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Examining the Effect of Exchange Rate Shocks on National Business Cycles under Linear and Nonlinear Regimes Using the Smooth Transition Regression Model

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ABSTRACT

The exchange rate is regarded as one of the most important macroeconomic variables in an economy. Likewise, the existence of business cycles is a plausible phenomenon in the economies of all countries and is influenced by other macroeconomic variables. In this article, the effect of exchange rate shocks on the country's business cycles under different regimes is examined using the threshold LSTAR model. The present study is applied in terms of purpose and descriptive–analytical in nature. The statistical population of the study consists of data related to the Iranian economy over the period 1978 to 2022, selected on an annual basis, and the Smooth Transition Regression (LSTAR) model estimated using EViews software is employed for model estimation. First, the stationarity of the research variables was examined, and subsequently, long-run relationships among the variables were analyzed using cointegration tests. In the LSTAR regression model, in the first section, the real exchange rate with a lag of three periods was selected as the threshold variable. The results of the analysis showed that exchange rate shocks have different effects on the country's business cycles across different regimes.

Keywords: Exchange rate shock, monetary policy, business cycles, threshold LSTAR model

Introduction

The exchange rate occupies a central position in macroeconomic analysis, particularly in open economies where external shocks are rapidly transmitted into domestic markets and macroeconomic aggregates. Fluctuations in the exchange rate influence production decisions, consumption patterns, investment incentives, and expectations, thereby shaping the phases of economic expansion and contraction over time. In developing and emerging economies, where structural rigidities, commodity dependence, and financial vulnerabilities are more pronounced, exchange rate movements often play a decisive role in amplifying business cycle volatility (1, 2). Consequently, understanding the mechanisms through which exchange rate dynamics interact with monetary policy and real economic activity has become a core concern of modern macroeconomic research.

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Business cycles, defined as recurrent but irregular fluctuations in economic activity around a long-term growth trend, are influenced by a complex set of internal and external factors. Classical and Keynesian traditions emphasize aggregate demand shocks, monetary conditions, and fiscal interventions as primary drivers of cyclical movements, while modern open-economy macroeconomics highlights the importance of external shocks, capital flows, and exchange rate regimes (3, 4). In this framework, exchange rate shocks are not merely nominal disturbances; they interact with price rigidities, expectations, and policy responses to generate persistent real effects. These interactions are particularly salient in economies with high import dependence, oil revenues, and exposure to global financial conditions.

Monetary policy constitutes another critical channel through which macroeconomic stability is pursued, especially in the presence of exchange rate volatility. Central banks attempt to mitigate excessive fluctuations in inflation, output, and financial markets through interest rate adjustments, liquidity management, and, in some cases, direct intervention in foreign exchange markets (5, 6). However, the effectiveness of monetary policy is often contingent on the state of the economy and the prevailing exchange rate regime. During periods of high volatility or structural imbalance, conventional policy tools may exhibit asymmetric or nonlinear effects, leading to different outcomes across expansionary and recessionary phases (7).

Theoretical contributions in international macroeconomics suggest that exchange rate movements can have both stabilizing and destabilizing effects on business cycles. On one hand, exchange rate flexibility may act as a shock absorber by facilitating external adjustment and reallocating resources in response to terms-of-trade shocks. On the other hand, sharp depreciations or appreciations can exacerbate inflationary pressures, disrupt balance sheets, and weaken investment, thereby intensifying cyclical downturns (2, 4). The net effect depends on structural characteristics such as financial depth, trade openness, and the credibility of monetary policy.

Empirical research has increasingly documented that the relationship between exchange rates, monetary policy, and business cycles is nonlinear and regime-dependent. Linear models, while analytically convenient, often fail to capture threshold effects, asymmetries, and structural breaks that characterize real-world economies. As a result, nonlinear econometric approaches—such as smooth transition regression and regime-switching models—have gained prominence in analyzing macroeconomic dynamics (7, 8). These models allow coefficients to vary across regimes, providing a more nuanced representation of how economic relationships change under different conditions.

Within this strand of literature, the smooth transition regression framework has been particularly useful in modeling gradual shifts between regimes driven by observable transition variables, such as exchange rates or inflation indicators. Unlike abrupt regime-switching models, smooth transition models capture continuous adjustments and threshold behavior, which are more consistent with the gradual nature of policy transmission and expectation formation (9, 10). This methodological advantage is especially relevant for economies experiencing frequent but uneven shocks, where policy responses and market adjustments unfold over time rather than instantaneously.

The Iranian economy provides a compelling context for examining these issues due to its structural characteristics and exposure to multiple sources of volatility. Heavy reliance on oil revenues, recurrent sanctions, managed exchange rate arrangements, and episodes of high inflation have collectively contributed to pronounced business cycle fluctuations. Exchange rate movements in Iran are often driven by external shocks, policy interventions, and expectations, making their impact on output and employment both significant and complex (11, 12). Consequently, analyzing exchange rate shocks within a nonlinear framework can yield valuable insights into the cyclical behavior of the economy.

A growing body of empirical evidence highlights the importance of exchange rate volatility in shaping macroeconomic outcomes in Iran and comparable economies. Studies focusing on business cycles have shown that exchange rate fluctuations can exert asymmetric effects on output gaps, with stronger impacts during downturns than expansions (9, 10). Similarly, research on unemployment, income distribution, and inflation underscores that exchange rate shocks interact with monetary policy and structural constraints, producing heterogeneous effects across economic regimes (13-15).

Recent international studies further corroborate the nonlinear nature of exchange rate effects. Evidence from oil-exporting and developing economies suggests that the transmission of exchange rate shocks to trade balances, export revenues, and financial markets depends on the level of volatility, policy credibility, and interest rate conditions (16-18). These findings reinforce the argument that a single linear specification may obscure critical dynamics, particularly in environments characterized by frequent regime changes and external disturbances.

The interaction between exchange rates and monetary policy uncertainty has also attracted increasing attention. Uncertainty regarding policy direction can amplify the adverse effects of exchange rate volatility on financial stability and real activity, as agents postpone investment and consumption decisions in response to heightened risk (19, 20). In such contexts, nonlinear models offer a powerful tool for disentangling the joint effects of policy and market-driven shocks across different states of the economy.

Despite the expanding literature, several gaps remain. First, many studies focus on isolated outcomes such as inflation, unemployment, or trade balances, rather than providing an integrated analysis of business cycles as a comprehensive macroeconomic phenomenon. Second, empirical evidence on the regime-dependent effects of exchange rate shocks and monetary policy in Iran remains limited, particularly within frameworks that allow for smooth transitions between regimes. Third, the identification of appropriate threshold variables—such as real exchange rate changes—requires careful econometric treatment to ensure robust inference (9, 10).

Addressing these gaps is essential for both academic understanding and policy design. From a policy perspective, recognizing that exchange rate shocks and monetary interventions may have different effects depending on the economic regime can enhance the effectiveness of stabilization strategies. Central banks operating in volatile environments need to account for nonlinearities and thresholds when formulating policy responses, especially in economies exposed to external shocks and structural constraints (5, 6). From an academic standpoint, employing advanced nonlinear econometric techniques contributes to a more accurate representation of macroeconomic dynamics and enriches the empirical literature on business cycles.

Accordingly, this study situates itself at the intersection of exchange rate economics, monetary policy analysis, and nonlinear business cycle modeling. By drawing on theoretical insights from open-economy macroeconomics and recent empirical advances, it adopts a smooth transition regression framework to capture regime-dependent dynamics in the Iranian economy. In doing so, it builds upon and extends prior research that has emphasized the importance of threshold effects and asymmetric responses to macroeconomic shocks (10-12).

The novelty of the present study lies in its explicit focus on the joint effects of exchange rate shocks and monetary policy on business cycles under linear and nonlinear regimes, using a threshold-based LSTAR approach. By allowing the impact of key macroeconomic variables to vary smoothly across regimes defined by exchange rate dynamics, the study provides a richer and more policy-relevant understanding of cyclical fluctuations in an oil-dependent, sanction-affected economy.

The aim of this study is to examine the effects of exchange rate shocks and central bank monetary policy on Iran's business cycles under linear and nonlinear regimes using a threshold smooth transition regression (LSTAR) model.

Methods and Materials

From the perspective of purpose, this study is applied, and in terms of data and information collection, it is descriptive with a causal orientation. The research methodology is ex post facto; practically, there is no possibility or capacity to intervene in or manipulate the existing realities of the population under study, and the research is confined solely to examining the current conditions. This study seeks to examine the effects of exchange rate shocks and the Central Bank's monetary policy on the country's business cycles under linear and nonlinear regimes by employing the threshold LSTAR model. In other words, with respect to its objective, the study is applied, and in terms of data and information collection, it is descriptive–causal. The methodology is ex post facto. In this research, an attempt is made—through theoretical elaboration, model design, and the application of econometric methods—to analyze the effects of exchange rate shocks and the Central Bank's monetary policy on the country's business cycles under linear and nonlinear regimes using the threshold LSTAR model via the LSTAR autoregressive framework, such that the empirical effects of this relationship are subjected to analysis.

Following the studies of Sidney et al. (2019) and Yıldırım Karaman (2017), this study investigates the effects of exchange rate shocks and the Central Bank's monetary policy on the output gap over business cycles. The general form of the LSTR model, given that the dependent variable is the country's business cycles and the explanatory variables include exchange rate shocks, monetary policy, the price of Iranian heavy crude oil, the real interest rate, and others, is specified as follows:

$GDPgap_t = \pi Z_t + (\theta Z_t)F(\gamma \cdot C \cdot S_t) + U_t$

$cons_t = \pi Z_t + (\theta Z_t)F(\gamma \cdot C \cdot S_t) + U_t$

$GDPgap_t = \varphi'(INT_t, EX_t) + (\theta' INT_t, EX_t)G(INT_t, \gamma, c) + u_t \qquad (1)$

$\{t = 1, \dots, T\}$

where the transition function F is defined as:

$F(\gamma, s_t, c) = (1 + \exp\{-\gamma(s_t - c)\})^{-1}, \quad \gamma > 0$

The final estimated model is expressed as:

$GDPgap_{i,t} = \alpha_0 + a_1\, INTF_t + a_2\, POILSH_t + a_3\, RET_t + a_4\, M_t + a_5\, DMEXP_t + a_6\, R_t + U_t \qquad (2)$

GDPgap: Output gap, representing potential output minus actual output; the Hodrick–Prescott filter is used to compute the output gap.

Z_t: Vector of variables.

$POILSH_t$: Oil price.

RET_t: Exchange rate (real price).

M_t: Money supply (nominal).

$INTF_t$: Inflation rate.

$DMEXP_t$: Sanctions dummy variable for the years of sanctions on the country's oil exports (approved and implemented by the European Union and the United States Senate against Iran in 2012); therefore, it takes the value of one for 2012 onward and zero for other years.

It is assumed that the dependent variable GDPgap is solely a function of its own lagged values. Under this assumption, and considering a two-regime transition function, the following relationship is obtained:

$$GDPgap_t = (\theta_0 + \theta_1 \, GDPgap_{t-1} + \cdots + \theta_p \, GDPgap_{t-p}) + (\varphi_0 + \varphi_1 \, GDPgap_{t-1} + \cdots + \varphi_p \, GDPgap_{t-p}) \, G(EX_t, \gamma, c) + u_t \qquad (3)$$

$$G(EX_t, \gamma, c) = 1 / (1 + \exp\{-\gamma(EX_t - c)\}) \qquad (4)$$

The time span of the present study consists of annual data from 1978 to 2022, and all data were extracted from the Central Bank's website.

Moreover, using the Wimark (1995) model and the study of Rui et al. (2016), the degree of Central Bank intervention is calculated as follows:

$$I_t = (\eta \Delta r_t) / EMP_t = (\eta \Delta r_t) / (\Delta e_t + \eta \Delta r_t) \qquad (5)$$

where r denotes the Central Bank's foreign reserves and e represents the exchange rate.

The Smooth Transition Regression (STR) model is a nonlinear time-series regression model that can be regarded as an extended form of the regime-switching regression model introduced by Bacon and Watts (1971). These researchers considered two regression lines and designed a model in which the transition from one line to another occurs smoothly. In the time-series literature, Chan and Tong (1986) were the first to elaborate on and propose the Smooth Transition Regression (STR) model in their studies. Prior to them, several economists, such as Goldfeld and Quandt (1972) and Maddala (1977), had referred to this type of nonlinear models in their work. In recent years, the use of nonlinear models has become more widespread, and many researchers have sought to develop these models, among whom Teräsvirta (1998) is one of the most prominent.

A standard STR model with a logistic transition function is generally specified as follows:

$$GRO_t = \pi Z_t + (\theta Z_t) F(\gamma \cdot c \cdot s_t) + U_t \qquad (6)$$

In this model, π is the vector of coefficients associated with the linear component, and θ is the vector of coefficients associated with the nonlinear component. Z_t is a vector of exogenous variables of the model, which typically includes lags of the dependent (endogenous) variable as well as other exogenous variables (Sidney et al., 2019). As stated earlier, F is the logistic transition function that determines how the system moves from one regime to another. In most previous studies, the common form of this function is defined as:

$$F(\gamma \cdot c \cdot s_t) = \left(1 + \exp\{-\gamma \textstyle\prod_{k=1}^{k} (s_t - c)\}\right)^{-1}, \quad \gamma > 0 \qquad (7)$$

The transition function F is a continuous and bounded function between zero and one and includes the slope parameter γ and the location parameter c. The slope parameter determines the speed of transition between the two extreme regimes, while the location parameter specifies the threshold separating these regimes. To examine the properties of the STR model with a logistic transition function, following van Dijk (1999), it is assumed that the dependent variable Y depends only on its own lagged values. Under this assumption, and considering a two-regime transition function, we obtain:

$$Y_t = (\pi_0 + \pi_1 Y_{t-1} + \cdots + \pi_p Y_{t-p}) + (\theta_0 + \theta_1 Y_{t-1} + \cdots + \theta_p Y_{t-p}) F(\gamma \cdot c \cdot s_t) + U_t$$

$$F(\gamma \cdot c \cdot s_t) = 1 / (1 + \exp\{-\gamma(s_t - c)\})$$

The results of this model are referred to as a two-regime LSTR model, where the location parameter c indicates the point of transition between the two extreme regimes $F(\gamma \cdot c \cdot s_t) = 0$ and $F(\gamma \cdot c \cdot s_t) = 1$, with $F(\gamma \cdot c \cdot s_t) = 0.5$ representing the midpoint of the transition. The slope parameter γ reflects the speed of transition between regimes, with larger values indicating a faster regime shift. When $\gamma \to \infty$ and $s_t > c$, F = 1, whereas when $s_t < c$, F = 0. Accordingly, equation (1) collapses into a threshold regression (TR) model. When $\gamma \to 0$, equation (1) reduces to a linear model. For a three-regime model in which two regime changes occur, the logistic function proposed by Johansen and Teräsvirta (1996) takes the following form:

$$F(\gamma \cdot c \cdot s_t) = 1 / (1 + \exp\{(s_t - C_1)(s_t - C_2)\}), \quad C_1 \le C_2, \quad \gamma > 0$$

In this case, if $\gamma \to 0$, the model converges to a linear specification. If $\gamma \to \infty$, then for $s_t < C_1$ and $s_t > C_2$, $F(\gamma \cdot c \cdot s_t) = 1$, and if $C_1 < s_t < C_2$, then $F(\gamma \cdot c \cdot s_t) = 0$. It should be noted that F is symmetric around the point $(C_1 + C_2)/2$, never equals zero, and its minimum values lie strictly between zero and one.

Findings and Results

In this section, the explanatory assessment of the effects of exchange rate shocks and the Central Bank's monetary policy on the country's business cycles under linear and nonlinear regimes is conducted using the threshold LSTR model. In this study, annual data were collected over the period 1978–2022 and analyzed in two parts: descriptive statistics and econometric estimation. For the analyses, the compiled data obtained from the Central Bank and official statistics were used; the descriptive and econometric analyses (using the LSTR approach) were performed in EViews 12, as described below.

Table 1 reports the central tendency and dispersion indices—specifically, the mean and standard deviation—for the research variables, as detailed below.

Table 1. Descriptive statistics of the study variables

Variable name	Observations	Mean	Median	Maximum	Minimum	Std. dev.	Jarque–Bera	Prob.
Business cycles (GDPgap)	45	1,452,184	1,308,649	22,735,505	845,529.2	542,942.8	4.205	0.070
Real exchange rate (Exrv)	45	49.890	45.306	142.990	8.268	30.722	13.059	0.001
Interest rate (R)	45	8.886	8.000	22.20	6.00	2.642	75.895	0.000
Money supply (M)	45	817,865.7	40,348.26	7,823,848	2,578.6	17,191	140.25	0.000
Oil price (poil)	45	43.461	31.570	109.060	14.560	28.922	6.387	0.041
Consumer Price Index (CPI)	45	275.601	58.200	2,166.450	1.000	499.844	123.409	0.000
Oil revenue (trexoi)	45	1,460,465	79,165.719	7,938,683	1,760.73	2,196,109	22.293	

Table 1 presents the central tendency and dispersion statistics for the national business cycle variable (GDPgap) based on the 2004 base year. The mean of the national business cycles is 1,452,184 (billion rials), and the median is 1,308,649 (billion rials). The minimum and maximum values of the national business cycles over the study period are 845,529.2 and 2,273,505 (billion rials), respectively. Figure 1 illustrates the national business cycles (GDPgap) over the period 1978–2022. In the years 1983, 1993, 2007, 2011, and 2017, the Iranian economy was in expansionary phases, whereas in 1981, 1988, 2001, 2013, 2015, and 2020, it experienced recessionary phases.

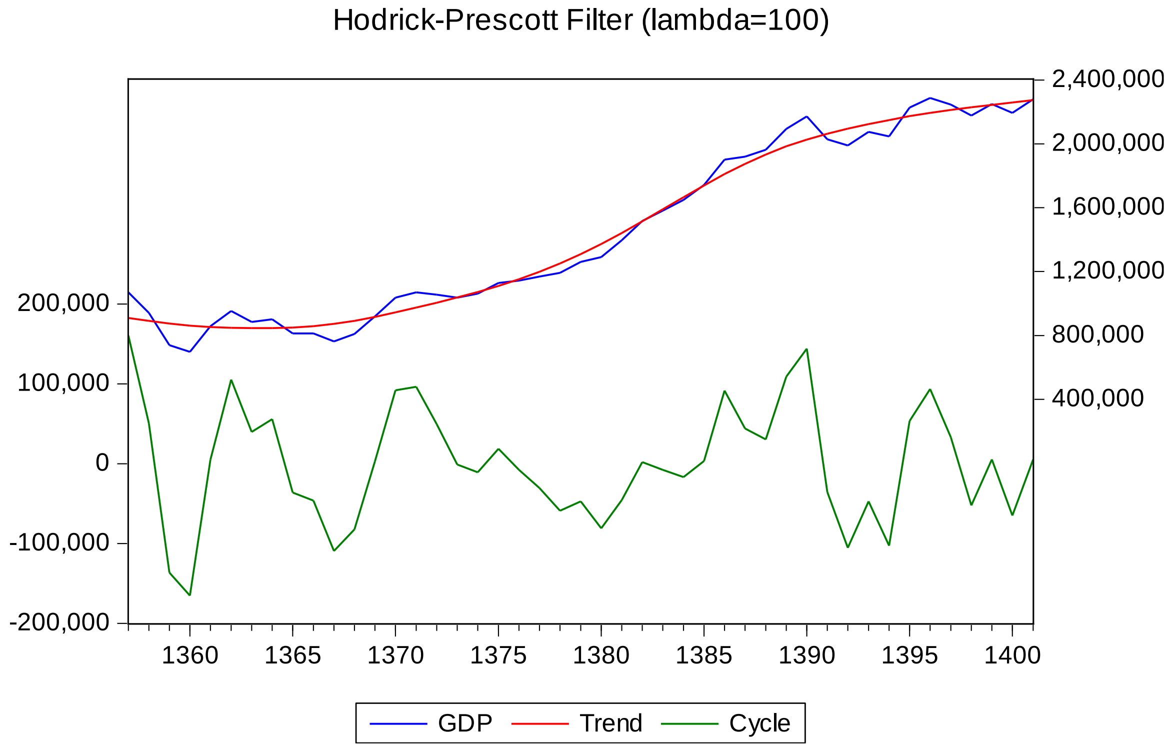

Figure 1. National business cycles (GDPgap) during 1978–2022 (billion rials) (Years equivalant: 1981-82, 1986-87, 1991-92, 1996-97, 2001-2, 2006-7, 2011-12, 2016-17, 2021-22)

Figure 1 reports the central tendency and dispersion statistics for the real exchange rate. The mean real exchange rate is 49.890%, and the median is 45.306%. The minimum and maximum real exchange rate values over the study period are 8.268% and 142.990%, respectively. Figure 2 illustrates the real exchange rate over the period 1978–2022. The highest and lowest real exchange rate levels correspond to 1992 and 1990, respectively.

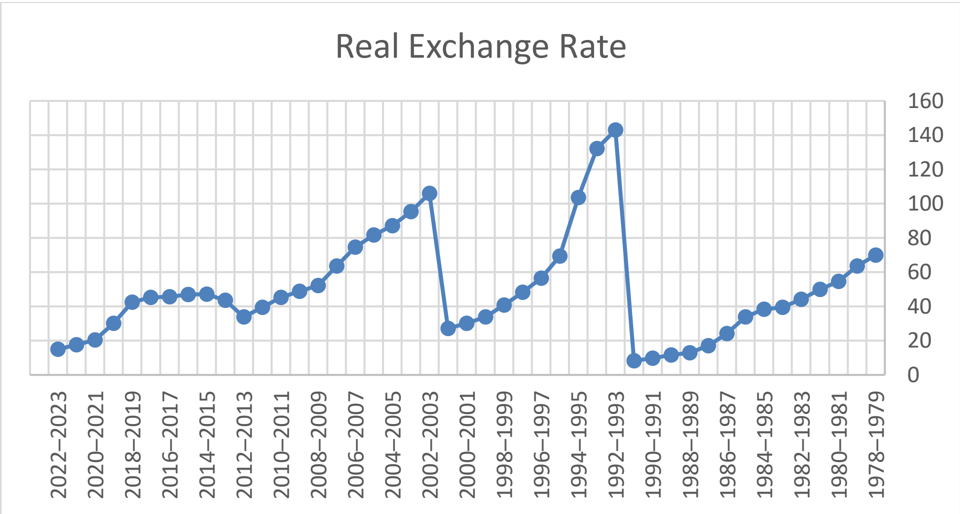

Figure 2. Exchange rate during 1978–2022 (rials)

Table 1 reports the central tendency and dispersion statistics for the short-term interest rate. The mean interest rate is 8.886%, and the median is 8.000%. The minimum and maximum interest rate values over the study period are 6.000% and 22.200%, respectively. Figure 3 illustrates the short-term interest rate over the period 1978–2022. The highest and lowest interest rate levels correspond to 2014 and 1985, respectively.

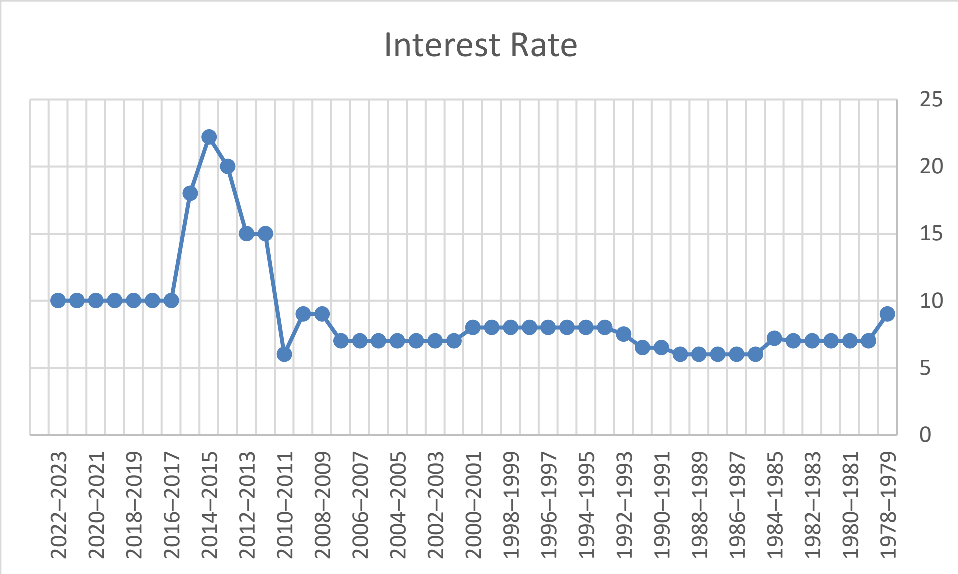

Figure 3. Interest rate during 1978–2022 (percent)

Table 1 reports the central tendency and dispersion statistics for the nominal money supply. The mean nominal money supply is 817,865.7 (billion rials), and the median is 40,348.26 (billion rials). The minimum and maximum money supply values over the study period are 2,578.6 and 7,823,848 (billion rials), respectively. Figure 4 illustrates the money supply over the period 1978–2022. The highest and lowest money supply levels correspond to 2014 and 1984, respectively.

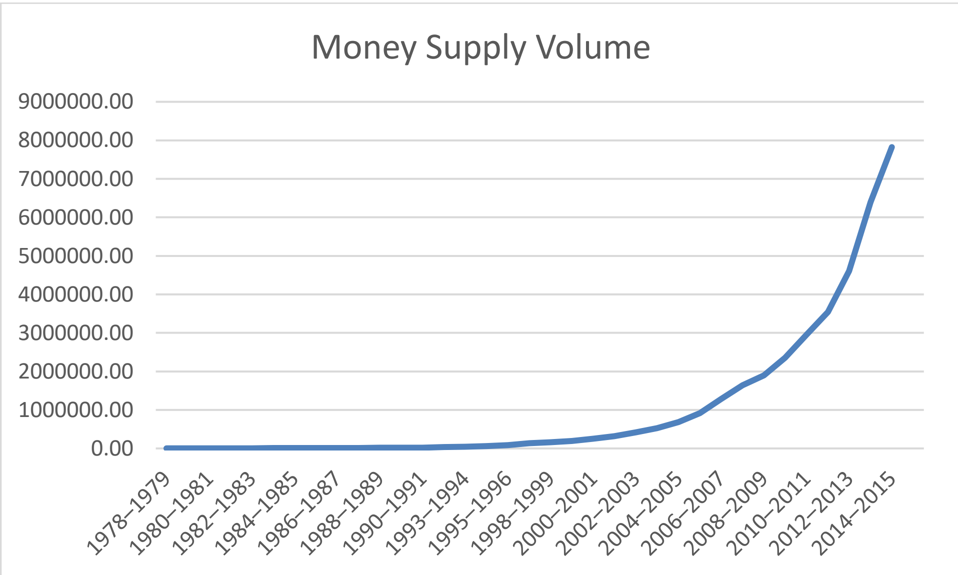

Figure 4. Money supply (M) during 1978–2022 (billion rials)

Table 1 reports the central tendency and dispersion statistics for the oil price (USD per barrel). The mean oil price is 43.461 (USD), and the median is 31.570 (USD). The minimum and maximum oil price values over the study period are 14.560 and 109.060 (USD), respectively. Figure 5 illustrates the oil price in U.S. dollars over the period 1978–2022.

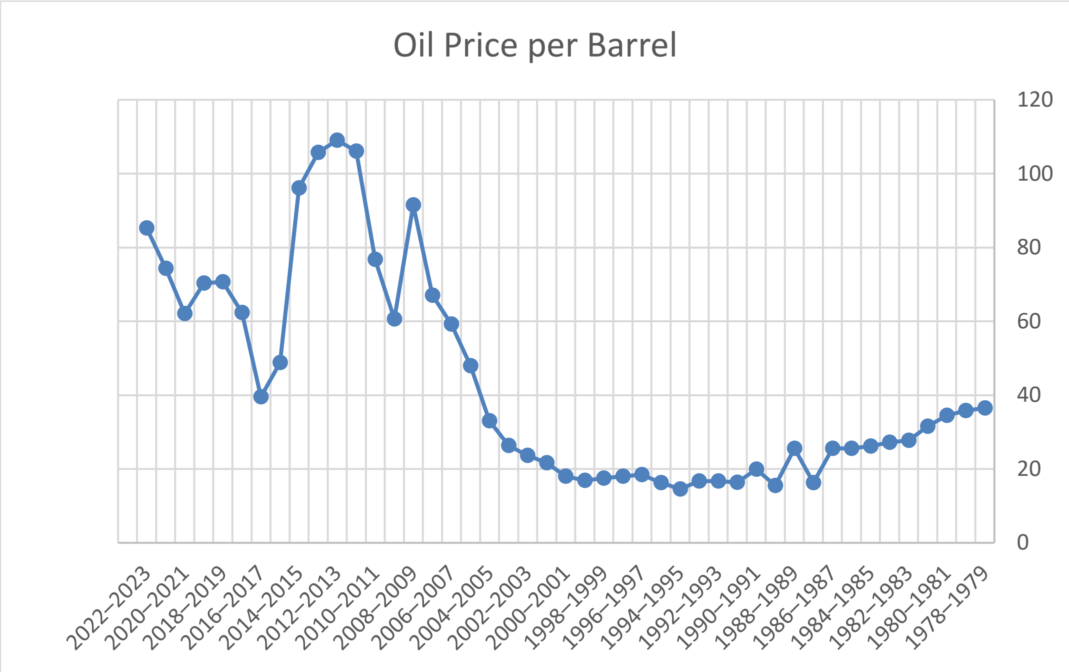

Figure 5. Oil price (poil) during 1978–2022 (USD)

Table 1 reports the central tendency and dispersion statistics for the Consumer Price Index (CPI) based on the 2004 base year. The mean CPI is 275.601, and the median is 58.200. The minimum and maximum CPI values over the study period are 1.000 and 2,166.450, respectively. Figure 6 illustrates the CPI over the period 1978–2022. The highest and lowest CPI levels correspond to 2022 and 1978, respectively.

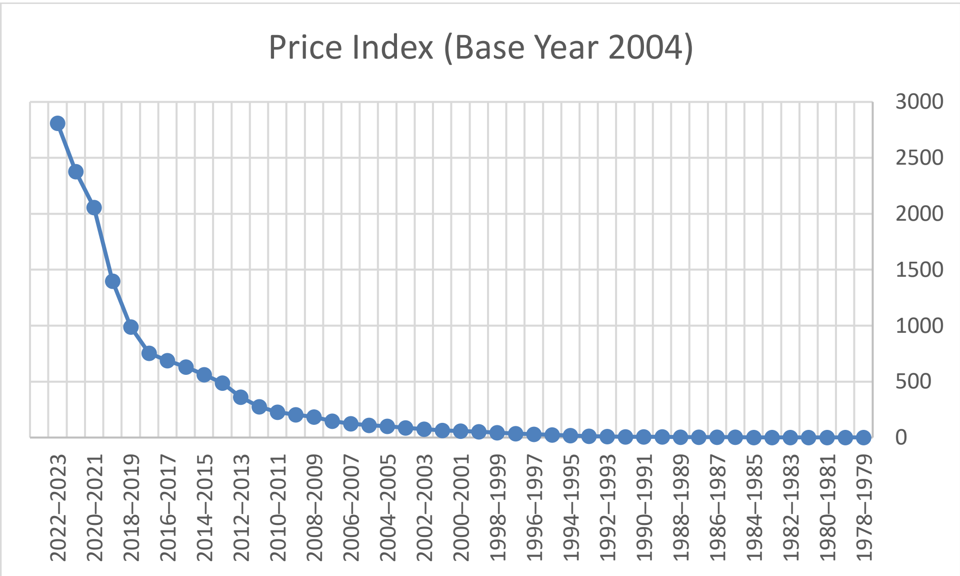

Figure 6. Consumer Price Index (CPI) during 1978–2022 (base year: 2004)

Table 1 reports the central tendency and dispersion statistics for oil revenue (in billion rials) based on the 2004 base year. The mean oil revenue is 1,460,465 (billion rials), and the median is 79,165.719 (billion rials). The minimum and maximum oil revenue values over the study period are 1,760.73 and 7,938,683 (billion rials), respectively. Figure 7 illustrates oil revenue (in billion rials) over the period 1978–2022. The highest and lowest oil revenues correspond to 2022 and 1978, respectively.

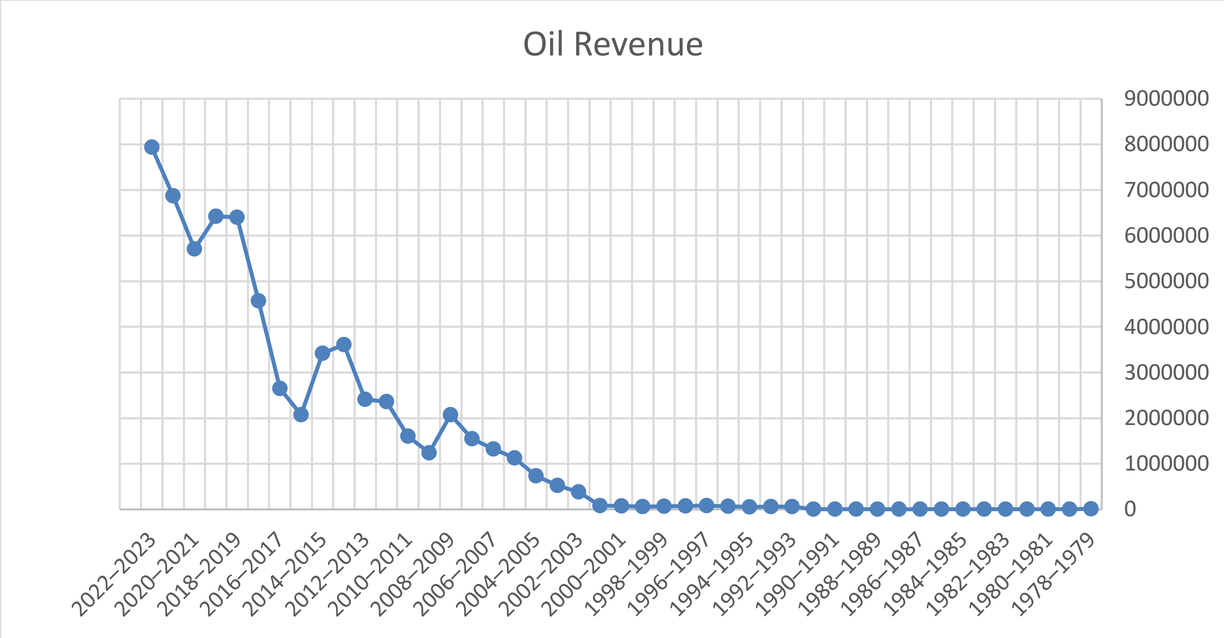

Figure 7. Oil revenue (trexoi) during 1978–2022 (billion rials)

In this study, the stationarity of the variables was first examined using the Dickey–Fuller test. The rationale for using the Dickey–Fuller test is its suitability for time-series data, because the presence of a stochastic trend in a time series can be readily diagnosed through this test, as reported in Table 2.

Table 2. Stationarity (unit-root) tests for the study variables (ADF test)

Variable name	ADF critical value (1%)	ADF critical value (5%)	ADF critical value (10%)	Test statistic (level)	Prob.	Result
Log of business cycles (GDPgap)	−4.211	−3.529	−3.196	−1.989	0.588	Non-stationary
Log of business cycles (GDPgap), first difference	−3.605	−2.936	−2.606	−3.201	0.0272	Stationary*
Log of exchange rate (LogExr)	−4.180	−3.515	−3.188	−2.513	0.320	Non-stationary
Log of exchange rate (LogExr), first difference	−4.186	−3.518	−3.189	−4.189	0.000	Stationary***
Log of interest rate (LogR)	−4.180	−3.515	−3.188	−2.764	0.217	Non-stationary
Log of interest rate (LogR), first difference	−4.186	−3.518	−3.189	−7.330	0.000	Stationary***
Log of money supply (M)	−4.243	−3.544	−3.204	−0.306	0.997	Non-stationary
Log of money supply (M), first difference	−4.243	−3.544	−3.204	−11.977	0.000	Stationary***
Log of oil price (Logpoil)	−4.180	−3.515	−3.188	−2.618	0.274	Non-stationary
Log of oil price (Logpoil), first difference	−4.186	−3.518	−3.189	−6.528	0.000	Stationary***
Log of price index (Logcpi)	−4.192	−3.520	−3.191	−0.083	0.993	Non-stationary
Log of price index (Logcpi), first difference	−2.621	−1.948	−1.611	−1.945	0.050	Stationary*
Log of oil revenue (Logtrexoi)	−4.180	−3.515	−3.188	−3.119	0.1146	Non-stationary
Log of oil revenue (Logtrexoi), first difference	−4.186	−3.518	−3.189	−6.409	0.000	Stationary***

*** $p \leq 0.001$; * $p \leq 0.05$.

Table 2 indicates that the test statistics for the study variables (log business cycles, log real exchange rate, log interest rate, log money supply, log oil price, log price index, and log oil revenue) at levels are greater than the corresponding critical values at the 1%, 5%, and 10% significance levels (i.e., the significance levels are greater than 0.05); therefore, the study variables are not stationary at level (I(0)) and must be differenced. It is observed that after one differencing (first difference), the study variables become stationary at I(1). Accordingly, it is concluded that the study variables (log business cycles, log real exchange rate, log interest rate, log money supply, oil price, log price index, and log oil revenue) are stationary at I(1).

Because the study variables are not stationary at the same level, the long-run relationships among the variables must be examined. To conduct the cointegration test, the Johansen cointegration procedure is used, and both the trace statistic and the maximum eigenvalue statistic are evaluated. In the Engle–Granger cointegration test, it is assumed that there is only one cointegrating vector; however, in a system, there may be more than one cointegrating vector, which constitutes a key limitation of the Engle–Granger approach. Therefore, to address the shortcomings of the Engle–Granger method, the Johansen method is applied. The Johansen test results in Table 3 indicate five long-run relationships based on the trace statistic and five long-run relationships based on the maximum eigenvalue statistic; thus, according to both the trace and maximum eigenvalue statistics, five long-run relationships exist among the study variables at the 0.95 confidence level, and these relationships are statistically supported.

Table 3. Johansen test results for determining the cointegration rank

Panel A. Trace statistic

H_0	H_1	Eigenvalue	Trace statistic	0.05 critical value	Prob.
R = 0*	R > 0	0.918	374.735	125.615	0.000***
R = 1*	R > 1	0.889	271.924	95.753	0.000***
R = 2*	R > 2	0.871	181.790	69.818	0.000***
R = 3*	R > 3	0.660	97.763	47.856	0.000***
R = 4*	R > 4	0.554	53.425	29.797	0.000***
R = 5*	R > 5	0.388	20.243	15.494	0.0089*
R = 6	R > 6	0.022	0.930	3.841	0.760

Panel B. Maximum eigenvalue statistic

H_0	H_1	Eigenvalue	Max-Eigen statistic	0.05 critical value	Prob.
R = 0*	R > 0	0.918	102.811	46.231	0.000***
R = 1*	R > 1	0.889	90.133	40.077	0.000***
R = 2*	R > 2	0.871	84.026	33.876	0.000***
R = 3*	R > 3	0.660	44.337	27.584	0.0002*
R = 4*	R > 4	0.554	33.182	21.131	0.0007*
R = 5*	R > 5	0.388	20.179	14.264	0.0052*
R = 6	R > 6	0.022	0.930	3.841	0.760

*** $p \leq 0.001$; * $p \leq 0.05$. Source: Estimation results of the study.

The first step in estimating a STAR model is to determine whether the model is linear or nonlinear based on the F-test statistic. If the null hypothesis of this test is rejected, linearity is rejected and nonlinearity is confirmed;

subsequently, the appropriate transition variable and the number of regimes in the nonlinear model must be determined based on the F1, F2, F3, and F4 test statistics.

Table 4. Nonlinearity test and selection of the transition variable

Null hypothesis	F statistic	Degrees of freedom	P-value
F_4: H_{04}: $\beta_1 = \beta_2 = \beta_3 = \beta_4 = 0$	6.586	(20, 16)	0.0001
F_3: H_{03}: $\beta_1 = \beta_2 = \beta_3 = 0$	6.586	(20, 16)	0.0001
F_2: H_{02}: $\beta_1 = \beta_2 = 0$	7.198	(22, 14)	0.0000
F_1: H_{01}: $\beta_1 = 0$	12.0484	(29, 7)	0.0000

Accordingly, the null hypothesis of model linearity in its general form can be specified as H_{01}: $\beta_1 = 0$. However, due to technical issues associated with testing this hypothesis, Luukkonen et al. recommend using a third-order Taylor approximation of the transition function $F(\gamma \cdot C \cdot s_t)$. The linearity null hypothesis is then tested via the Lagrange Multiplier (LM) statistic or the corresponding F-ratio under H_{03}: $\beta_1 = \beta_2 = \beta_3 = 0$. The estimation results for this stage are reported in Table 4, and based on these results, the research model is nonlinear.

In this study, the real exchange rate with lags of zero, one, two, and three periods is considered as the candidate threshold variable. The sum of squared residuals (SSR) for the real exchange rate under lag lengths of zero, one, two, and three periods is presented in Figure 8.

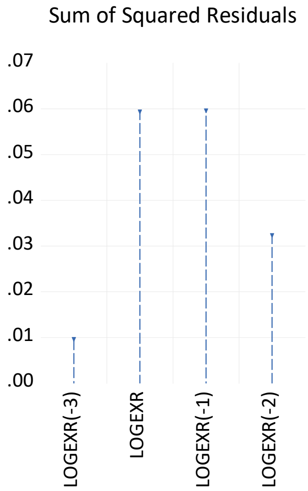

Figure 8. Sum of squared residuals (SSR)

The SSR plot indicates that the real exchange rate with a three-period lag yields the smallest SSR value (0.00968). Therefore, the real exchange rate with three lags is selected as the threshold variable.

Based on the p-values of the F-test statistics reported in Table 5, the nonlinearity assumption for the first lag of the study variable is accepted. According to the obtained results, the hypothesis of a nonlinear relationship among the study variables is confirmed.

Table 5. Type of transition-function specification

Null hypothesis	F statistic	Degrees of freedom	P-value
F_3: H_{03}: $\beta_3 = 0$	1.233	(2, 20)	0.312
F_2: H_{02}: $\beta_2 = 0 \mid \beta_3 = 0$	1.248	(7, 22)	0.320
F_1: H_{01}: $\beta_1 = 0 \mid \beta_2 = \beta_3 = 0$	12.404	(7, 29)	0.000

If H_{02} is rejected while the other two null hypotheses are accepted, an ESTAR model is selected. If H_{03} or H_{01} is rejected, the model is LSTAR. If all three null hypotheses are rejected, then, based on the p-values, the strongest rejection of the null hypothesis is considered. According to the proposed decision rule, if H_{02} is rejected most strongly, the model is ESTAR; otherwise, the model is LSTAR (Yıldırım Karaman, 2017). Based on the results reported in Table 5, the appropriate specification for the real exchange rate with three lags is selected as the smooth transition regression model with an exponential transition function (LSTAR), because H_{01} is rejected ($p = 0.000 < 0.05$), and the p-value for H_{02} is 0.320 (greater than 0.05); therefore, the estimated model is a smooth transition regression with an exponential transition function (LSTAR).

The estimation of the research model consists of two stages. In the first stage, initial values are selected for the smoothness (slope) parameter (γ) and the threshold value (C). The second stage involves the final estimation of the research specification, which is linear in C (the location parameter) and log-linear in γ (the slope parameter). The initial values are obtained through a grid search over C (location parameter) and a linear–logarithmic search over γ (slope parameter). For each combination of γ and C, the sum of squared errors is computed, and the pair of parameter values that yields the minimum sum of squared residuals (SSR) is chosen as the starting point for the algorithm. In general, for estimating the parameters of nonlinear equations, one may either minimize the residual sum of squares or maximize the likelihood function. In this study, nonlinear least squares was employed using the Newton–Raphson algorithm, and the results are reported in Table 6. Based on the reported results, most estimated coefficients are statistically significant at the 95% confidence level and are consistent with the stated theoretical foundations. Another important point is the high explanatory power of the estimated model, as indicated by the adjusted coefficient of determination of 4.007%. To address the research hypotheses regarding the capability of nonlinear regression to model and estimate the effect of the exchange rate on the country's business cycles and the superior precision of the nonlinear regression approach relative to classical regression methods, the linear LSTAR model was also estimated, as presented in Table 6.

Table 6. Model estimation results

Threshold variables (linear part)

Variable	Coefficient	Std. error	t-statistic	Prob.
Log price index	−0.0011	0.0265	−0.0429	0.9661
Log exchange rate	−0.1422	0.0245	−5.794	0.000***
Log money supply	−0.0185	0.0069	−2.672	0.0126**
Log interest rate	−0.0230	0.0338	−0.6859	0.4986
Log oil revenue	0.1615	0.0239	6.7499	0.000***
Sanctions	−0.2660	0.0613	−4.351	0.0002***
Constant	10.801	0.3684	29.316	0.000***

Threshold variables (nonlinear part)

Variable	Coefficient	Std. error	t-statistic	Prob.
Log price index	−0.2575	0.0518	−4.9650	0.0000***
Log exchange rate	0.1131	0.0670	1.671	0.106
Log money supply	0.3520	0.0690	5.0955	0.000***
Log interest rate	−0.0212	0.1300	−0.162	0.8721
Log oil revenue	0.1576	0.0610	−2.564	0.0162*
Sanctions	9.9898	16.131	0.6192	0.5409
Constant	0.2485	0.8288	0.2998	0.766

Slope / Threshold

Parameter	Estimate	Std. error	t-statistic	Prob.
SLOPE (γ)	23.4088	9.5352	2.4549	0.0208*
THRESHOLD (C)	4.0078	0.0630	63.5354	0.000***

Model estimation in Regime 2

Variable	Coefficient	Std. error	t-statistic	Prob.
Log price index	−0.2587	0.0784	−3.299***	—
Log exchange rate	−0.0291	0.0921	−0.3159	—
Log money supply	0.3337	0.0760	4.390***	—
Log interest rate	−0.0444	0.1640	0.2707	—
Log oil revenue	0.0039	0.0853	0.0457	—
Sanctions	9.723	16.193	0.600	—
Constant	11.0495	1.9721	5.603***	—

The final estimated values for the smoothness parameter (γ) and the threshold value (C) are 23.4088 and 4.0078%, respectively. Therefore, the transition function plot is as follows.

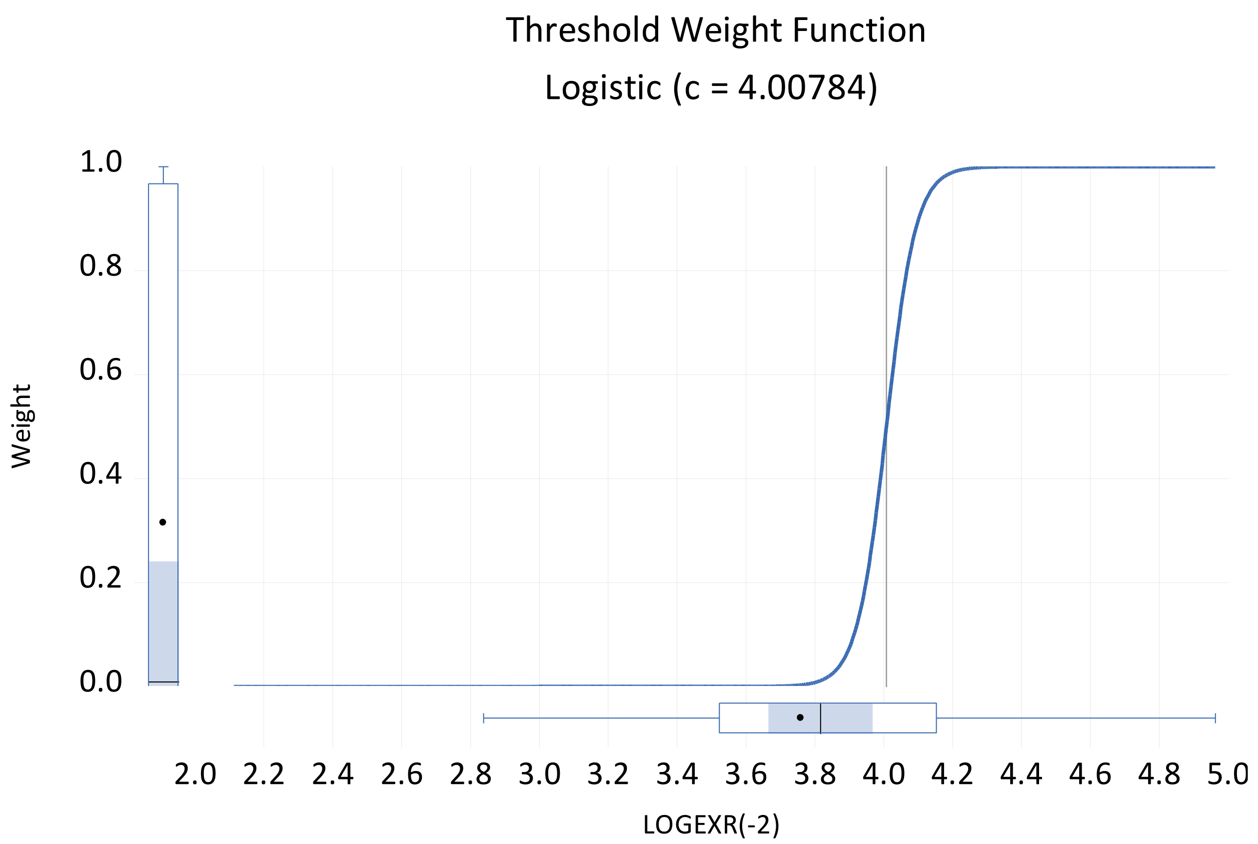

Figure 9. Transition function

In Regime 1, G = 0, and in Regime 2, G = 1. Accordingly, for Regime 1, we have:

LogGDPgap = −0.0011 logcpi − 0.1422 logexr − 0.0185 logm − 0.0232 logR + 0.161 logTroil − 0.266 Thrim + 10.80

For Regime 2, which is obtained by summing the estimated coefficients of the linear and nonlinear components, we have:

$$LogGDPgap = -0.2587\ logcpi - 0.0291\ logexr + 0.3337\ logm - 0.0444\ logR + 0.0039\ logTroil + 9.723\ hrim + 11.0495$$

Based on the estimated regression equations of the LSTAR model, the differing coefficients across the two regimes indicate that the effects of the explanatory variables vary by regime.

Changes in the log of the price index have different effects on the business cycle across the two regimes: in Regime 1, changes in the log price index exert a negative and statistically insignificant effect on the business cycle; after crossing the threshold (C = 4.007), in Regime 2, changes in the log price index have a negative and statistically significant effect on the business cycle.

Changes in the log of the exchange rate have different effects on the business cycle across the two regimes: in Regime 1, changes in the log exchange rate exert a negative and statistically significant effect on the business cycle; after crossing the threshold (C = 4.007), in Regime 2, changes in the log exchange rate exert a positive and statistically insignificant effect on the business cycle.

The effects of changes in the log of money supply differ across the two regimes: in Regime 1, changes in money supply exert a negative and statistically significant effect on the business cycle; after crossing the threshold (C = 4.007), in Regime 2, the effect of the log of money supply on the business cycle becomes positive and statistically significant.

The effects of changes in the log of the interest rate are similar across both regimes and are statistically insignificant for the business cycle.

The effects of changes in the log of oil revenue differ across the two regimes: in Regime 1, changes in the log of oil revenue exert a positive and statistically significant effect on the business cycle; after crossing the threshold (C = 4.007), in Regime 2, changes in the log of oil revenue remain positive but are not statistically significant.

Changes in sanctions conditions have different effects on the business cycle across the two regimes: in Regime 1, changes in sanctions conditions exert a negative and statistically significant effect on the business cycle; after crossing the threshold (C = 4.007), in Regime 2, changes in sanctions conditions exert a positive and statistically insignificant effect on the business cycle.

Given that the exchange rate is selected as the threshold variable, it is concluded in this section that changes in the exchange rate exert different effects on the business cycle across the two regimes.

Discussion and Conclusion

The findings of this study provide robust evidence that the effects of exchange rate shocks and central bank monetary policy on business cycles are fundamentally nonlinear and regime-dependent. The estimation results from the LSTAR model clearly indicate that macroeconomic relationships in the Iranian economy cannot be adequately captured by linear specifications, as the magnitude, direction, and statistical significance of key variables differ markedly across regimes defined by the exchange rate threshold. This result confirms the rejection of linearity and underscores the importance of threshold-based approaches in analyzing macroeconomic fluctuations in open and structurally constrained economies (9, 10).

One of the most salient results concerns the asymmetric impact of exchange rate changes on business cycles across regimes. In the low-regime state, where the exchange rate remains below the estimated threshold, exchange rate depreciation exerts a negative and statistically significant effect on the output gap, indicating that exchange rate shocks are contractionary in relatively stable or controlled conditions. This finding is consistent with theoretical arguments emphasizing cost-push effects, import dependence, and balance-sheet channels in developing and oil-dependent economies, where currency depreciation raises production costs and weakens aggregate demand (2, 4). Empirically, similar contractionary effects of exchange rate volatility on real activity have been reported for Iran and other emerging economies using nonlinear frameworks (10, 11).

However, once the exchange rate crosses the identified threshold and the economy enters the high-regime state, the effect of exchange rate changes on business cycles becomes positive but statistically insignificant. This shift suggests that beyond a certain level of exchange rate adjustment, the economy partially adapts to currency movements, possibly through price adjustments, expectations, or changes in trade behavior. Such regime-dependent behavior aligns with the notion that severe or persistent exchange rate movements may alter agents' expectations and induce structural adjustments that dampen the marginal impact of further shocks (3, 20). This finding also resonates with evidence from oil-exporting economies, where exchange rate shocks exhibit diminishing real effects once economies adjust to sustained external pressures (16).

The results related to monetary variables further reinforce the presence of nonlinear dynamics. Changes in money supply exert a negative and significant effect on business cycles in the low regime, while in the high regime the effect turns positive and significant. This reversal highlights the state-contingent effectiveness of monetary expansion. In relatively stable regimes, increases in liquidity may fuel inflationary pressures without stimulating real output, thereby worsening cyclical conditions. In contrast, during high-volatility regimes characterized by exchange rate stress, monetary expansion may alleviate liquidity constraints, support production, and mitigate recessionary pressures. These findings are consistent with the monetary transmission literature, which emphasizes that the impact of monetary policy depends critically on the prevailing macroeconomic environment (5, 7). Empirical support for such asymmetric monetary effects has been documented in studies focusing on policy uncertainty and regime shifts in Iran (19).

Interest rate effects, by contrast, appear weak and statistically insignificant across both regimes. This result suggests limited effectiveness of the interest rate channel in influencing business cycles within the Iranian context, possibly due to financial repression, administered interest rates, or shallow credit markets. Similar conclusions have been reached in studies of developing economies where structural constraints weaken the traditional interest rate transmission mechanism (1, 8). The persistence of insignificance across regimes implies that nonlinearities in this channel are less pronounced than in exchange rate and liquidity channels.

Oil-related variables play a critical role in shaping business cycles, reflecting the structural dependence of the Iranian economy on hydrocarbon revenues. The positive and significant effect of oil revenues in the low regime indicates that increases in oil income stimulate economic activity and contribute to cyclical expansion under relatively normal conditions. This finding aligns with extensive literature documenting the pro-cyclical role of oil revenues in oil-exporting countries (16, 21). However, in the high regime, the effect of oil revenues becomes statistically insignificant, suggesting that during periods of severe exchange rate instability or external pressure, the ability of oil revenues to stabilize business cycles diminishes. This result may reflect constraints imposed by sanctions, inefficiencies in revenue transmission, or increased uncertainty that weakens the expansionary impact of oil income (11, 14).

The price level also exhibits regime-dependent effects. In the low regime, changes in the price index have a negative but insignificant effect on business cycles, whereas in the high regime the effect becomes negative and statistically significant. This pattern indicates that inflationary pressures become more detrimental to real activity once the economy crosses the exchange rate threshold. Such findings are consistent with the view that inflation exacerbates macroeconomic instability in high-volatility environments, eroding purchasing power and discouraging investment (1, 22). Empirical studies in developing economies similarly report stronger contractionary effects of inflation during periods of heightened exchange rate misalignment (12, 23).

Sanctions emerge as another critical factor with asymmetric effects. In the low regime, sanctions exert a negative and significant impact on business cycles, reflecting their direct adverse effects on production, trade, and investment. In the high regime, however, the effect becomes positive but insignificant, suggesting partial adaptation or the dominance of other macroeconomic forces during periods of intense exchange rate pressure. This finding is consistent with arguments that prolonged external constraints may induce structural adjustments that reduce the marginal impact of sanctions over time, although at substantial long-term cost (11, 13).

Overall, the results of the LSTAR model provide strong empirical support for the hypothesis that exchange rate shocks and monetary policy effects on business cycles are nonlinear and state-dependent. By allowing coefficients to vary smoothly across regimes, the model captures dynamics that would be obscured in linear specifications. These findings corroborate and extend prior research emphasizing threshold effects and asymmetries in macroeconomic relationships in Iran and comparable economies (9, 10, 19). Importantly, the results also highlight the limitations of one-size-fits-all policy prescriptions and underscore the need for regime-aware macroeconomic management.

Despite its contributions, this study is subject to several limitations. First, the analysis relies on annual data, which may mask short-term dynamics and limit the ability to capture rapid adjustments in exchange rates and monetary policy. Second, the model focuses on a specific set of macroeconomic variables and does not explicitly incorporate financial market indicators or expectations-based measures that may further influence business cycles. Third, while the LSTAR framework captures smooth regime transitions, it assumes a specific functional form for the transition mechanism, which may not fully reflect more complex or abrupt structural changes in the economy.

Future studies could extend this research by employing higher-frequency data to explore short-run dynamics and policy transmission mechanisms more precisely. Incorporating additional variables such as financial development indicators, capital flows, or expectation measures could also enrich the analysis. Moreover, comparative studies across countries or regions using similar nonlinear frameworks would help assess the generalizability of the findings. Finally, exploring alternative nonlinear models, such as Markov regime-switching or time-varying parameter models, may provide complementary insights into the dynamics of exchange rate shocks and business cycles.

From a practical perspective, policymakers should explicitly account for regime-dependent effects when designing exchange rate and monetary policies. Exchange rate management strategies should recognize that the real effects of currency movements differ across states of the economy, and monetary interventions should be calibrated accordingly. Strengthening institutional frameworks, enhancing policy credibility, and improving data transparency can also help mitigate the adverse effects of exchange rate shocks. Finally, diversification away from oil dependence and the development of financial markets may reduce vulnerability to external shocks and contribute to more stable business cycle dynamics.

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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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References

1. Mankiw NG. Principles of economics. Boston, MA: Cengage Learning; 2020.
2. Obstfeld M, Rogoff K. Foundations of international macroeconomics. Cambridge, MA: MIT Press; 1996.
3. Frankel JA, Razin A. Fiscal policies in open economies. Cambridge, MA: MIT Press; 1996.
4. Frankel JA, Razin A. International economics and macroeconomic policy. Cambridge, MA: MIT Press; 1998.
5. Mishkin FS. Symposium on monetary transmission mechanism. Journal of Economic Perspectives. 1995;9(4):3-10. doi: 10.1257/jep.9.4.3.
6. Weimark T. Measuring central bank intervention: An application to the Swedish Riksbank. Scandinavian Journal of Economics. 1995;97(4):547-60. doi: 10.2307/3440582.
7. Caballero R. The “other” macroeconomic policy. American Economic Review. 2010;100(2):1-26. doi: 10.1257/aer.100.2.1.
8. Logo F. Real interest rates and private investment in developing countries. World Development. 2006;34(11):1927-38. doi: 10.1016/j.worlddev.2006.05.006.
9. Bahmani-Eskooi MR, Moshehji M. Exchange rate and monetary policies in Iran: A threshold approach. Economic Research Journal. 2020;25(1):45-70.
10. Neshat M, Yousefi S. The effect of exchange rate volatility on Iran's business cycles using an LSTR threshold model. Iranian Economic Research Quarterly. 2019;23(4):75-102.
11. Nikzad M, Yazdani M. The effects of current account and capital account changes on recession and boom cycles in Iran's economy. Iranian Economic Research Journal. 2025;19(2):45-72.
12. Tayebi M, Zamani A. Real exchange rate misalignment and its determinants in the Iranian economy. Quarterly Journal of Economic Research. 2024;21(3):55-88.
13. Abdi Seyyed Kalaei M, Mo'tameni M, Mosleh H. The effect of exchange rate fluctuations on unemployment in MENA countries. Biannual Scientific Journal of Economic Studies and Policies. 2025.
14. Rahimi S, Jafari M. Examining the Effect of Exchange Rate Fluctuations on Income Inequality in Iran. Quarterly Journal of Macroeconomics and Development. 2024(27):80-105.
15. Rezaei A, Hosseini N. The Impact of Exchange Rates on Income Inequality in Developing Countries. Quarterly Journal of Development Economics. 2024(53):120-45.
16. Helmi H. Exchange rate shocks and trade balance: Evidence from oil-exporting economies. Energy Economics. 2023;117:106517. doi: 10.1016/j.eneco.2023.106517.

17. Urgessa O. Effects of real effective exchange rate volatility on export earnings in Ethiopia: symmetric and asymmetric effect analysis. Heliyon. 2024;10. doi: 10.1016/j.heliyon.2023.e23529.

18. Rady AHA, Essam F, Yahia H, Shalaby M. The Dynamic Relationship Between Exchange Rate Volatility and Stock Prices in the Egyptian Real Estate Market and the Moderating Effect of Interest Rates. European Journal of Business Management and Research. 2024;9(5):31-44. doi: 10.24018/ejbmr.2024.9.5.2230.

19. Saeedi A, Fallahi M, Esmaeilpour Moghadam H. The Impact of Uncertainty from Monetary and Exchange Rate Policies on Financial Stability: A Markov Regime-Switching Approach in Iran's Economy. Monetary and Financial Economics. 2025. doi: 10.22067/mfe.2025.93017.1537.

20. Candian G, De Leo P. Imperfect exchange rate expectations. Review of Economics and Statistics. 2025;107(5):1406-23. doi: 10.1162/rest_a_01343.

21. Shavvalpour Arani S, Mohseni R, Kordtabar Firoozjaei Ha-D. An Asymmetric Analysis of the Impact of Oil Price, Exchange Rate, and Their Uncertainties on Unemployment in Oil-Exporting Countries. Macroeconomic Research Journal. 2024;19(2):37-70.

22. Pirpour H, Samsami Mazra'e Akhound H. Evaluating the effect of exchange rate fluctuations on core inflation using the stochastic Mundell-Fleming model and diagnosing the reciprocal relationship between exchange rate fluctuations and inflation deviations in Iran. Quantitative Economics. 2025.

23. Tejesh HR. Impact of Inflation and Exchange Rate on Stock Market Returns in India: An ARDL Approach. Theoretical & Applied Economics. 2024;31(2):25-36. doi: 10.2139/ssrn.4889476.