 1.003883 | 0.000469 | 2142.666 | 0.0000 |
| MA(1) | -0.113128 | 0.031363 | -3.607082 | 0.0003 |

Based on the above models, current and capital expenditures are related to their own first lag through an AR(1) process and to the first lag of the error term through an MA(1) process.

To examine the presence of conditional heteroskedasticity in current and capital expenditures, the ARCH test is employed. The results of this test are reported in the following tables.

**Table 3. ARCH Test for the Current Expenditure Model**

| Test Statistic | Value | Probability |
|---|---|---|
| F-statistic | 2.214044 | Prob. F(1,34) = 0.0364 |
| Obs*R-squared | 2.218810 | Prob. Chi-Square(1) = 0.0363 |

**Table 4. ARCH Test for the Capital Expenditure Model**

| Test Statistic | Value | Probability |
|---|---|---|
| F-statistic | 3.198927 | Prob. F(1,34) = 0.0277 |
| Obs*R-squared | 3.210755 | Prob. Chi-Square(1) = 0.0271 |

Given the obtained probabilities, the null hypothesis of no conditional heteroskedasticity in current and capital expenditures is rejected. Therefore, both current and capital government expenditures exhibit conditional heteroskedasticity.

Finally, to obtain shocks to current and capital government expenditures, the EGARCH model proposed by Nelson (1991) is employed. One of the major limitations of the ARCH and GARCH approaches lies in their symmetry assumption; that is, positive and negative shocks of equal magnitude are assumed to have identical effects on volatility. However, fluctuations in current and capital government expenditures do not respond symmetrically to the nature of news (positive versus negative shocks). Accordingly, to address this limitation and to analyze the behavior of volatility, it is necessary to use an asymmetric model.

$$\ln\,\sigma_t^2=\alpha_0+\alpha_1\frac{|\,u_{t-1}\,|}{\sqrt{\sigma_{t-1}^2}}+\beta\ln\,\sigma_{t-1}^2+\gamma\frac{u_{t-1}}{\sqrt{\sigma_{t-1}^2}},\alpha_0=\omega-\alpha\sqrt{\frac{2}{\pi}},\alpha_1=\alpha$$

This model has several advantages. First, the dependent variable $\sigma_t^2$is expressed in logarithmic form; therefore, the coefficients of the explanatory variables may be either positive or negative, while $\sigma_t^2$remains strictly positive. Consequently, there is no need to impose non-negativity constraints on the coefficients. Second, this model explicitly accounts for asymmetric shock effects. The parameter $\gamma$is the coefficient on $u_{t-1}$, which may be positive or negative. Thus, $\gamma$captures the asymmetric effects of positive and negative shocks, whereas $\alpha$is associated only with the absolute value $|\,u_{t-1}\,|$. If $\gamma=0$, the model is symmetric; otherwise, it is asymmetric. The effect of positive shocks equals $\alpha + \gamma$, while the effect of negative shocks equals $\alpha - \gamma$. If $\gamma$ is negative, it indicates that negative shocks have a stronger effect than positive shocks, and vice versa.

**Table 5. EGARCH Model for Current Government Expenditures**

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| AR(1) | 1.014930 | 0.011936 | 85.03441 | 0.0000 |
| MA(1) | -0.275550 | 0.037883 | -7.273665 | 0.0000 |
| C(3) | -1.460988 | 0.302531 | -4.829211 | 0.0000 |
| C(4) | -0.557479 | 0.164041 | -3.398404 | 0.0007 |
| C(5) | 0.118678 | 0.138461 | 0.857123 | 0.3914 |
| C(6) | -0.842610 | 0.171831 | -4.903725 | 0.0000 |

Dependent Variable: GOVEXP

Method: ML ARCH – Normal distribution (BFGS / Marquardt steps

LOG(GARCH) = C(3) + C(4)*ABS(RESID(-1)/@SQRT(GARCH(-1))) + C(5)*RESID(-1)/@SQRT(GARCH(-1)) + C(6)*LOG(GARCH(-1))

**Table 6. EGARCH Model for Capital Government Expenditures**

| Variable | Coefficient | Std. Error | z-Statistic | Prob. |
|---|---|---|---|---|
| AR(1) | 1.003883 | 0.000469 | 2142.666 | 0.0000 |
| MA(1) | -0.113128 | 0.031363 | -3.607082 | 0.0003 |
| C(3) | 0.347601 | 4.2E-104 | 8.4E+102 | 0.0000 |
| C(4) | -0.863386 | 0.048117 | -17.94366 | 0.0000 |
| C(5) | -0.452962 | 0.010840 | -41.78519 | 0.0000 |
| C(6) | 0.740037 | 0.040186 | 18.41519 | 0.0000 |

Dependent Variable: GOVOMR

Method: ML ARCH – Normal distribution (BFGS / Marquardt steps)

LOG(GARCH) = C(3) + C(4)*ABS(RESID(-1)/@SQRT(GARCH(-1))) + C(5)*RESID(-1)/@SQRT(GARCH(-1)) + C(6)*LOG(GARCH(-1))

At this stage, the magnitude of shocks arising from current and capital government expenditures is calculated. For use in the main model, these magnitudes are transformed into variables. To this end, the *Make variance GARCH* command is applied to convert the estimated shocks into independent variables.

In this study, financial cycles are computed using the ratio of liquidity to gross domestic product, and the Hodrick–Prescott filter is applied for the period 1987–2024. The results of this procedure are illustrated in the figure below, and the resulting series is entered into the main model as the variable **Y**.

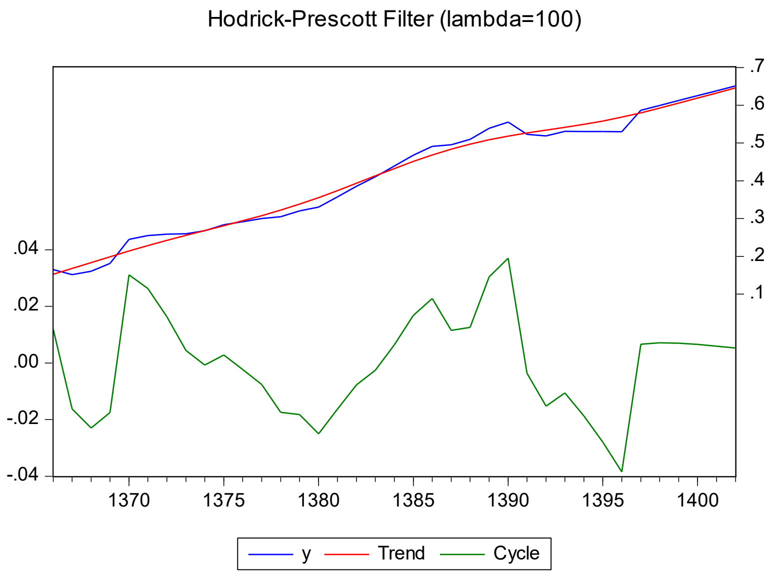


**Figure 1. Graphical representation of the gap in the liquidity-to-GDP ratio to illustrate the financial cycle**

According to Figure (1), the cyclical component obtained is used to identify recessionary and expansionary periods. To determine these phases based on the cyclical component, it is necessary to identify peaks and troughs. In this study, the approaches proposed by Hamburg and Verastendig (2008) and Chen, Guo, and Miller (2000) are employed to determine turning points. The results indicate that the Iranian economy experienced a total of six cycles over the period 1987–2024.

Specifically, the period from 1987 to 1990 corresponds to a recession, 1991 to 1995 to an expansion, 1996 to 2004 to a recession, 2004 to 2012 to an expansion, 2012 to 2018 to a recession, and from 2018 to 2024 to an expansion in Iran's financial cycle. The recent expansion in the financial cycle can be attributed to the excessive growth of liquidity volume in the country.

This article employs the conventional Phillips–Perron (PP) unit root test.

**Table 7. Results of the PP Test at Levels of the Model Variables**

| Variable | PP Statistic | Probability | Status |
|---|---|---|---|
| CFNAI | -0.266390 | 0.5831 | Non-stationary |
| D(CFNAI) | -14.02772 | 0.0000 | I(1) |
| CONSUMPTION | -1.343491 | 0.1627 | Non-stationary |
| D(CONSUMPTION) | -10.84401 | 0.0000 | I(1) |
| GDP | -2.991803 | 0.1483 | Non-stationary |
| D(GDP) | -8.391937 | 0.0000 | I(1) |
| INTEREST | -1.568711 | 0.1084 | Non-stationary |
| D(INTEREST) | -7.181535 | 0.0000 | I(1) |
| INV | -0.690786 | 0.4106 | Non-stationary |
| D(INV) | -8.286790 | 0.0000 | I(1) |
| NFCLS | -1.024842 | 0.2691 | Non-stationary |
| D(NFCLS) | -8.206334 | 0.0000 | I(1) |
| PPIC | -0.412226 | 0.5278 | Non-stationary |
| D(PPIC) | -6.158476 | 0.0000 | I(1) |
| Y | -0.305372 | 0.5686 | Non-stationary |
| D(Y) | -10.86299 | 0.0000 | I(1) |

Based on the theoretical foundations of stationarity tests, the null hypothesis $H_0$ in these tests is the presence of a unit root (non-stationarity). Considering the test results, it can be concluded that all variables in the model are stationary at first differences, that is, they are integrated of order one, I(1).

Since the variables in the model have the same order of integration, I(1), the Johansen–Juselius cointegration test is employed to determine the existence of a long-run equilibrium relationship among the variables. To conduct this test, the number of cointegrating vectors must be specified. For this purpose, an appropriate specification regarding the presence or absence of an intercept and a deterministic trend in the cointegrating vector must be selected. In this context, five specifications are considered: (1) no intercept and no trend; (2) restricted intercept and no trend; (3) unrestricted intercept and no trend; (4) unrestricted intercept and restricted trend; and (5) unrestricted intercept and unrestricted trend. These specifications range from the most restrictive (Model 1) to the least restrictive (Model 5).

Subsequently, the null hypothesis of no cointegrating vector is tested against the alternative of at least one cointegrating vector, followed by tests for the existence of additional vectors up to $n-1$, where $n$ denotes the number of variables. A summary of the trace statistic ($\lambda_{\text{Trace}}$) and the maximum eigenvalue statistic ($\lambda_{\text{Max}}$) for the five specifications is reported in Table 8. As shown, the null hypothesis of no cointegration is rejected in favor of at least one cointegrating vector among the variables.

**Table 8. Summary of the Number of Cointegrating Vectors**

| Model | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| Trace Test | 4 | 6 | 3 | 3 | 4 |
| Max-Eigenvalue Test | 2 | 3 | 3 | 5 | 2 |

The estimation results and the corresponding cointegration tests for the selected specification are reported in Table 9. According to the trace test, the existence of three cointegrating vectors is confirmed, and the maximum eigenvalue test likewise confirms the presence of three cointegrating vectors at the 5% significance level.

**Table 9. Johansen Cointegration Test Results**

| Max-Eigen Statistic | 95% Critical Value | Probability | Trace Statistic | 95% Critical Value | Probability | $H_1$ | $H_0$ |
|---|---|---|---|---|---|---|---|
| 67.57026 | 52.36261 | 0.0007 | 228.9846 | 159.5297 | 0.0000 | r = 1 | r = 0 |
| 61.49758 | 46.23142 | 0.0006 | 161.4143 | 125.6154 | 0.0001 | r = 2 | r ≤ 1 |
| 46.14144 | 40.07757 | 0.0092 | 99.91674 | 95.75366 | 0.0250 | r = 3 | r ≤ 2 |
| 24.76886 | 33.87687 | 0.4008 | 53.77530 | 69.81889 | 0.4713 | r = 4 | r ≤ 3 |
| 15.00804 | 27.58434 | 0.7476 | 29.00644 | 47.85613 | 0.7680 | r = 5 | r ≤ 4 |
| 10.75474 | 21.13162 | 0.6717 | 13.99840 | 29.79707 | 0.8407 | r = 6 | r ≤ 5 |
| 2.469964 | 14.26460 | 0.9756 | 3.243663 | 15.49471 | 0.9547 | r = 7 | r ≤ 6 |
| 0.773699 | 3.841466 | 0.3791 | 0.773699 | 3.841466 | 0.3791 | r = 8 | r ≤ 7 |

The Markov-switching model is an appropriate estimation framework only if the underlying data-generating process is nonlinear. To verify nonlinearity in the data pattern, the LR (likelihood ratio) test is applied. The test statistic is computed from the maximum likelihood values of two competing models: a single-regime model (linear) and a two-regime model (nonlinear). The LR statistic follows a chi-square distribution. If the computed statistic exceeds the critical value at the desired confidence level, one may conclude that the linear model is not adequate at that confidence level and that the nonlinear model should be used instead.

**Table 10. LR Test Results**

| Probability | Degrees of Freedom | Test Statistic |
|---|---|---|
| 0.0000 | 7 | 45.326 |

As shown in the table above, the LR test statistic is greater than its critical value at the 5% significance level. Therefore, it can be concluded that, rather than linear models, the nonlinear Markov-switching approach is more appropriate for estimating the model. The following table presents the estimation results of the Markov-switching model for the above equation.

**Table 11. Parameter Estimates of the Markov Switching MS(2)-AR(1) Model during Recession and Expansion**

| Variable | Coefficient | Std. Error | t-Statistic | Probability |
|---|---|---|---|---|
| c1 | 0.329059 | 0.136048 | 2.418698 | 0.0234 |
| c2 | -0.582868 | 0.239013 | -2.438644 | 0.0149 |
| σ1 | 0.096233 | 0.026454 | 3.637702 | 0.0003 |
| σ2 | 0.433451 | 0.159418 | 2.718959 | 0.0218 |
| ΔY(-1) | 0.393667 | 0.142781 | 2.757139 | 0.0201 |
| CFNAI (1) | -0.110132 | 0.055542 | -1.982849 | 0.0491 |
| CFNAI (2) | 0.016839 | 0.006021 | 2.796881 | 0.0049 |
| NFCLS (1) | -0.016559 | 0.007091 | -2.335214 | 0.0355 |
| NFCLS (2) | 0.089278 | 0.027937 | 3.195642 | 0.0015 |
| PPIC (1) | 0.197885 | 0.085146 | 2.324058 | 0.0201 |
| PPIC (2) | 0.022717 | 0.010298 | 2.205968 | 0.0274 |
| Interest (1) | -0.012901 | 0.004822 | -2.675502 | 0.0077 |
| Interest (2) | 0.235516 | 0.085117 | 2.762255 | 0.0053 |

| | | | | |
|---|---|---|---|---|
| GDP (1) | -0.014838 | 0.006607 | -2.245690 | 0.0378 |
| GDP (2) | -0.120704 | 0.060604 | -1.991677 | 0.0487 |
| INV (1) | -0.521556 | 0.300193 | -1.737401 | 0.0823 |
| INV (2) | -0.047188 | 0.007428 | -6.352493 | 0.0000 |
| Consumption (1) | -0.035862 | 0.014372 | -2.495362 | 0.0129 |
| Consumption (2) | 0.289564 | 0.208206 | 1.390754 | 0.1664 |

Based on the Markov model estimation results, most coefficients are statistically significant at the 95% confidence level, and their signs are consistent with theoretical foundations. The intercept equals 0.32 in Regime 1 and −0.58 in Regime 2. According to Hamilton (1988), a regime with a negative intercept represents a recessionary regime, whereas a regime with a positive intercept represents an expansionary regime. Accordingly, in this study, Regime 1 represents the expansion period and Regime 2 represents the recession period. The variance of the disturbance terms is 0.09 in Regime 1 and 0.43 in Regime 2. These values indicate that Regime 1 (expansion) exhibits lower volatility than Regime 2 (recession) in the present study. The Markov estimation results in the table above indicate that negative shocks to current and capital government expenditures, consumption, and inflation have negative effects on the financial cycle in both expansionary and recessionary periods. Shocks to current and capital government expenditures reduce the liquidity-to-GDP ratio in Regime 1 (expansion), whereas in Regime 2 (recession) this effect is positive. The output variable also exerts a negative effect on the liquidity-to-GDP ratio in both regimes (recession and expansion).

In Iran, because financial markets are typically limited and imperfect, investment depends on the feasibility of mobilizing financing from domestic savings. As a result, an increase in bond yields can expand the capacity to finance projects. Therefore, an increase in bond yields, on the one hand, may reduce investment by raising the cost of investment and consumption; on the other hand, by facilitating greater access to project financing, it may increase investment. Overall, this renders the net effect of bond yields ambiguous. Moreover, under conditions of financial repression (a controlled banking system), due to negative real returns on bank-based bond yields (interest) arising from inflation and controlled credit conditions, investment and output are considered the primary casualties of financial repression.

Since 2006, due to directive policies of the central bank, the use of participation-based contracts such as civil partnership, legal partnership, and *mudarabah* has increased. However, the key issue concerns the correct implementation of participation-based contracts in a manner that safeguards banks' interests. Addressing this issue requires rigorous analysis and proper selection of investment projects, continuous monitoring and management of project implementation, and, ultimately, the calculation and determination of the actual realized profit of the partnership. These requirements necessitate expertise, technical knowledge, and close operational engagement with participation-based projects; nevertheless, banks generally lack specialized human resources as well as sufficient knowledge and skills in these areas.

On the other hand, an examination of the time-series patterns of the bond yield rate, output, and the participation profit rate corroborates the model estimation results: despite the increased use of participation-based contracts and investment in this domain, the bond yield rate has risen, yet this increase has not translated into the real sector of the Iranian economy. As the Markov model estimation results for recession and expansion also indicate, in Regime 1 and with increases in current and capital government expenditures, the financial cycle rises during the expansion period. However, a critical issue in Iran is the persistence (duration) of recessionary and expansionary phases. By referring to Table 5, it can also be concluded that, over the period examined in this study, the economy experienced

22 expansionary periods compared to 15 recessionary periods, which, under existing conditions, reduces the risk associated with adopting participation-based contracts. Since expectations play a crucial role in investment decisions, any political and economic instability can negatively affect investment and reduce long-term investment risk.

The following figure shows the probability that each year in the sample belongs to each of the two regimes. The dashed lines in the two lower graphs represent these probabilities. As the figure indicates, the probabilities of Regime 1 and Regime 2 sum to one in each year. The shaded areas in the graphs indicate the classification of years across the two regimes.

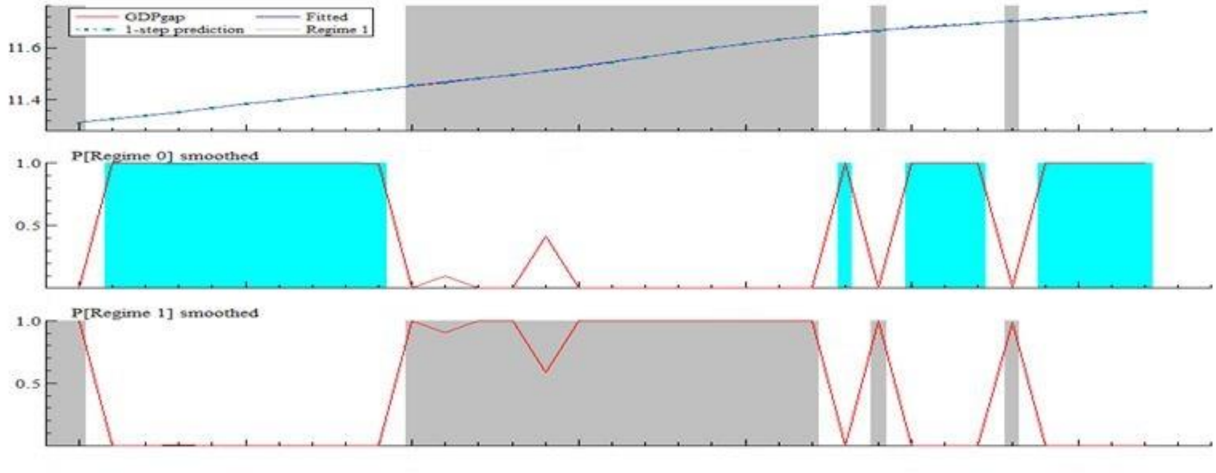


**Figure 2. Probability of Each Year Falling into the Two Extracted Regimes (Financial Cycle of Business Cycles Model)**

The following table reports the years classified under each regime; in other words, it reflects the performance of Iran's credit facilities in using participation-based contracts during recessionary and expansionary periods. As the table indicates, for the financial-cycle business-cycle model, the occurrence of oil and monetary shocks in the years 1987–1995, 2009, 2011–2013, and 2015–2024 increased the financial cycle gap, whereas in 1996–2008, 2010, and 2014 the output gap decreased.

**Table 12. Years Classified under Each Regime for the Model**

| Regime | Iran |
|---|---|
| Regime 1 | 9 years: 1987–1995 |
| | 1 year: 2009 |
| | 3 years: 2011–2013 |
| | 9 years: 2015–2024 |
| Regime 2 | 13 years: 1996–2008 |
| | 1 year: 2010 |
| | 1 year: 2014 |

To examine the effects of government expenditure shocks and Islamic participation bonds on the financial cycle, it is important to consider the impacts of inflation, consumption, and gross domestic product as well. From 1990 to 1997, global oil consumption increased by 2.6 million barrels per day. The decline in Russia's oil production also played a significant role in the price increases during this period; specifically, between 1990 and 1996 Russia's oil production fell from about 11 million barrels per day to less than 6 million barrels per day. However, the price increases ended rapidly because OPEC did not adequately take into account the effects of the Asian economic crisis in 1997 and 1998. Oil prices in August 1990 rose by another 13% due to the war between Iraq and U.S.-allied forces. After the war ended, the United Nations imposed sanctions on Iraq's oil exports. This action caused oil prices to reach a new peak again within nine months.

The Asian economic crisis of 1997 and 1998 occurred following the Thai government's decision to float its currency—the Thai baht—on July 2, 1997. As a result of this policy, which was suddenly implemented without the necessary preconditions, the value of the baht depreciated. This abrupt depreciation severely damaged the country's economy and caused Thailand's export goods to lose their competitiveness in global markets. In the early 2000s, the United States faced a severe financial credit crisis, which was effectively initiated by banks. American borrowers encountered repayment formulas that appeared unusual to them and changed over time. In the initial years, loan interest rates were low and mortgage borrowers were able to service their debts. As loan interest rates increased, many U.S. borrowers were no longer able to repay their debts, and a number of major U.S. financial institutions—including Goldman Sachs, JPMorgan, Merrill Lynch, and Lehman Brothers—faced a major economic crisis. Ultimately, on September 15, 2008, Lehman Brothers' bankruptcy inflicted a substantial shock on the U.S. economy.

The upward trend in oil prices from 2002 coincided with problems in Venezuela due to a sharp decline in that country's production. In March 2003, after Venezuela's output had partially recovered, U.S. military operations in Iraq began. In October 2004, oil prices surpassed the threshold of USD 50 per barrel, and in June of that year prices surged and eventually crossed the psychological barrier of USD 60 per barrel. At the end of 2005 and the beginning of 2006, owing to unrest in Nigeria and intensifying disputes between Iran and Western countries over Iran's nuclear activities, oil prices continued their upward trajectory; by the end of January 2006, oil prices rose to more than USD 68 per barrel. In May 2007, each barrel of oil traded at USD 73.93, an increase attributed to unrest in Nigeria. Although oil markets experienced substantial volatility in 2007 and even recorded a peak of USD 99.29, the realization of "USD 100 oil" ultimately occurred in the first days of 2008. The rise in oil prices continued throughout 2008 and reached USD 148 per barrel in July 2008.

A relevant shock also occurred in 2014–2015, when a decline in oil demand and the resistance of some OPEC members to reduce production pushed oil prices to their lowest level in recent years. During 1992–1996, as well as in 2000, 2002, and 2005, major monetary shocks exhibited the highest probability of regime occurrence in Iran's economy. In 1992–1996, two factors contributed to an expansion in money supply and liquidity and the emergence of major monetary shocks in Iran. The first factor was the establishment of the foreign exchange obligations reserve account following the move toward a floating exchange rate during those years, which increased the government's liabilities to the central bank and thereby facilitated monetary expansion. The second factor was the rise in state-owned enterprises' debt to the banking system, which caused the change in the outstanding debt balance of the public sector to the banking system to increase from 2,254 billion rials in 1992 to 10,848 billion rials in 2006. In 2002, the change in the outstanding debt balance of the public sector to the banking system again grew by 332% compared to the previous year. On the other hand, the debtor balance of the foreign exchange obligations reserve account, which had declined during 2000–2002 due to partial repayment of the government's debts to the central bank, increased again in 2004–2005. Therefore, examining the origins of major monetary shocks in Iran during the study period indicates that the government's fiscal policy and the substantial debt of the public sector to the banking system were key sources of major monetary shocks.

Moreover, growth in oil revenues is one of the factors driving output growth and credit expansion in Iran, and consequently, it contributes to higher returns on participation-based contracts. For example, between 1999 and 2007, as oil prices rose, Iran's economy experienced strong growth; conversely, one of the reasons for the slowdown in economic growth in recent years has been the lack of robust growth in oil prices.

Table 6 reports the transition probabilities from one regime to another and the regime duration. As can be observed, based on the transition probability functions for Iran's estimated model in the table below, if Iran's economy is in recession at time *t*, then—given current and capital expenditure shocks and inflation—there is a 0.25 probability that it will remain in the same state, and a 0.75 probability that under other factors it will shift to the recessionary state. If the economy is in expansion at time *t*, then—given current and capital expenditure shocks and inflation—there is a 0.73 probability that it will remain in the same state at time *t + 1*, and a 0.27 probability that under other factors it will منتقل (shift) to a recessionary state. In addition, the economy's exposure to expansion in the present study is 22 periods compared to 15 recessionary periods. Thus, it is evident that persistence in expansion, as well as a return from recession to expansion, is associated with a relatively high probability in Iran's economy, and these factors reduce the risk of using participation-based contracts and their returns for investment in participation-based projects.

**Table 13. Transition Probabilities from One Regime to Another**

| | Regime 1 | Regime 2 |
|---|---|---|
| Regime 1 | 0.73 | 0.27 |
| Regime 2 | 0.25 | 0.75 |

As noted in the model introduction, the disturbance terms in the Markov-switching model should be normally distributed and free from autocorrelation and heteroskedasticity. The results of the tests corresponding to these properties are presented below.

**Table 14. Results of Diagnostic Tests**

| Test Type | Test Statistic | Test Value | Probability |
|---|---|---|---|
| No autocorrelation test | $\chi^2(3)$ | 6.236 | 0.29 |
| Normality test | $\chi^2(2)$ | 5.329 | 0.18 |
| Homoskedasticity (equal variance) test | F(1,1) | 0.415 | 0.91 |

The no-autocorrelation test results indicate that, at the 5% significance level, the null hypothesis of no autocorrelation cannot be rejected; therefore, it can be inferred that the disturbance terms are free from autocorrelation. The normality test also indicates that the null hypothesis of normally distributed residuals for the estimated model is not rejected. In addition, the homoskedasticity test results show that the null hypothesis of equal variance of the disturbance terms is not rejected.

## Discussion and Conclusion

The findings of this study provide robust evidence that government expenditure shocks and Islamic participation bonds exert asymmetric and regime-dependent effects on the financial cycle in Iran. Consistent with the nonlinear nature of macro-financial dynamics emphasized in recent literature, the results confirm that the impacts of fiscal shocks differ markedly between expansionary and recessionary regimes, thereby validating the use of a Markov regime-switching framework. In line with the estimated parameters, government current and capital expenditure shocks reduce financial-cycle volatility during expansionary periods while exerting expansionary or stabilizing effects during recessions, highlighting the state-contingent effectiveness of fiscal policy (2, 3). These findings resonate with the argument that fiscal multipliers and financial responses are inherently asymmetric and depend on prevailing macroeconomic conditions (1, 4).

A central result of the study is the differential role of Islamic participation bonds across regimes. During expansionary phases, increases in participation-based financing are associated with moderation in the liquidity-to-GDP ratio, suggesting that risk-sharing instruments may curb excessive financial exuberance by tying financing more closely to real economic activity. This outcome aligns with theoretical and empirical arguments that Islamic finance can dampen speculative dynamics and reduce leverage-driven volatility (8, 9). Conversely, in recessionary regimes, participation bonds appear to support the financial cycle by facilitating access to non-interest-based financing when conventional credit channels are constrained. This countercyclical role is consistent with evidence from other Islamic and emerging economies, where participation-based instruments have been shown to mitigate downturns by sustaining investment and consumption (12, 13).

The observed asymmetry in the response of the financial cycle to government expenditure shocks also reflects the interaction between fiscal policy and uncertainty. During recessions, heightened uncertainty amplifies the effects of expansionary fiscal actions on financial conditions, whereas in expansions, similar shocks may be partially neutralized by tighter financial constraints or inflationary pressures. This interpretation is consistent with studies emphasizing the role of uncertainty and news shocks in shaping macro-financial responses (5, 15). The stronger sensitivity of the financial cycle to negative shocks, compared to positive ones, further corroborates the asymmetric transmission mechanisms highlighted in recent macroeconomic research (7, 16).

Inflation emerges as a critical mediating variable in the fiscal–financial nexus. The results indicate that inflationary pressures associated with government spending shocks tend to dampen the financial cycle during expansions, likely by eroding real returns and tightening effective financing conditions. During recessions, however, moderate inflation may ease real debt burdens and support liquidity, thereby reinforcing the stabilizing role of fiscal policy. These findings are in line with empirical evidence suggesting that inflation can both constrain and facilitate financial activity depending on the macroeconomic regime (3, 4). They also underscore the importance of coordinated fiscal and monetary policy, particularly in economies facing recurrent supply-side and external shocks.

Consumption and output responses further illustrate the asymmetric dynamics uncovered in this study. The negative impact of consumption shocks on the financial cycle across both regimes suggests that household behavior, shaped by income uncertainty and credit constraints, plays a key role in transmitting fiscal shocks to financial variables. This outcome is consistent with the view that consumption responses are more pronounced during downturns, when precautionary savings motives dominate (5). Similarly, the negative association between output and the liquidity-to-GDP ratio reflects structural characteristics of the Iranian economy, where increases in real activity may not be matched by proportional improvements in financial intermediation, particularly under conditions of sanctions and financial repression (18).

The results also highlight the relevance of oil price dynamics and external shocks in shaping fiscal–financial interactions. Periods characterized by oil revenue booms coincide with stronger expansionary regimes and greater use of participation-based instruments, while oil price busts amplify recessionary dynamics and increase reliance on government spending to stabilize the economy. This pattern is consistent with global evidence on the trade and financial effects of oil price shocks (11, 14). In Iran’s case, oil-driven fiscal expansions appear to have supported financial cycles primarily through liquidity effects rather than sustained real-sector investment, echoing concerns raised in the literature about the limited trickle-down effects of resource-based fiscal spending (17).

From a methodological perspective, the study's findings reinforce the importance of nonlinear modeling approaches in fiscal policy analysis. Linear models would have obscured the regime-dependent effects identified here, potentially leading to misleading policy conclusions. The significant likelihood ratio test results and the distinct parameter estimates across regimes support the growing consensus that macro-financial relationships are best captured through nonlinear and state-dependent frameworks (6, 19). This is particularly relevant for economies experiencing frequent structural breaks, policy shifts, and external shocks.

Comparing the present findings with international evidence reveals both commonalities and context-specific features. Similar to results reported for other Islamic and emerging economies, fiscal shocks in Iran exhibit stronger effects during recessions than expansions (12, 13). However, the role of Islamic participation bonds appears more nuanced, reflecting institutional constraints, limited project evaluation capacity, and the coexistence of conventional and Islamic banking practices. While global studies highlight the resilience of sukuk markets to certain shocks, they also emphasize increasing integration with global financial cycles (7, 10), a trend that may heighten vulnerability to external disturbances.

Overall, the results suggest that government expenditure shocks and Islamic participation bonds jointly shape Iran's financial cycle in a complex, asymmetric manner. Fiscal expansions can stabilize financial conditions during downturns but risk fueling imbalances during booms if not accompanied by effective financial governance. Participation-based instruments offer potential stabilizing benefits, yet their effectiveness depends on proper implementation, project selection, and institutional capacity. These conclusions align with broader debates on the role of fiscal policy and alternative finance in promoting macro-financial stability in uncertain and volatile environments (3, 16).

Despite its contributions, this study is subject to several limitations. First, the analysis relies on aggregate macroeconomic and financial indicators, which may mask sectoral heterogeneity and micro-level transmission mechanisms. Second, data constraints limit the ability to distinguish between different types of Islamic participation instruments and their specific contractual features. Third, the regime-switching framework, while capturing nonlinearities, assumes discrete regimes and may not fully reflect gradual transitions or overlapping states. Finally, country-specific factors such as sanctions, institutional reforms, and informal financial practices may introduce unobserved influences that are not explicitly modeled.

Future research could extend this analysis by incorporating sectoral or firm-level data to better understand the micro-foundations of fiscal–financial interactions. Comparative studies across multiple Islamic and non-Islamic economies would also help to isolate the role of institutional and financial structures in shaping asymmetric responses. Moreover, integrating time-varying parameter models or smooth transition frameworks could capture more nuanced regime dynamics. Finally, future work could explicitly model the interaction between fiscal policy, monetary policy, and macroprudential tools to provide a more comprehensive assessment of policy coordination.

From a practical perspective, policymakers should account for regime-dependent effects when designing fiscal interventions, recognizing that expansionary spending may yield different financial outcomes depending on the state of the economy. Strengthening the institutional framework governing participation-based financing, including project evaluation, monitoring, and transparency, can enhance the stabilizing role of Islamic instruments. Finally, improving coordination between fiscal authorities, financial regulators, and the banking sector can help ensure that government spending supports sustainable financial cycles rather than amplifying volatility.

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## Authors' Contributions

All authors equally contributed to this study.

## Declaration of Interest

The authors of this article declared no conflict of interest.

## Ethical Considerations

All ethical principles were adheried in conducting and writing this article.

## Transparency of Data

In accordance with the principles of transparency and open research, we declare that all data and materials used in this study are available upon request.

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