

Identification of Structural Monetary Policy Shocks and Evaluation of the Dynamic Response of Stock Returns: Emphasizing Investor Participation and Investor Sentiment

Davood Bazrafshan¹ and Farid Askari^{2,*}



¹ PhD Student, Department of Economic Sciences, Ab.C., Islamic Azad University, Abhar, Iran; 

² Department of Economic Sciences, Ab.C., Islamic Azad University, Abhar, Iran; 

* Correspondence: faridaskari@iau.ac.ir

Citation: Bazrafshan, D. & Askari, F. (2027). Identification of Structural Monetary Policy Shocks and Evaluation of the Dynamic Response of Stock Returns: Emphasizing Investor Participation and Investor Sentiment. *Business, Marketing, and Finance Open*, 4(8), 1-23.

Received: 18 June 2026

Revised: 04 September 2026

Accepted: 12 September 2026

Initial Publish: 20 September 2026

Final Publish: 01 August 2027



Copyright: © 2027 by the authors. Published under the terms and conditions of Creative Commons Attribution-NonCommercial 4.0 International (CC BY-NC 4.0) License.

Abstract: This study aimed to identify structural monetary policy shocks in Iran and evaluate the dynamic response of Tehran Stock Exchange returns to these shocks while considering investor sentiment, risk aversion, institutional conditions, exchange-rate dynamics, and market regimes. This quantitative time-series study used monthly data from December 2008 to December 2024. The baseline dataset comprised 193 observations, with 184 effective observations in the final VAR/SVAR estimation after applying transformations and lags. The endogenous vector included changes in annual inflation, monetary-base growth, investor sentiment, changes in a market-based risk-aversion proxy, and stock-market returns, while governance, the global financial crisis, and monthly seasonal dummies were treated as exogenous controls. Stationarity was assessed using ADF, Phillips–Perron, and KPSS tests. A VAR(3) was selected based on diagnostic adequacy, and structural monetary shocks were identified through recursive Cholesky restrictions. Dynamic effects were evaluated using Wild Bootstrap impulse responses and forecast error variance decomposition. Robustness was examined by adding exchange-rate changes, and a two-state Markov Switching model was estimated to investigate regime-dependent effects. The stock-return response to a positive structural monetary shock was initially negative, reaching its largest baseline decline around the third month, but no response from impact through 12 months was statistically significant at the 5% level. Monetary shocks explained only 1.25% of return forecast-error variance after 12 months, whereas return-specific shocks explained 90.57%. The exchange-rate robustness model yielded the same inference. Changes in risk aversion significantly predicted returns, whereas investor sentiment did not show a significant overall effect. Markov Switching identified a calm/low-return regime and a volatile/high-return regime with expected durations of 15.68 and 9.77 months, respectively, but regime differences in the effects of risk aversion, sentiment, and monetary shocks were not significant at 5%. Structural monetary policy shocks had a limited direct effect on Tehran Stock Exchange returns, while market-specific shocks and changes in risk pricing were more influential, and nonlinear market behavior was expressed mainly through differences in return levels, volatility, and regime persistence rather than through significant regime-dependent monetary effects.

Keywords: Structural Monetary Policy Shock; Stock Returns; Investor Sentiment; Risk Aversion; SVAR; Markov Switching; Tehran Stock Exchange

1. Introduction

Stock markets play a central role in modern financial systems by allocating capital, transmitting information about expected economic conditions, and converting changes in macroeconomic policy into changes in asset prices and required rates of return. Because stock prices reflect the discounted value of expected future cash flows, any

policy intervention that changes financing conditions, inflation expectations, discount rates, liquidity, or perceptions of economic risk can potentially alter stock returns. Monetary policy is therefore one of the major macroeconomic forces through which financial-market valuations may be affected. Yet the response of stock returns to monetary policy is neither necessarily immediate nor uniform. The magnitude and even the direction of the response can depend on the nature of the policy innovation, the state of the financial market, investors' expectations, their willingness to bear risk, and the institutional environment in which policy is implemented. This complexity has increasingly shifted empirical research away from simple correlations between observed monetary aggregates and stock prices toward the identification of genuinely unexpected monetary policy shocks and the assessment of their dynamic effects through structural time-series models [1-4].

A fundamental issue in this literature is that an observed change in a monetary-policy variable cannot automatically be interpreted as a monetary policy shock. Changes in money supply, monetary base, or policy interest rates may partly reflect systematic responses by monetary authorities to inflation, economic activity, exchange-rate pressures, or financial conditions. If these predictable or endogenous responses are treated as exogenous shocks, the estimated stock-market effect may be contaminated by reverse causality and simultaneous macroeconomic adjustments. Structural vector autoregressive models address this identification problem by separating reduced-form innovations into economically interpretable structural disturbances subject to explicit contemporaneous restrictions. Such methods have been used to examine the interaction between monetary policy, exchange rates, and asset prices and have shown that the response of equity markets may differ substantially depending on the identification strategy, the policy measure employed, and the economic environment [3, 5, 6]. In Iran specifically, earlier studies have documented meaningful links among monetary developments, exchange-rate movements, stock-market cycles, and stock prices, while also emphasizing that these relationships may be nonlinear, asymmetric, or time varying [7-9].

The dynamic transmission of monetary policy to equity returns operates through several channels. A monetary expansion can reduce financing costs, support liquidity, raise nominal demand, and alter expectations about corporate profitability; however, it can simultaneously generate inflationary or exchange-rate pressures that change discount rates and increase uncertainty. Conversely, a contractionary monetary disturbance can reduce liquidity and increase required returns, yet it may also strengthen confidence in price stability. Consequently, the sign of the stock-market response is ultimately an empirical question. Evidence from different markets confirms that monetary-policy actions can generate heterogeneous responses across periods and market conditions, while the effect may change when policy uncertainty or unexpected policy surprises are explicitly considered [4, 10]. Research on the Tehran Stock Exchange similarly indicates that monetary policy is connected to stock-market movements, but that the strength of this relationship depends on the specification and state of the market [2, 6, 9].

One reason for this heterogeneity is that monetary policy does not affect markets independently of investor behavior. Financial prices emerge from the interaction between macroeconomic information and the way investors interpret that information. Investor sentiment captures broad fluctuations in optimism and pessimism that may not be fully explained by contemporaneous fundamentals. Such sentiment can influence trading intensity, willingness to hold risky assets, and the speed with which information is incorporated into stock prices. A large international literature shows that investor sentiment contains relevant information for subsequent stock returns and that its effect may vary across markets, horizons, and states of the economy [11, 12]. Multi-scale evidence further indicates that the sentiment–return relationship can change across investment horizons, suggesting that short-term behavioral pressures and longer-term valuation dynamics should not be treated as identical phenomena [13]. In

the Iranian market, evidence of sentiment-related and momentum-driven behavior has likewise been reported, particularly in relation to herding and emotionally motivated trading [14, 15].

The interaction between monetary policy and investor sentiment is especially important because identical policy innovations may generate different stock-market consequences when investors are optimistic compared with when they are pessimistic. Time-varying analyses have shown that sentiment can modify the stock-return response to monetary-policy shocks, meaning that the policy transmission mechanism is partly behavioral rather than purely macroeconomic [16]. Cross-country evidence also indicates that stock markets may react differently to conventional and unconventional monetary policies across sentiment regimes [17]. More recent firm-level evidence confirms that monetary shocks and sentiment can interact in shaping equity returns, reinforcing the argument that a policy shock should not be interpreted independently of prevailing investor expectations [18]. These findings imply that an analysis focused only on monetary variables and stock returns may conceal an important behavioral component of policy transmission.

A second behavioral dimension concerns risk aversion. Investor sentiment and risk aversion are related but conceptually distinct. Sentiment reflects optimism, pessimism, and expectations about market prospects, whereas risk aversion concerns the compensation investors require for bearing uncertainty. Changes in risk appetite can therefore influence expected returns even when beliefs about future fundamentals remain unchanged. Time variation in risk appetite and economic uncertainty has been shown to be substantial, with important implications for asset valuation and financial-market dynamics [19]. Similarly, evidence on time-varying risk aversion demonstrates that shifts in the price of risk can help predict future returns [20]. Risk aversion also affects how investors respond to market liquidity and changing investment opportunities, strengthening the link between financial conditions and expected stock returns [21]. The relation between risk compensation and returns may itself depend on investor sentiment, further illustrating the interaction between behavioral mechanisms [22].

Recent studies increasingly treat sentiment and risk aversion jointly rather than as interchangeable concepts. Evidence suggests that sentiment can alter the conventional risk–return relationship, implying that the return compensation associated with risk depends partly on prevailing behavioral conditions [23]. In nonlinear models, risk aversion, investor sentiment, and monetary-policy shocks have all been identified as potentially relevant determinants of stock returns, but their effects need not remain constant over time [24]. This issue is especially relevant for the Iranian capital market, where evidence has indicated a nonlinear relationship between risk aversion and stock returns [25]. Accordingly, treating risk aversion as a fixed structural characteristic of investors or assuming that it affects returns identically in all market conditions may impose an unrealistic restriction on the empirical model.

These behavioral considerations lead naturally to the possibility of regime dependence. Stock markets often alternate between relatively calm periods and episodes of unusually high volatility, rapid price increases, or sharp reversals. In such circumstances, a single linear model may average across economically distinct states and conceal important nonlinearities. Two-state Markov-switching models provide a useful framework because they allow parameters, conditional means, or variances to differ across latent market regimes without imposing *ex ante* classifications. Early applications demonstrated that the relationship between volatility and equity returns can differ across statistically identified regimes [26]. Regime-switching approaches have also been applied to interactions between exchange-rate changes and stock returns, showing that macro-financial transmission can depend on the underlying market state [27]. When monetary policy is incorporated, regime-based analyses further suggest that the response of asset prices can change according to prevailing inflation or market conditions [10].

These findings motivate a framework in which monetary shocks, sentiment, and risk aversion are evaluated not only on average but also across endogenously identified states of the stock market.

The institutional environment represents another dimension through which stock-market responses may vary. Governance quality can influence investor confidence, property-right protection, regulatory predictability, information transparency, corruption risk, and the credibility of economic policy. Cross-country evidence indicates that governance quality is related to equity-market risk and can help predict stock-market returns [28, 29]. More recent evidence suggests that stronger country-level governance can enhance equity-market stability by reducing vulnerability to destabilizing shocks [30]. Within Iran, governance has also been connected to competitiveness and the broader functioning of the Tehran Stock Exchange [31]. Governance therefore provides an important institutional control when evaluating the effect of monetary and behavioral variables, particularly because apparently similar policy shocks may be interpreted differently in environments with different levels of regulatory quality, political stability, government effectiveness, and rule of law.

External financial disturbances further complicate the identification of domestic policy effects. The global financial crisis demonstrated that emerging and frontier stock markets can be affected through international financial linkages, volatility spillovers, and changing correlations with global markets even when the domestic origin of the crisis is limited. Evidence from emerging Middle Eastern and North African markets documented substantial changes in return and volatility linkages during the global financial crisis [32]. For a study beginning in late 2008, controlling explicitly for the crisis period is therefore necessary to avoid attributing externally generated market disturbances entirely to domestic monetary conditions, investor behavior, or institutional factors. This is especially important because the earliest segment of the sample coincides with an extraordinary episode in international financial markets.

Exchange-rate dynamics are equally relevant in the Iranian context. Monetary expansion, inflation expectations, and exchange-rate depreciation can be closely interconnected, while exchange-rate changes may affect listed firms differently according to import dependence, export exposure, foreign-currency liabilities, and the domestic valuation of real assets. Iranian evidence has long shown that monetary policy, exchange rates, gold prices, and stock-market behavior interact dynamically [7]. Structural evidence also supports important linkages among monetary-policy shocks, exchange-rate shocks, and stock prices in Iran [3]. International evidence from regime-switching structural models similarly indicates that the relationship between exchange-rate movements and equity returns may vary across states [27]. Consequently, even when exchange-rate data limitations prevent the variable from entering the longest baseline sample, its inclusion in a shorter robustness specification is important for evaluating whether an estimated monetary-policy effect is actually capturing an omitted currency-market channel.

Despite substantial progress in the literature, several gaps remain when the Iranian capital market is considered. First, many studies examine observed monetary indicators rather than maintaining a strict distinction between changes in monetary conditions and exogenous structural monetary shocks. Second, monetary-policy transmission, investor sentiment, and risk aversion are often examined separately, although evidence increasingly suggests that they interact dynamically [16, 18, 24]. Third, linear models can obscure the possibility that return behavior differs sharply between calm and turbulent periods. Fourth, institutional conditions and external episodes such as the global financial crisis are not always integrated explicitly into the same empirical framework. Finally, the Iranian literature has generated important evidence on monetary-policy effects, stock-market cycles, risk aversion, governance, investor behavior, and nonlinear dynamics, but these elements have rarely been combined within a

single framework that first identifies structural monetary innovations and then evaluates their dynamic and regime-dependent transmission [5, 9, 25].

The present study addresses these issues by combining structural identification with behavioral and regime-switching analysis. Monthly stock returns are modeled jointly with changes in inflation, monetary-base growth, investor sentiment, and changes in a market-based risk-aversion proxy, while governance quality, the global financial crisis, and calendar effects are treated as external controls. A structural VAR is used to isolate the unexpected component of monetary-base growth and trace the dynamic response of stock returns through impulse-response functions and forecast-error variance decomposition. Because stock-market volatility and behavioral relationships may change across time, a two-state Markov Switching model is additionally used to identify latent market regimes and test whether the associations of risk aversion, sentiment, and structural monetary shocks with returns differ across those regimes. This combined design follows the broader empirical movement toward time-varying, nonlinear, and behaviorally informed analyses of asset pricing and policy transmission [17, 23, 26], while adapting the framework to the specific monetary, institutional, and financial characteristics of the Iranian market.

Accordingly, the aim of this study is to identify structural monetary policy shocks in Iran and evaluate the dynamic response of Tehran Stock Exchange returns to these shocks while examining the roles of investor sentiment, risk aversion, institutional conditions, exchange-rate dynamics, and market regimes.

2. Methods and Materials

This study employed a quantitative, applied, and time-series research design to identify exogenous structural monetary policy shocks and evaluate the dynamic response of stock market returns to such shocks, with particular emphasis on the roles of investor participation and investor sentiment. The empirical framework was designed to distinguish an unexpected monetary policy innovation from systematic policy responses to contemporaneous macroeconomic and financial conditions. Accordingly, the unit of analysis consisted of consecutive time-series observations for the Iranian economy and capital market rather than individual investors or listed firms. The statistical sample included all observations for which a consistent set of monetary, macroeconomic, stock-market, investor-participation, and investor-sentiment variables was simultaneously available. A monthly frequency was adopted because it provides an appropriate compromise between capturing relatively rapid financial-market responses to monetary innovations and maintaining adequate consistency in macroeconomic and monetary data. The principal stock-market variable was the return on the broad stock-market index, calculated as the logarithmic change in the index between two consecutive periods. This specification transformed the index level into a stationary return measure and allowed the estimated impulse responses to be interpreted in terms of changes in market performance following an unexpected monetary policy disturbance.

The empirical design treated monetary policy shocks as unexpected components of monetary-policy decisions after accounting for the information available to monetary authorities at the time of policy formulation. Depending on the operational characteristics of monetary policy in Iran, the monetary-policy stance was represented by an observable policy-related monetary or interest-rate indicator capable of reflecting contractionary and expansionary policy conditions. Rather than interpreting every observed change in this variable as an exogenous policy shock, the study isolated its structural innovation within a multivariate system containing macroeconomic and financial information. This distinction was essential because observed monetary-policy variables are endogenous and may themselves react to inflation, economic conditions, liquidity developments, or financial-market fluctuations. The baseline system therefore incorporated the monetary-policy indicator together with variables representing price

conditions, real economic activity where sufficiently reliable high-frequency data were available, liquidity or monetary conditions, stock-market returns, investor participation, and investor sentiment. Parsimonious specification was maintained to preserve degrees of freedom and avoid the inclusion of variables that were conceptually unrelated to the identification of monetary shocks or the transmission of those shocks to the stock market.

Investor participation was incorporated as a market-activity dimension reflecting the degree to which investors were actively involved in stock-market transactions. It was operationalized using an observable trading-based indicator, such as the number or intensity of active transactions, trading participation, market turnover, or another consistently available measure reflecting investors' actual market engagement. Investor sentiment was treated separately because participation and sentiment represent conceptually distinct aspects of investor behavior. Whereas participation reflects observable engagement in trading, sentiment captures the prevailing degree of optimism or pessimism among market participants. The inclusion of both constructs made it possible to determine whether the transmission of monetary-policy shocks to stock returns was associated not only with changes in fundamental financing conditions but also with changes in investors' willingness to participate and their collective expectations about market prospects. The analysis was conducted over the maximum common sample period permitted by the simultaneous availability and comparability of all variables, and observations with inconsistent definitions or structural discontinuities attributable solely to changes in data-reporting procedures were harmonized before estimation.

The study relied exclusively on secondary quantitative data obtained from official monetary, macroeconomic, and capital-market databases. Monetary and macroeconomic variables were extracted from official statistical publications and time-series databases of the Central Bank of the Islamic Republic of Iran and, where necessary for complementary macroeconomic information, other official Iranian statistical sources. Capital-market data, including the broad market index, transaction information, trading activity, and other variables required to construct stock returns and investor-participation indicators, were obtained from official Tehran Stock Exchange and Iranian capital-market information systems. All raw series were matched by observation date and converted to a common monthly frequency before estimation. When an underlying series was available at a higher frequency, monthly values were constructed according to the economic meaning of the variable; end-of-period values were used for stock variables when appropriate, while averages or monthly aggregates were used for flow and activity variables. No questionnaire or interview instrument was used because investor participation and investor sentiment were measured from observable market data rather than self-reported perceptions.

Stock-market return was calculated as the natural logarithmic difference of the market index, multiplied by 100 when percentage interpretation was required. Monetary and macroeconomic level variables characterized by substantial scale differences were expressed in logarithmic form where economically meaningful, while rate variables were retained in their original units unless transformation was necessary to achieve statistical adequacy. Investor participation was constructed from trading information so that higher values indicated stronger investor involvement in the market. To remove the mechanical effect of changes in total market size, the participation measure was normalized when required using variables such as market capitalization, total eligible accounts, total trading activity, or another relevant scaling factor. Investor sentiment was constructed as a market-based proxy derived from observable indicators associated with speculative demand, trading intensity, market breadth, turnover, valuation pressure, or related manifestations of investor optimism and pessimism. Where several sentiment-related indicators were simultaneously available, they were standardized before aggregation so that

differences in measurement scale did not dominate the resulting sentiment measure. The direction of each component was adjusted so that larger values consistently represented more optimistic sentiment. If a composite sentiment index was required, dimensionality-reduction procedures such as principal component analysis were used to extract the common variation among the standardized sentiment proxies rather than assigning arbitrary subjective weights.

Before econometric estimation, the assembled database was screened for missing observations, inconsistent frequencies, extreme recording errors, and changes in variable definitions. Nominal macroeconomic series were converted into real terms where their economic interpretation required adjustment for changes in the general price level. Variables displaying strong deterministic seasonality were seasonally adjusted when the original source did not already provide adjusted data. Logarithmic transformations, growth rates, or first differences were applied according to the statistical properties and economic interpretation of each series. However, transformations were not performed mechanically; the final treatment of each variable was selected jointly on the basis of time-series properties and the need to preserve the interpretation of the structural monetary-policy shock and the corresponding stock-market response.

Data analysis was carried out in several sequential stages. Initially, descriptive statistics and time-series plots were examined to identify the distributional characteristics, volatility patterns, potential outliers, and possible structural changes in the variables. The stationarity properties of the series were subsequently evaluated using conventional unit-root procedures, including the augmented Dickey–Fuller and Phillips–Perron tests, with appropriate deterministic components selected according to the characteristics of each variable. Because conventional unit-root tests may have limited power in the presence of major economic disruptions or regime shifts, the stability of the series and model parameters was also considered when interpreting the preliminary results. Variables were transformed when necessary to obtain a statistically coherent multivariate system, while avoiding unnecessary differencing that could eliminate economically meaningful information concerning the transmission of monetary policy.

The core empirical analysis was based on a Structural Vector Autoregressive model. In reduced form, the vector of endogenous variables was modeled as a function of its own lagged values, deterministic components where required, and a vector of reduced-form innovations. The optimal lag length was selected by jointly considering standard information criteria, particularly the Akaike information criterion, Schwarz Bayesian criterion, and Hannan–Quinn criterion, together with residual-diagnostic results and the need to maintain model parsimony. The selected specification was then subjected to stability and residual-diagnostic tests to verify that the estimated dynamic system was suitable for structural interpretation. Serial correlation of the residuals was assessed through multivariate autocorrelation tests, and the stability condition was examined by verifying that the inverse roots of the characteristic polynomial were located within the unit circle.

Identification of the monetary-policy shock constituted the central econometric stage. The structural representation was imposed on the covariance structure of the reduced-form innovations so that an innovation in the monetary-policy equation could be interpreted as an unexpected policy disturbance rather than an endogenous response to contemporaneous economic or financial developments. The identification restrictions were determined by economic theory and the timing of information transmission among macroeconomic, monetary, and financial variables. Slow-moving macroeconomic variables were assumed not to respond fully within the same observation period to an unexpected monetary-policy innovation, whereas financial-market variables, including stock returns, investor participation, and investor sentiment, were permitted to react more rapidly. At the same time, the

monetary authority was allowed to respond contemporaneously to the macroeconomic information assumed to be observable when policy decisions were made. This recursive or theory-consistent structural ordering separated systematic monetary-policy behavior from the orthogonal structural innovation interpreted as the monetary-policy shock. Where necessary, alternative theoretically plausible orderings were estimated as robustness checks to ensure that the conclusions were not mechanically generated by a single contemporaneous ordering assumption.

After identifying the structural monetary-policy innovation, impulse response functions were estimated to trace the dynamic effects of a standardized monetary-policy shock on stock returns over subsequent periods. The impulse responses indicated the direction, magnitude, timing, persistence, and eventual dissipation of the stock-market reaction. Confidence intervals around the responses were obtained through simulation or bootstrap procedures so that statistically meaningful responses could be distinguished from sampling uncertainty. Particular attention was paid to the initial reaction of stock returns, the period in which the maximum response occurred, the persistence of the effect, and the horizon over which the response returned toward its pre-shock equilibrium. Forecast error variance decomposition was also employed to determine the proportion of future uncertainty in stock returns attributable to structural monetary-policy shocks relative to other innovations contained in the system.

To address the study's emphasis on investor participation and investor sentiment, their dynamic responses to the same identified monetary-policy shock were estimated jointly with the response of stock returns. This approach made it possible to assess whether monetary disturbances were followed by systematic changes in market participation and investor sentiment and whether these behavioral-market variables evolved in a manner consistent with the observed response of stock returns. The analysis therefore distinguished the direct dynamic response of equity returns from the accompanying adjustment of investor behavior. Additional specifications incorporated participation and sentiment separately and jointly in the structural system to examine the robustness of the estimated stock-return response to their inclusion. Comparison of impulse responses across these specifications provided evidence regarding whether investor participation and sentiment materially altered the magnitude or persistence of monetary-policy transmission to the equity market. Robustness analyses further considered alternative lag structures, reasonable alternative proxies for the monetary-policy stance, participation, and sentiment, and alternative structural identification restrictions. Statistical inference was conducted at conventional significance levels, while substantive interpretation emphasized the economic magnitude and temporal profile of the estimated responses rather than relying exclusively on statistical significance.

3. Findings and Results

The empirical analysis was based on monthly observations from December 2008 through December 2024. The underlying database contained 193 monthly observations, although the effective sample varied across variables because monthly returns, lag-based behavioral proxies, first differences, and model lags necessarily removed observations from the beginning of the sample. These reductions followed directly from variable construction and were not caused by discretionary deletion of observations. The empirical framework distinguished observed monetary expansion from a structural monetary policy shock. In particular, monthly monetary-base growth was treated as an observed monetary variable, whereas the monetary policy shock was defined as the unexpected orthogonal innovation identified subsequently in the Structural Vector Autoregressive model. The main stock-market variable was the monthly logarithmic return on the Tehran Stock Exchange Price and Dividend Index (TEDPIX), calculated as

$$RET_t = 100 \times \ln \left(\frac{TEDPIX_t}{TEDPIX_{t-1}} \right).$$

Because computation of the first return required the preceding month's index, the return series began in January 2009 and contained 192 observations. Inflation was measured primarily as the twelve-month percentage change in the consumer price index,

$$INF_t = 100 \times \left(\frac{CPI_t}{CPI_{t-12}} - 1 \right),$$

while the monetary variable was expressed as the monthly logarithmic growth rate of reserve money,

$$MPG_t = 100 \times [\ln(MB_t) - \ln(MB_{t-1})].$$

The distinction between MPG_t and the structural monetary shock is fundamental: MPG_t contains systematic, seasonal, endogenous, and unexpected monetary movements, whereas the structural shock represents only the orthogonal unexpected component extracted from the SVAR.

Investor sentiment was represented by a reproducible market-based proxy constructed exclusively from information available before month t , thereby avoiding mechanical contemporaneous correlation with current stock returns. It combined volatility-adjusted short-run momentum with the proportion of positive-return months:

$$SENT_t = 0.5 \left[\frac{RET_{t-3:t-1}}{SD(RET_{t-6:t-1})} \right] + 0.5 \left[2 \left(\frac{N(RET > 0)}{N} \right) - 1 \right].$$

A minimum of three preceding observations was required, making April 2009 the first month for which SENT could be calculated. The first component captures recent return momentum relative to its volatility, whereas the second distinguishes periods dominated by positive rather than negative returns. Risk aversion was also represented by a backward-looking market-based proxy based on the relationship between historical mean return and return variance. Up to 24 preceding monthly returns were used, with an expanding window at the start of the sample and a minimum of four prior returns. Consequently, the first usable risk-aversion observation was obtained in May 2009. The proxy was interpreted as a market-implied measure of the price of risk rather than a direct psychological measure of individual utility. All information entering both SENT and RA preceded the current month.

Institutional quality was represented by a composite World Governance Indicators measure constructed as

$$GOV_y = \frac{VA_y + PV_y + GE_y + RQ_y + RL_y + CC_y}{6},$$

where the components respectively represented Voice and Accountability, Political Stability and Absence of Violence, Government Effectiveness, Regulatory Quality, Rule of Law, and Control of Corruption. Annual values were repeated across the corresponding months and treated as slow-moving exogenous institutional information. The global financial crisis was controlled through an exogenous binary indicator equal to one from December 2008 through December 2009 and zero otherwise. Because usable return observations began in January 2009, 12 crisis observations remained in return-based models. The free-market exchange rate was retained for robustness analysis rather than the baseline SVAR because continuous coverage began only in late 2011. Its monthly logarithmic change was calculated from monthly averages of daily exchange-rate observations.

The descriptive statistics indicate pronounced heterogeneity in the distributions of the financial, behavioral, monetary, and macroeconomic variables. Monthly stock returns averaged 3.002% with a standard deviation of 8.568%, while the observed range extended from -14.086% to 47.211%. The return distribution was strongly positively skewed and leptokurtic. Monetary-base growth also displayed marked positive skewness and exceptionally high kurtosis. Twelve-month inflation averaged 25.389%, while its first difference was centered very close to zero. SENT was the only main variable for which the Jarque–Bera test did not reject normality at the conventional 5% level. RA itself was mildly negatively skewed, whereas its first difference displayed substantial excess kurtosis. Exchange-rate changes were likewise positively skewed and fat-tailed.

Table 1. Descriptive Statistics of the Study Variables

Variable	N	Mean	Median	SD	Minimum	Maximum	Skewness	Kurtosis	Jarque–Bera	p-value
RET	192	3.002	2.111	8.568	-14.086	47.211	1.741	9.610	423.911	<0.001
MP Growth	193	1.718	1.699	2.899	-6.659	22.001	2.371	19.846	2336.074	<0.001
INF	193	25.389	25.064	13.977	5.303	54.540	0.192	1.805	12.655	0.0018
Δ INF	192	0.020	-0.131	2.451	-11.119	10.693	0.077	7.053	122.954	<0.001
SENT	189	0.306	0.215	0.649	-2.124	2.301	0.218	3.773	5.565	0.0619
RA	188	5.651	5.311	4.441	-12.770	17.351	-0.213	4.237	12.160	0.0023
Δ RA	187	0.027	-0.012	1.564	-6.238	9.704	0.878	13.480	829.567	<0.001
GOV	193	-1.073	-1.090	0.147	-1.288	-0.805	0.575	2.171	16.125	0.0003
$\Delta \ln(\text{EXR}) \times 100$	157	2.567	1.210	6.871	-13.672	34.414	1.126	6.941	125.977	<0.001

The descriptive pattern of stock returns revealed substantial episodic volatility. The highest monthly return, 47.21%, occurred in April 2020, whereas the lowest, -14.09%, occurred in October 2020. Monetary-base growth ranged from -6.66% in January 2013 to 22.00% in March 2009. Inflation reached its minimum of approximately 5.30% in December 2016 and its maximum of 54.54% in April 2023. SENT ranged from -2.124 in April 2009 to 2.301 in February 2018. RA ranged from -12.77 in January 2016 to 17.35 in June 2018. For the exchange-rate subsample, the strongest monthly depreciation was approximately 34.41% in September 2018, whereas the largest appreciation was approximately -13.67% in December 2018. These patterns were consistent with the high kurtosis observed in the descriptive results and indicated that volatility was concentrated in specific episodes rather than being constant throughout the sample.

Formal seasonality tests were subsequently applied because predictable calendar movements in the monetary variable could otherwise contaminate the innovation interpreted as a structural monetary policy shock. One-way ANOVA and Kruskal–Wallis tests produced no evidence of statistically significant monthly seasonality in stock returns, Δ INF, SENT, or Δ RA. Monetary-base growth, however, exhibited a strong calendar pattern under both tests. The ANOVA statistic was 6.895 with $p < 0.000001$, while the Kruskal–Wallis statistic was 39.145 with $p = 0.000050$. Accordingly, 11 monthly dummy variables were included as exogenous controls in the baseline VAR/SVAR, with January serving as the reference month. This specification retained the original monetary data while removing predictable calendar effects from the residual innovation.

Unit-root testing was conducted using the Augmented Dickey–Fuller, Phillips–Perron, and KPSS procedures. RET and SENT were stationary in levels. Annual inflation was clearly nonstationary, while its first difference was stationary under all three tests. RA produced conflicting level results, with the ADF and KPSS tests failing to support level stationarity; therefore, its first difference was used. Monetary-base growth generated mixed evidence because ADF and PP strongly rejected a unit root whereas KPSS marginally rejected stationarity. Because this variable was already a logarithmic growth rate and exhibited a demonstrable seasonal pattern, it was retained in

growth form and seasonality was controlled explicitly rather than imposing a second difference. Exchange-rate log changes were stationary.

Table 2. Seasonality and Stationarity Results

Variable	ANOVA p-value	Kruskal–Wallis p-value	ADF	ADF p-value	PP	PP p-value	KPSS	KPSS p-value	Final Treatment
RET	0.2621	0.1160	−10.046	<0.001	−10.242	<0.001	0.078	0.100	Level
MP Growth	<0.000001	0.000050	−3.527	0.0073	−15.279	<0.001	0.501	0.0414	Growth form + monthly dummies
INF	—	—	−1.461	0.5525	−1.898	0.3328	0.804	0.010	First difference
ΔINF	≈1.000	≈1.000	−4.469	0.0002	−6.736	<0.001	0.059	0.100	Stationary
SENT	0.9410	0.9831	−3.612	0.0055	−5.703	<0.001	0.144	0.100	Level
RA	—	—	−2.398	0.1424	−3.089	0.0273	0.656	0.0176	First difference
ΔRA	0.5756	0.9433	−10.621	<0.001	−11.468	<0.001	0.112	0.100	Stationary
Δln(EXR) × 100	—	—	−8.754	<0.001	−8.540	<0.001	0.079	0.100	Level

The resulting baseline endogenous vector was therefore specified as

$$Y_t = [\Delta INF_t \quad MPG_t \quad SENT_t \quad \Delta RA_t \quad RET_t]'$$

Lag selection considered one through six monthly lags. All four information criteria favored a first-order VAR. However, the residual-whiteness test rejected the null of no residual serial correlation for VAR(1), with $p = 0.0037$, and for VAR(2), with $p = 0.0294$. VAR(3) was the first specification for which the whiteness null was not rejected at the 5% level, $p = 0.0915$. Although VAR(4) also passed this diagnostic, it used more parameters and had less favorable information criteria. Accordingly, VAR(3) was selected as the most parsimonious specification that adequately captured the serial dynamics.

Table 3. VAR Lag Selection and Diagnostic Adequacy

Lag	AIC	BIC	HQIC	FPE	Whiteness p-value	Stable
1	5.7012	7.3487	6.3688	300.2871	0.0037	Yes
2	5.7105	7.7993	6.5570	304.2382	0.0294	Yes
3	5.7189	8.2524	6.7457	308.6272	0.0915	Yes
4	5.7472	8.7287	6.9557	320.2165	0.0982	Yes
5	5.9214	9.3543	7.3131	385.7435	0.0361	Yes
6	6.0967	9.9844	7.6729	467.0239	0.0019	Yes

The final VAR(3) used 184 effective monthly observations after accounting for behavioral-variable construction, differencing, and three lags. GOV, GFC, and the 11 monthly dummy variables were included as exogenous controls. Its general form was

$$Y_t = c + A_1 Y_{t-1} + A_2 Y_{t-2} + A_3 Y_{t-3} + B X_t + u_t,$$

where X_t contained the exogenous institutional, crisis, and monthly-seasonality controls. The system satisfied the stability condition. The minimum modulus of the characteristic roots was 1.3385; equivalently, the largest inverse-root modulus was approximately

$$\frac{1}{1.3385} \approx 0.7471 < 1.$$

The 12-lag Portmanteau/Whiteness statistic was 253.7223 with $p = 0.0915$, so residual serial correlation was not detected at the conventional 5% level. Multivariate normality, however, was strongly rejected, with a test statistic of 1698.614 and $p < 0.001$. ARCH-LM testing with six lags showed no significant conditional heteroskedasticity in the ΔINF equation ($LM = 4.767, p = 0.5740$), the monetary-growth equation ($LM = 2.966, p = 0.8131$), or the SENT equation ($LM = 3.803, p = 0.7033$). Significant ARCH effects were detected for ΔRA ($LM = 31.922, p < 0.001$) and RET ($LM = 50.042, p < 0.001$). Consequently, inferential evaluation of impulse responses relied on Wild Residual Bootstrap confidence intervals rather than solely on Gaussian asymptotic intervals.

Structural identification was performed recursively using the ordering

$$\Delta INF_t \rightarrow MPG_t \rightarrow SENT_t \rightarrow \Delta RA_t \rightarrow RET_t.$$

The relationship between reduced-form innovations and orthogonal structural shocks was represented as

$$u_t = P\varepsilon_t,$$

where P was the lower-triangular Cholesky factor of the reduced-form innovation covariance matrix. Equivalently,

$$\varepsilon_t = P^{-1}u_t.$$

Under this ordering, ΔINF did not respond contemporaneously to shocks occurring later in the system, monetary-base growth could respond contemporaneously to ΔINF but not to SENT, ΔRA , or RET, SENT could respond contemporaneously to inflation and monetary innovations, ΔRA could react to inflation, monetary conditions, and sentiment, and stock returns were permitted to respond contemporaneously to all preceding structural shocks. The second element of ε_t , rather than observed monetary-base growth or the raw reduced-form monetary residual, was defined as the structural monetary policy shock:

$$MPShock_t = \varepsilon_{MP,t}.$$

Thus, conceptually,

$$MPG_t = \text{predictable component} + \text{seasonal component} + \text{systematic policy response} + \text{unexpected component},$$

while

$$MPShock_t = \text{orthogonal unexpected structural monetary innovation}.$$

For illustration, estimated monetary shocks were approximately 0.100 in September 2018, 0.390 in October 2018, and -0.518 in November 2018, demonstrating that the structural innovation need not move proportionately with observed monetary-base growth.

The impulse-response analysis showed that stock returns initially tended to respond negatively to a positive structural monetary innovation. The strongest baseline point response occurred around the third month and was approximately -0.8275 percentage points. Nevertheless, its 95% Wild Bootstrap interval, approximately [-1.7537, 0.3238], contained zero. More generally, none of the confidence intervals from horizon 0 through month 12 excluded zero. The point responses therefore suggested an initially negative but transitory adjustment, whereas the bootstrap evidence did not establish a statistically significant direct return response at the 5% level.

Forecast error variance decomposition reinforced this conclusion. At a one-month horizon, 97.47% of return forecast-error variance was explained by the return equation's own structural innovation and only 0.06% by the structural monetary shock. The monetary contribution increased gradually but remained small, reaching approximately 1.18% at six months and 1.25% at twelve months. Changes in inflation and risk aversion accounted for somewhat larger shares at medium horizons, but the return-specific innovation remained overwhelmingly dominant throughout.

Table 4. Forecast Error Variance Decomposition of Stock Returns

Horizon	ΔINF	Monetary Shock	SENT	ΔRA	Return-Specific Shock
1 month	0.87%	0.06%	1.07%	0.53%	97.47%
3 months	1.00%	0.28%	0.98%	2.71%	95.02%
6 months	3.20%	1.18%	1.13%	3.45%	91.04%
12 months	3.55%	1.25%	1.16%	3.47%	90.57%

The limited direct role of the structural monetary shock was also supported by a robustness specification explicitly incorporating exchange-rate movements. Because continuous free-market exchange-rate observations were available only from December 2011 onward, the robustness analysis used the December 2011–December 2024 subsample and an expanded endogenous vector,

$$Y_t^R = [\Delta INF_t \quad MPG_t \quad \Delta EXR_t \quad SENT_t \quad \Delta RA_t \quad RET_t]'$$

The recursive ordering placed exchange-rate changes immediately after the monetary variable, allowing the exchange rate to respond contemporaneously to the monetary shock while retaining stock returns in the final position. GOV and monthly dummy variables remained exogenous. GFC was omitted because it was identically zero throughout the shorter subsample. A VAR(2) was selected for this robustness specification.

The response remained initially negative after explicitly controlling for the exchange-rate channel. It was -0.2985 on impact, -0.4731 after one month, and reached its largest negative value of -0.5148 after two months. It became slightly positive at month four and converged rapidly toward zero thereafter. Crucially, every 95% confidence interval from impact through month 12 contained zero.

Table 5. Robustness IRF of Stock Returns to the Structural Monetary Shock with the Exchange Rate Included

Horizon	RET Response	95% Lower Bound	95% Upper Bound
0	-0.2985	-0.8712	0.3257
1	-0.4731	-1.7060	0.8702
2	-0.5148	-1.4392	0.5450
3	-0.0699	-0.6608	0.6141
4	0.2382	-0.1902	0.6758
5	0.0951	-0.2489	0.3950
6	-0.0309	-0.3043	0.1747
7	-0.0190	-0.1898	0.1261
8	-0.0006	-0.1074	0.1272
9	-0.0068	-0.0880	0.0897
10	-0.0088	-0.0801	0.0607
11	-0.0052	-0.0634	0.0466
12	-0.0033	-0.0446	0.0333

These results indicate that the weak direct monetary-shock effect was not simply an artifact of excluding exchange-rate dynamics. Although the precise timing of the minimum response differed from the baseline model, statistical inference was unchanged: no horizon from impact through one year produced a 95% confidence interval excluding zero.

The possibility of nonlinear stock-market dynamics was then examined using a two-state Markov Switching model. The effective regime-switching sample ran from September 2009 through December 2024 and contained 184 monthly observations. Regimes were not assigned *ex ante* labels such as expansion or recession. Instead, their interpretation followed estimation and was based on conditional return means, volatility, persistence, and smoothed-state probabilities. The first-order transition process was defined as

$$P(S_t = j \mid S_{t-1} = i) = p_{ij}.$$

Continuous regressors were standardized before estimation, so regime-specific coefficients represented the change in stock returns associated with a one-standard-deviation change in the corresponding regressor.

The two regimes differed sharply in their return and volatility characteristics. Regime 0 comprised 120 months, or 65.22% of the effective sample, with an average return of 0.439% and an observed standard deviation of 4.519%. Regime 1 comprised 64 months, or 34.78%, and exhibited a substantially higher mean return of 7.714% and a standard deviation of 11.978%. Model-implied conditional variances were 18.179 and 127.768, corresponding to conditional standard deviations of 4.264 and 11.303, respectively. On this basis, Regime 0 was interpreted as a calm/low-return state and Regime 1 as a volatile/high-return state.

Persistence was high in both states. The transition matrix was

$$P = \begin{bmatrix} 0.9362 & 0.0638 \\ 0.1024 & 0.8976 \end{bmatrix}.$$

The expected duration of state i was therefore

$$D_i = \frac{1}{1 - p_{ii}}.$$

Hence,

$$D_0 = \frac{1}{1 - 0.9362} = 15.68 \text{ months},$$

and

$$D_1 = \frac{1}{1 - 0.8976} = 9.77 \text{ months}.$$

The calm/low-return state was therefore both more frequent and more persistent. Smoothed probabilities also produced meaningful temporal clustering rather than erratic monthly switching. For example, the probability of belonging to the volatile/high-return state increased from approximately 0.4695 in October 2024 to 0.9817 in November and 0.9882 in December 2024. Conversely, an individual positive-return month did not automatically imply assignment to the high-return regime; October 2009, for example, produced a return of approximately 7.89% while retaining a smoothed probability of approximately 0.8197 for the calm state.

The regime-specific coefficients provided weaker evidence of changes in the behavioral and monetary relationships than did the regime-specific mean and variance parameters. In the calm state, ΔRA had a positive and

statistically significant coefficient of 1.5893 ($p = 0.0032$). Its coefficient in the volatile regime remained positive at 1.5321 but was not statistically significant ($p = 0.3056$). Critically, the direct cross-regime coefficient-difference test yielded $p = 0.9730$, demonstrating that significance in one regime and nonsignificance in the other did not constitute evidence of a statistically different coefficient. SENT was -0.4991 ($p = 0.3516$) in the calm regime and 2.5696 ($p = 0.0902$) in the volatile regime. The difference test yielded $p = 0.0637$, providing only marginal evidence at the 10% level and not at 5%. The structural monetary-shock coefficient was -0.3815 ($p = 0.3680$) in the calm state and 0.1415 ($p = 0.9265$) in the volatile state, with no significant cross-regime difference ($p = 0.7461$).

Table 6. Markov-Switching Regime Characteristics and Regime-Dependent Effects

Indicator	Calm/Low-Return Regime	Volatile/High-Return Regime
Months	120	64
Share of sample	65.22%	34.78%
Mean monthly return	0.439%	7.714%
Observed return SD	4.519%	11.978%
Model conditional variance	18.179	127.768
Model conditional SD	4.264	11.303
Probability of remaining in regime	0.9362	0.8976
Expected duration	15.68 months	9.77 months
ΔRA coefficient	1.5893	1.5321
ΔRA p-value	0.0032	0.3056
SENT coefficient	-0.4991	2.5696
SENT p-value	0.3516	0.0902
Monetary-shock coefficient	-0.3815	0.1415
Monetary-shock p-value	0.3680	0.9265

Formal coefficient-difference tests reinforced this interpretation. The cross-regime difference for ΔRA was 0.0572 ($z = 0.0338$, $p = 0.9730$); the SENT difference was -3.0687 ($z = -1.8539$, $p = 0.0637$); and the monetary-shock difference was -0.5230 ($z = -0.3238$, $p = 0.7461$). A complementary comparison between a model in which ΔRA , SENT, and the monetary-shock coefficients were all allowed to vary across regimes and a restricted model in which they were common across regimes reached the same conclusion. The unrestricted model had a log-likelihood of -617.876 , AIC of 1265.751, and BIC of 1313.975, whereas the restricted model had a log-likelihood of -620.144 , a lower AIC of 1264.289, and a substantially lower BIC of 1302.868. The auxiliary likelihood-ratio statistic was 4.5376 with three degrees of freedom and $p = 0.2090$. Thus, regime switching was pronounced in the conditional level and variance of stock returns, but the data did not require statistically distinct coefficients for the three principal explanatory variables at the 5% level.

The institutional and crisis controls were also statistically insignificant. In the Markov Switching specification, the standardized GOV coefficient was approximately -0.3904 with a standard error of 0.4120, $z = -0.948$, and $p = 0.3432$. The GFC coefficient was approximately -0.8332 with a standard error of 2.7112, $z = -0.307$, and $p = 0.7586$. Because the Markov sample retained only September–December 2009 from the crisis period, an auxiliary HAC regression was estimated over the broader January 2009–December 2024 return sample. GOV remained negative but statistically insignificant, $b = -6.5464$, $HAC SE = 4.1161$, $p = 0.1117$, and GFC was also negative and insignificant, $b = -1.1700$, $HAC SE = 2.2858$, $p = 0.6087$. Monetary-base growth ($b = 0.0189$, $p = 0.8839$) and ΔINF ($b = 0.2880$, $p = 0.3737$) were likewise insignificant in this auxiliary specification. The auxiliary regression was not intended as a comprehensive return model; its role was to assess whether the lack of significant GOV and GFC effects persisted in the longer sample.

Granger-predictive tests added a further distinction between behavioral variables. Lagged ΔRA significantly contributed to predicting stock returns, with a test statistic of 9.0831 and $p = 0.0282$. In contrast, lagged SENT did not provide significant predictive content for returns, and the corresponding probability was approximately 0.843. Monetary growth likewise showed no significant predictive contribution, with $p = 0.6433$. These findings should be interpreted as predictive rather than structural causality.

The combined hypothesis evidence is summarized below. The strongest support concerned the association between changes in the market-based risk-aversion proxy and subsequent returns, although that result remains conditional on the proxy definition and does not establish structural causality. SENT did not exhibit a statistically significant overall effect at 5%, despite marginal evidence in the volatile regime. The structural monetary policy shock produced neither a statistically significant baseline IRF nor a significant exchange-rate-augmented IRF, and its 12-month FEVD share was only 1.25%. GOV and GFC were also statistically insignificant. Structural shocks other than the return-specific innovation collectively explained an increasing but still minority share of forecast-error variance, rising from 2.53% at one month to 9.43% at twelve months. Finally, although the Markov model clearly identified two distinct stock-return regimes, the effects of ΔRA , SENT, and the structural monetary shock did not differ significantly across those regimes at the 5% level.

Table 7. Summary of Empirical Hypothesis Tests

Hypothesis	Principal Evidence	Statistical Result	Conclusion
Risk-aversion changes are significantly associated with stock returns	Granger test; Markov coefficient	Granger $p = 0.0282$; calm-regime $p = 0.0032$	Conditionally supported
Investor sentiment significantly affects stock returns	Granger test; regime coefficients	Granger $p \approx 0.843$; volatile-regime $p = 0.0902$	Not supported at 5%
Structural monetary policy shocks significantly affect stock returns	Bootstrap IRF; FEVD; robustness model	No 95% IRF excludes zero; 12-month FEVD = 1.25%	Not supported
Governance significantly affects stock returns	Markov and HAC estimates	$p = 0.3432$; HAC $p = 0.1117$	Not supported
Global financial crisis significantly affected stock returns	Markov and HAC estimates	$p = 0.7586$; HAC $p = 0.6087$	Not supported
Structural shocks explain part of return forecast-error variance	FEVD	Non-return shocks = 2.53% at 1 month and 9.43% at 12 months	Supported
Risk-aversion effect differs across market regimes	Wald coefficient-difference test	$p = 0.9730$	Not supported
Sentiment effect differs across market regimes	Wald coefficient-difference test	$p = 0.0637$	Not supported at 5%; marginal at 10%
Monetary-shock effect differs across market regimes	Wald coefficient-difference test	$p = 0.7461$	Not supported

Taken together, the empirical results identify three central patterns. First, the structural monetary policy innovation had a small and statistically indistinguishable-from-zero direct effect on Tehran stock-market returns over horizons up to twelve months. The corresponding FEVD contribution remained modest, and explicit inclusion of exchange-rate changes did not alter this inference. Second, stock-return dynamics displayed substantial state dependence in their conditional mean, volatility, and persistence: the calm/low-return regime accounted for 65.22% of observations and lasted an expected 15.68 months, whereas the volatile/high-return regime accounted for 34.78% and lasted approximately 9.77 months. Third, this pronounced regime dependence in market conditions did not translate into statistically robust regime dependence in the coefficients of risk aversion, sentiment, or structural monetary policy shocks. Among the behavioral measures, changes in the market-based risk-aversion proxy provided the clearest predictive information for returns, while SENT generated only marginal evidence in the

volatile regime. The dominant quantitative result of the SVAR remained the large contribution of return-specific innovations, which explained 90.57% of return forecast-error variance even at a twelve-month horizon.

4. Discussion and Conclusion

The present study examined the dynamic response of Tehran Stock Exchange returns to structurally identified monetary policy shocks while simultaneously accounting for investor sentiment, changes in a market-based risk-aversion proxy, governance conditions, the global financial crisis, exchange-rate dynamics, and possible regime switching in stock-market behavior. The empirical results provide a differentiated picture of monetary and behavioral transmission in the Iranian capital market. The principal finding is that a positive structural monetary policy shock generated an initially negative response in stock returns, with the strongest point response occurring within the first few months, but this response was not statistically significant at the 5% level at any horizon from impact through twelve months. Forecast error variance decomposition reinforced this result: the monetary shock explained only 0.06% of return forecast-error variance after one month and approximately 1.25% after twelve months, while return-specific innovations accounted for more than 90% even at the one-year horizon. The same conclusion survived the robustness model that explicitly incorporated exchange-rate changes. At the same time, changes in the risk-aversion proxy showed significant predictive content for returns, whereas investor sentiment did not display a statistically robust overall effect. The Markov Switching analysis identified two clearly differentiated market states in terms of mean return, volatility, and persistence, but the coefficients of risk aversion, sentiment, and monetary shocks did not differ significantly between regimes at the conventional 5% level.

The weak direct response of stock returns to the structural monetary policy shock is one of the most important findings of the study. This result does not imply that monetary policy is irrelevant to the Iranian stock market. Rather, it indicates that after predictable monetary behavior, seasonal movements, contemporaneous macroeconomic interactions, and other systematic components were removed, the remaining unexpected monetary innovation did not produce a statistically distinguishable direct return response over the examined horizons. This distinction is important because observed monetary growth and structural monetary shocks are not equivalent. Earlier studies have shown that the estimated stock-market response to monetary policy depends substantially on how the policy innovation is identified and whether systematic policy responses are separated from unexpected shocks [1, 4]. Iranian evidence has likewise demonstrated significant interactions among monetary conditions, exchange rates, and stock prices, but these relationships are sensitive to model specification, temporal variation, and policy asymmetry [3, 5, 6, 9]. The present results are therefore consistent with the view that monetary transmission to equity returns may be indirect, time varying, and conditional rather than mechanically strong in every period.

The shape of the estimated impulse response is also economically meaningful. Stock returns responded negatively in the early months following a positive monetary innovation, but the effect diminished and converged toward zero. Such a pattern can arise when monetary expansion initially increases uncertainty, inflation expectations, or concerns about future macroeconomic instability rather than being interpreted purely as a liquidity stimulus. In the Iranian setting, where monetary developments are closely linked to inflation and exchange-rate pressures, the same expansionary innovation may simultaneously increase nominal liquidity and worsen expectations regarding the purchasing power of money. This dual mechanism can weaken or even reverse the conventional positive asset-price response expected from easier monetary conditions. Previous Iranian studies have similarly reported that the stock-market effects of monetary policy can be asymmetric across market cycles and

depend on broader macro-financial conditions [2, 8, 9]. Time-varying evidence also suggests that the effect of policy shocks on financial markets is not stable across periods, which provides a plausible explanation for the absence of a strong average response in a long sample spanning several distinct episodes [5, 6].

The forecast error variance decomposition further clarifies the limited quantitative importance of the direct monetary channel. Even after twelve months, the monetary shock accounted for only about 1.25% of return forecast-error variance, whereas approximately 90.57% was attributable to return-specific innovations. This finding indicates that a large proportion of monthly stock-market fluctuations originated from forces not directly represented by the identified monetary, inflation, sentiment, or risk-aversion shocks. Such forces may include firm-specific information, political developments, regulatory interventions, changes in market microstructure, commodity-price movements, sectoral disturbances, trading restrictions, and unobserved expectations. The dominance of idiosyncratic or market-specific shocks is consistent with studies showing that monetary variables explain only part of equity-return variation and that stock markets remain highly responsive to financial and behavioral forces beyond standard macroeconomic fundamentals [1, 24]. The result is particularly plausible in an emerging market such as Iran, where institutional, regulatory, and political developments can generate large price movements independently of contemporaneous monetary conditions.

The robustness model containing exchange-rate changes produced essentially the same inferential conclusion. Although the initial point response of stock returns remained negative, none of the confidence intervals excluded zero. This finding is important because exchange rates are among the most relevant potential transmission channels linking Iranian monetary conditions to asset prices. Earlier studies have established strong dynamic relationships between monetary policy, exchange rates, and equity-market behavior in Iran [3, 7]. Regime-based evidence from other emerging markets also shows that exchange-rate changes can interact with stock returns in nonlinear ways [27]. The persistence of the insignificant monetary response after exchange-rate inclusion suggests that the baseline result was not merely an artifact of omitting this channel. Instead, the direct effect of the identified monetary innovation appears genuinely weak relative to other sources of return variation.

A second major finding concerns risk aversion. Lagged changes in the market-based risk-aversion proxy significantly improved the explanation of future stock returns in the VAR framework, and the Markov model showed a positive and significant coefficient in the calm/low-return regime. This result aligns with the broader asset-pricing literature emphasizing that time variation in investors' willingness to bear risk is central to expected-return dynamics. Risk appetite and uncertainty are known to vary considerably over time, and shifts in required compensation for risk can affect both valuations and subsequent returns [19]. Similarly, time-varying risk aversion has been shown to contain predictive information for future equity returns [20]. The present finding is also consistent with evidence that the relationship between liquidity, returns, and risk appetite depends on fluctuations in investors' required compensation for uncertainty [21]. In Iran, the nonlinear relationship previously reported between risk aversion and stock returns provides additional support for the relevance of this behavioral-financial channel [25].

At the same time, the positive coefficient on the change in risk aversion should not be interpreted mechanically as evidence that greater psychological risk aversion causes higher stock prices. The proxy used in this study is market based and reflects the relationship between past returns and return variance; therefore, it is more accurately interpreted as an indicator of changes in the market price of risk than as a direct measurement of individual utility curvature. Under this interpretation, a positive association with subsequent returns is plausible because higher required compensation for bearing risk may be followed by higher realized returns. This interpretation accords

with studies linking risk compensation to equity returns and showing that the return required for risk bearing is partly conditioned by investor sentiment and market circumstances [22, 23]. The finding therefore supports the proposition that behavioral risk pricing contains information relevant to future stock-market performance.

Investor sentiment produced weaker evidence. In the linear VAR, lagged SENT did not significantly predict returns. In the Markov Switching model, the sentiment coefficient was insignificant in the calm regime and positive but only marginally significant at the 10% level in the volatile/high-return regime. This pattern is broadly consistent with literature showing that sentiment effects are state dependent and may become more pronounced in periods of heightened market activity or instability. International evidence demonstrates that investor sentiment can explain stock returns, but the magnitude of this relationship varies across countries and time horizons [11-13]. Other studies show that sentiment interacts with the risk-return relation rather than exerting a uniform independent effect [22, 23]. The comparatively stronger coefficient observed in the volatile regime is therefore compatible with the argument that sentiment matters more when uncertainty, trading intensity, and behavioral reactions are elevated.

The Iranian literature also provides support for this interpretation. Prior research has documented a role for sentiment and momentum in herding behavior on the Tehran Stock Exchange and has linked emotional investor decisions to the occurrence of market crises [14, 15]. These studies suggest that behavioral forces may become particularly influential when market conditions depart from normality. However, the present results indicate that the evidence is not strong enough to establish a robust sentiment-return effect at the 5% level across the full sample. One reason may be measurement. The sentiment variable employed here was deliberately constructed from lagged market information to avoid contemporaneous mechanical correlation with returns. Although methodologically preferable, such a conservative market-based proxy may capture only part of the broader psychological component reflected in survey-based expectations, trading-account flows, news tone, or search behavior. Thus, the limited statistical effect should be interpreted as evidence about this specific market-based proxy rather than as a general rejection of behavioral sentiment mechanisms.

The interaction between sentiment and monetary transmission also deserves attention. Previous research has found that the stock-market response to monetary-policy shocks can vary systematically with investor sentiment [16]. Cross-country evidence further shows that conventional and unconventional monetary policies generate different equity responses across sentiment regimes [17]. More recent evidence at the firm level similarly confirms that monetary policy shocks and investor sentiment jointly influence stock returns [18]. In the present study, however, neither the structural monetary shock nor its regime-specific coefficient was statistically significant. This divergence may reflect differences in market structure, monetary-policy implementation, data frequency, or sentiment measurement. In particular, the Iranian monetary system relies less heavily on a transparent single policy-rate signal than many advanced economies, which may weaken the speed with which investors translate monetary surprises into stock-price revisions.

The Markov Switching results provide strong evidence that Tehran stock returns themselves are nonlinear. The calm/low-return state accounted for approximately 65% of the effective sample and was characterized by an average monthly return of only 0.439% and a conditional standard deviation of about 4.264%. In contrast, the volatile/high-return state produced an average monthly return of 7.714% and a conditional standard deviation of approximately 11.303%. Both regimes were highly persistent, with expected durations of 15.68 and 9.77 months, respectively. These results accord with the broader literature demonstrating that stock-market returns and volatility are often better represented by regime-switching processes than by a single invariant distribution [26]. They are also consistent with evidence that asset-price dynamics and macro-financial transmission can change across latent

regimes [10, 27]. The high persistence found here further indicates that the regimes represent sustained market environments rather than transitory outliers.

Importantly, however, the existence of distinct return regimes did not imply that the explanatory coefficients of risk aversion, sentiment, and monetary shocks differed significantly between states. Direct Wald tests showed no significant regime difference for risk aversion or the monetary shock, while the sentiment difference was only marginal at the 10% level. The restricted Markov specification with common coefficients was also preferred by both AIC and BIC, and the auxiliary likelihood-ratio test did not reject coefficient equality. This result supports an important distinction: nonlinear behavior may be concentrated in the conditional mean, volatility, and persistence of returns rather than in every structural slope coefficient. Dahmene et al. similarly emphasize that nonlinearity in stock returns can arise through complex interactions among risk aversion, sentiment, and monetary shocks, but not every channel must switch equally strongly across regimes [24]. The present evidence suggests that regime switching is especially relevant for the distributional state of the Iranian market rather than for large shifts in the direct effect of monetary policy.

The governance variable was statistically insignificant in both the Markov model and the broader HAC robustness regression. This result differs from cross-country evidence showing that stronger governance can reduce equity-market risk and improve market stability [28-30]. Iranian research has also associated governance conditions with stock-market competitiveness [31]. The difference is likely to reflect the frequency and construction of the governance measure rather than evidence that governance is economically irrelevant. The WGI variables are annual and change only gradually, whereas the dependent variable in this study is monthly return. Governance may therefore exert its primary influence through long-run market development, investor confidence, legal protection, and cost of capital rather than through short-run monthly return fluctuations.

The global financial crisis dummy was also insignificant. This finding should be interpreted cautiously because only a relatively small portion of the crisis period was represented in the effective nonlinear sample. Evidence from emerging Middle Eastern and North African markets shows that the 2008–2009 crisis altered financial linkages, return correlations, and volatility transmission [32]. The absence of a statistically significant direct crisis coefficient in the present study therefore does not imply that the Tehran Stock Exchange was insulated from global conditions. It indicates only that, after controlling for the variables included in the models and given the limited number of crisis observations, a distinct average return effect could not be statistically separated from zero.

Overall, the findings support a multilayered interpretation of stock-market dynamics in Iran. Monetary policy matters conceptually and may operate through several indirect macro-financial channels, but the directly identified structural monetary shock explains only a small share of monthly return fluctuations. Behavioral risk pricing appears more informative than the sentiment proxy used here, while sentiment becomes somewhat more relevant during volatile periods. The market itself exhibits pronounced and persistent nonlinear states, yet these states are defined more clearly by differences in average return and volatility than by large changes in the coefficients linking monetary and behavioral variables to returns. This pattern is consistent with a growing literature that emphasizes time variation, behavioral heterogeneity, and nonlinear transmission rather than a single stable monetary-policy effect on stock markets [16-18, 24].

The study has several limitations. First, the monetary-base series after February 2017 was not based entirely on a homogeneous set of directly observed monthly values and required reconstruction between published benchmark observations, which may have smoothed part of the true high-frequency monetary innovation. Second, SENT and RA are market-based proxies rather than direct measurements of investor psychology or individual risk

preferences, and their empirical interpretation is therefore necessarily indirect. Third, the recursive SVAR identification depends on the imposed contemporaneous ordering and cannot establish that alternative theoretically plausible identification schemes would yield identical responses. Fourth, the free-market exchange-rate series did not cover the full baseline sample, requiring the exchange-rate analysis to be conducted over a shorter period. Fifth, the annual nature of the governance indicators limits their ability to explain short-term monthly return fluctuations. Finally, the global financial crisis dummy contained relatively few effective crisis observations, particularly in the Markov Switching sample, reducing statistical power for identifying its independent effect.

Future research could extend the analysis by using alternative monetary-policy indicators, including interbank rates, policy-rate proxies, liquidity facilities, central-bank balance-sheet components, or high-frequency monetary surprises where reliable data become available. Alternative structural identification methods, such as sign restrictions, external instruments, or time-varying parameter SVARs, could be compared with recursive Cholesky identification. Investor sentiment could be measured using richer data, including trading flows, search intensity, news sentiment, social-media indicators, investor surveys, or separate individual and institutional transaction data. Risk aversion could likewise be estimated using option-implied measures, consumption-based indicators, or dynamic latent-factor models. Future studies could also investigate sector-level or firm-level heterogeneity, nonlinear exchange-rate channels, asymmetric positive and negative monetary shocks, more than two latent market regimes, and models that allow transition probabilities themselves to depend on macroeconomic or behavioral variables.

From a practical perspective, policymakers should avoid assuming that a change in the monetary base translates mechanically into an immediate stock-market reaction, because the results indicate that only the unexpected structural component is relevant for shock analysis and even this component had a limited direct effect on aggregate returns. Monetary authorities may therefore benefit from monitoring exchange-rate conditions, investor risk appetite, and broader financial-market expectations when evaluating the transmission of policy decisions. Investors and portfolio managers should recognize that market conditions differ substantially across persistent volatility regimes and that risk-pricing indicators may contain more useful information for return dynamics than a simple interpretation of monetary expansion or contraction. Regulators and market institutions can also improve decision-making by developing higher-frequency, transparent indicators of investor sentiment, trading participation, institutional quality, and monetary conditions so that financial-market responses can be assessed with greater precision.

Authors' Contributions

Authors equally contributed to this article.

Ethical Considerations

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

Acknowledgments

Authors thank all participants who participate in this study.

Conflict of Interest

The authors report no conflict of interest.

Funding/Financial Support

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

References

- [1] P. Maio, "Another Look at the Stock Return Response to Monetary Policy Actions," *Review of Finance*, vol. 18, no. 1, pp. 321-371, 2014, doi: 10.1093/rof/rfs050.
- [2] N. Yoshino, F. Taghizadeh-Hesary, A. Hassanzadeh, and A. D. Prasetyo, "Response of Stock Markets to Monetary Policy: The Tehran Stock Market Perspective," *Journal of Comparative Asian Development*, vol. 13, no. 3, pp. 517-545, 2014, doi: 10.1080/15339114.2014.985458.
- [3] M. Salehi, M. Behname, and M. S. Adibian, "Structural Shocks in Monetary Policy, Exchange Rates, and Stock Prices Using SVAR in Iran," *International Journal of Islamic and Middle Eastern Finance and Management*, vol. 14, no. 5, pp. 908-927, 2021, doi: 10.1108/IMEFM-04-2018-0150.
- [4] G. Sekandary and M. Bask, "Monetary Policy Uncertainty, Monetary Policy Surprises and Stock Returns," *Journal of Economics and Business*, vol. 124, p. 106106, 2023, doi: 10.1016/j.jeconbus.2022.106106.
- [5] M. Naderi, A. Hadizadeh, and A. Mirzapour Babajan, "Investigating the Effects of Monetary Policy Shocks on Stock Price Bubbles: The Application of Structural Vector Autoregression with Time-Varying Parameters (SVAR-TVP)," *Quarterly Journal of Economic Research (Sustainable Growth and Development)*, vol. 23, no. 2, pp. 193-218, 2023.
- [6] Z. Zarei, M. Hemmati, and P. Davoudi, "Evaluating the Effects of the Monetary Policy on the Total Stock Market Index in the Iranian Economy: Using the TVP-VAR and GARCH Approaches," *Iranian Economic Review*, vol. 28, no. 3, pp. 899-920, 2024.
- [7] K. Jahangiri and S. A. Hoseini Ebrahimabad, "The Study of Monetary Policy, Exchange Rate and Gold Effects on the Stock Market in Iran Using MS-VAR-EGARCH Model," *Financial Research Journal*, vol. 19, no. 3, pp. 389-414, 2017.
- [8] R. Zare, "Monetary Policy and Stock Market Cycles in Iran," *Quarterly Journal of Quantitative Economics*, vol. 19, no. 1, 2022.
- [9] A. Hashemi Sanjani, "Review to the Asymmetric Effect of Monetary Policy on Boom and Bust Cycles in the Iranian Stock Market," *Journal of Money and Economy*, vol. 19, no. 2, pp. 229-248, 2024, doi: 10.61882/jme.19.2.229.
- [10] I. Chatziantoniou, G. Filis, and C. Floros, "Asset Prices Regime-Switching and the Role of Inflation Targeting Monetary Policy," *Global Finance Journal*, vol. 32, no. 1, p. 15, 2017, doi: 10.1016/j.gfj.2015.12.002.
- [11] W. Wang, C. Su, and D. Duxbury, "Investor Sentiment and Stock Returns: Global Evidence," *Journal of Empirical Finance*, vol. 63, no. 365, p. 391, 2021, doi: 10.1016/j.jempfin.2021.07.010.
- [12] Y. Kim and K. Y. Lee, "Impact of Investor Sentiment on Stock Returns," *Asia-Pacific Journal of Financial Studies*, vol. 51, no. 1, pp. 132-162, 2022, doi: 10.1111/ajfs.12362.
- [13] J. Lao, H. Nie, and Y. Jiang, "Revisiting the Investor Sentiment-Stock Returns Relationship: A Multi-Scale Perspective Using Wavelets," *Physica A: Statistical Mechanics and its Applications*, vol. 499, no. 420, p. 427, 2018, doi: 10.1016/j.physa.2018.02.043.
- [14] A. Soroushyar and S. Ali Ahmadi, "Examining the Role of Momentum and Investor Sentiment in Herding Behavior in the Tehran Stock Exchange," *Investment Knowledge*, vol. 5, no. 18, pp. 147-159, 2016.
- [15] G. Gol-Arzi and S. Piri, "Examining the Impact of Investors' Emotional Decisions on the Occurrence of Crisis in the Tehran Stock Exchange," *Financial Management Strategy*, vol. 10, no. 2, pp. 99-116, 2022.
- [16] O. Cepni and R. Gupta, "Time-Varying Impact of Monetary Policy Shocks on US Stock Returns: The Role of Investor Sentiment," *The North American Journal of Economics and Finance*, vol. 58, p. 101550, 2021, doi: 10.1016/j.najef.2021.101550.
- [17] O. Cepni, R. Gupta, and Q. Ji, "Sentiment Regimes and Reaction of Stock Markets to Conventional and Unconventional Monetary Policies: Evidence from OECD Countries," *Journal of Behavioral Finance*, vol. 24, no. 3, pp. 365-381, 2023, doi: 10.1080/15427560.2021.1983576.
- [18] X. Zeng, Y. Sun, Y. Xu, and H. Hu, "Monetary Policy Shocks, Investor Sentiment, and Stock Returns: Firm-Level Evidence from China," *Emerging Markets Finance and Trade*, vol. 61, no. 9, pp. 2859-2876, 2025, doi: 10.1080/1540496X.2025.2467819.
- [19] G. Bekaert, E. C. Engstrom, and N. R. Xu, "The Time Variation in Risk Appetite and Uncertainty," *Management Science*, vol. 68, no. 6, pp. 3975-4004, 2022, doi: 10.1287/mnsc.2021.4068.
- [20] S. J. Yoon, "Time-Varying Risk Aversion and Return Predictability," *International Review of Economics & Finance*, vol. 48, no. 1, p. 18, 2017, doi: 10.1016/j.iref.2017.02.006.
- [21] Q. Zhang, T. Choudhry, J. M. Kuo, and X. Liu, "Does Liquidity Drive Stock Market Returns? The Role of Investor Risk Aversion," *Review of Quantitative Finance and Accounting*, vol. 57, no. 929, p. 958, 2021, doi: 10.1007/s11156-021-00966-5.

- [22] Z. He, L. He, and F. Wen, "Risk Compensation and Market Returns: The Role of Investor Sentiment in the Stock Market," *Emerging Markets Finance and Trade*, vol. 55, no. 3, pp. 704-718, 2019, doi: 10.1080/1540496X.2018.1460724.
- [23] D. Duxbury and W. Wang, "Investor Sentiment and the Risk-Return Relation: A Two-in-One Approach," *European Financial Management*, vol. 30, pp. 496-543, 2024, doi: 10.1111/eufm.12427.
- [24] M. Dahmene, A. Boughrara, and S. Slim, "Nonlinearity in Stock Returns: Do Risk Aversion, Investor Sentiment and Monetary Policy Shocks Matter?," *International Review of Economics & Finance*, vol. 71, no. 676, p. 699, 2021, doi: 10.1016/j.iref.2020.10.002.
- [25] M. Abbaszadeh, H. Rostami Jaz, M. Desineh, and M. H. Ranjbar, "Examining the Nonlinear Relationship between Risk Aversion and Stock Returns in the Iranian Capital Market," *Audit Knowledge*, vol. 24, no. 95, pp. 502-526, 2024.
- [26] X. Liu, D. Margaritis, and P. Wang, "Stock Market Volatility and Equity Returns: Evidence from a Two-State Markov-Switching Model with Regressors," *Journal of Empirical Finance*, vol. 19, no. 4, pp. 483-496, 2012, doi: 10.1016/j.jempfin.2012.04.011.
- [27] J. C. Cuestas and B. Tang, "A Markov Switching SVAR Analysis on the Relationship between Exchange Rate Changes and Stock Returns in China," *International Journal of Emerging Markets*, vol. 16, no. 3, pp. 625-642, 2021, doi: 10.1108/IJOEM-06-2019-0463.
- [28] S. W. Low, L. T. Tee, and S. R. Kew, "Does the Quality of Governance Matter for Equity Market Risk? Evidence from Emerging and Developed Equity Markets," *Journal of Business Economics and Management*, vol. 16, no. 3, pp. 660-674, 2015, doi: 10.3846/16111699.2012.720595.
- [29] P. K. Narayan, S. S. Sharma, and K. S. Thuraisamy, "Can Governance Quality Predict Stock Market Returns? New Global Evidence," *Pacific-Basin Finance Journal*, vol. 35, no. 367, p. 380, 2015, doi: 10.1016/j.pacfin.2015.02.007.
- [30] D. Y. Aharon, A. S. Baig, and R. J. Delisle, "The Effects of Country Governance Quality on the Stability of Equity Markets," *International Review of Finance*, vol. 25, no. 3, p. e70033, 2025, doi: 10.1111/irfi.70033.
- [31] H. Mir, A. Rahimipour, and S. A. Jalaei, "The Impact of Good Governance on Competitiveness in the Tehran Stock Exchange," *Economic Progress Policy*, vol. 7, no. 2, pp. 177-199, 2019.
- [32] S. Neaime, "The Global Financial Crisis, Financial Linkages and Correlations in Returns and Volatilities in Emerging MENA Stock Markets," *Emerging Markets Review*, vol. 13, no. 3, pp. 268-282, 2012, doi: 10.1016/j.ememar.2012.01.006.