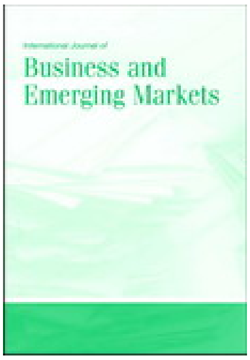

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Monetary policy's impact on Vietnamese stock market bubbles

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Abstract: This study investigates the relationship between monetary policy and stock market bubbles in Vietnam over the period 2010–2022. Using the supremum augmented Dickey-Fuller (SADF) and generalised supremum augmented Dickey-Fuller (GSADF) tests, the paper first confirms the presence of stock market bubbles in both the Ho Chi Minh City (VNI) and Hanoi (HNX) indices during several subperiods, notably 2017–2018 and 2020–2022. Subsequently, a vector autoregression (VAR) model is employed to examine the influence of key monetary policy instruments – overnight interbank interest rate, refinancing rate, money supply, inflation, and industrial production index – on bubble dynamics, represented by the P/E ratio. The results show that higher interbank and refinancing rates are associated with shrinking bubbles, while money supply, inflation, and industrial production exert relatively weak or insignificant impacts. The findings highlight the critical role of interest rate management in mitigating stock market bubbles and offer policy implications for regulators and investors in emerging markets like Vietnam.

Keywords: monetary policy; stock market bubbles; vector autoregression; Vietnam.

JEL codes: G10, G12, G14, G41, E44, E52.

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1 Introduction

Stock market bubbles have long been a recurring feature of global financial markets and are familiar to investors. Since its inception, the Vietnamese stock market has experienced pronounced ups and downs that have affected the real economy. Over time, it has affirmed its role in mobilising capital for development and has served as an effective channel for the government to implement macroeconomic policies. Episodes of bubbles have indeed occurred in Vietnam, and several studies have documented their presence.

A ‘bubble’ arises when the price of an asset increases irrationally without fundamental justification. Speculative buying fuels self-reinforcing gains: as prices climb, more participants anticipate further increases, enlarging the bubble until a sharp correction ensues. According to Garber (1990), a bubble is the portion of asset-price fluctuations unexplained by fundamentals – economic factors such as cash flows and interest rates. In the conventional view, a stock market bubble is the deviation of prices from intrinsic value. Yet this definition is impractical in real-time because bubbles are difficult to identify clearly until after they burst (Greenspan, 2002). Jackson and Schmidt (2021) characterise such episodes as exponential run-ups beyond intrinsic value followed by abrupt declines once prices reach a critical threshold.

Various approaches have been proposed to evaluate drivers of bubble formation. Notably, Le and Do (2016) show that common calendar effects significantly influence the Vietnamese market. While these effects do not create bubbles per se, the volatility they generate helps explain important market dynamics. Our review of studies in Vietnam and abroad identifies several contributors to bubbles, including economic growth, macroeconomic policies, investor psychology, and the breadth of market participation. Among these, macroeconomic policy – particularly monetary policy – plays a prominent role in shaping bubble formation.

This study improves upon prior Vietnamese evidence in two ways. First, relative to Do and Le (2024) – who primarily dated bubbles using SADF/GSADF without embedding a policy-transmission framework – and Nguyễn (2022) – who discussed 2020–2022 developments qualitatively – our analysis unifies formal bubble dating with monetary-policy transmission in a single empirical design. Methodologically, we combine SADF/GSADF tests (with Monte Carlo critical values) to date bubbles on VNI and HNX with a monthly VAR that models bubble dynamics – proxied by the P/E ratio – jointly with overnight interbank and refinancing rates, money supply (M2), inflation (CPI), and industrial production (IIP); we also examine two-way interactions via Granger causality and trace impulse responses and variance decomposition.

Second, we provide new quantitative evidence for the COVID-19 period (2020–2022). We identify pandemic-era bubble episodes on both exchanges and estimate how monetary instruments relate to bubble measures during this time. The results indicate that higher interbank and refinancing rates are associated with shrinking bubbles, whereas money supply, inflation, and industrial production exert relatively weak or statistically small effects. By integrating bubble identification and policy transmission, the study offers a clearer basis for policy evaluation in an emerging-market context.

2 Theoretical framework

Monetary policy comprises the instruments used by central banks to steer short-term funding conditions and the price of liquidity – through the policy rate and operational tools such as open-market operations and standing facilities – with the aim of stabilising inflation and output (Woodford, 2003; Bindseil, 2014). In Vietnam, the statutory objectives are specified in the 2010 Law on the State Bank of Vietnam.

Monetary policy is both a driver of, and a potential remedy for, stock market imbalances. In a discounted-cash-flow perspective, policy moves equity values via the discount factor and via expectations about future cash flows. Smirlock and Yawitz (1985) argue that changes in policy rates transmit to market rates, directly shifting discount rates and indirectly altering expected profitability. Galí and Gambetti (2015) suggest that tighter policy can compress the scale of bubbles, especially over longer horizons, by raising required returns and dampening speculative forces.

Two transmission channels are central to this study. First, the *interest-rate channel*: policy rates anchor short-term funding costs and, through expectations, the term structure, shaping discount rates and financing conditions for households and firms. Under expansionary policy and rising broad money, real rates decline and the cost of capital falls. Lower deposit returns reduce the incentive to save and encourage portfolio shifts toward higher-return assets such as equities; persistent inflows can elevate prices and, under favourable conditions, feed bubble dynamics. Second, the *credit (bank-lending) channel*: an accommodative stance – looser lending standards, higher bank reserves, expanding deposits – increases the supply of credit. Firms borrow more to finance investment; stronger activity and earnings expectations then feed back into stock prices. Together, these mechanisms align with the *risk-taking channel* of monetary policy, whereby easier conditions relax balance-sheet constraints and compress perceived risk premia, raising appetite for risk (Borio and Zhu, 2012; Bruno and Shin, 2015).

These financing channels interact with a *behavioural layer* that shapes the intensity and persistence of bubbles. When sentiment is elevated – especially for assets that are hard to arbitrage – prices deviate more from fundamentals (Baker and Wurgler, 2006). Investors also extrapolate recent returns; past run-ups raise expected returns and invite further demand, generating momentum that can power self-reinforcing price increases

(Greenwood and Shleifer, 2014). Limits to arbitrage – funding frictions, short-sale constraints, and career concerns – impede rapid correction by rational traders, allowing mispricing to persist even as financing conditions normalise. Thus, monetary policy sets the backdrop by easing or tightening financial conditions, while behavioural forces determine whether valuations merely adjust or escalate into speculative episodes.

Synthesising these elements, the framework proceeds in three steps. First, policy instruments – overnight interbank and refinancing rates, alongside liquidity tools that influence broad money (M2) – affect the interest-rate and credit channels, altering discount rates, liquidity, leverage, and access to external finance. Second, these financing conditions interact with the behavioural layer: easier conditions raise risk appetite and the salience of recent returns, strengthening extrapolative expectations and encouraging reallocation toward equities. Third, the interaction between financing and behaviour is reflected in *market outcomes*: valuation metrics (e.g., the price-to-earnings, P/E, ratio), indicators of explosive behaviour, and trading activity. Feedback may then arise: price run-ups bolster sentiment and risk-taking, while authorities adjust policy in response to overheating signals.

3 Literature review

Stock market bubbles have long attracted sustained attention from economists, and a wide range of empirical strategies has been proposed to verify their existence across countries and periods. Broadly, the literature follows two complementary paths:

1. testing whether prices are anchored by fundamentals
2. dating episodes of explosive behaviour directly; a related strand then examines how monetary policy and macro-financial conditions shape those episodes.

In the first path, researchers ask whether stock prices move with cash-flow fundamentals. The Johansen (1998) cointegration framework tests for a stable long-run relation between stock prices and dividends; when such a vector exists, price increases can, in principle, be justified by dividend growth. Applying this logic, Herrera and Perry (2001) studied Latin America over 1980–2001 and identified intervals of overpricing consistent with bubble behaviour when deviations from cointegration widened. The appeal of this approach is its close link to valuation theory, though results can be sensitive to dividend measurement, structural breaks, and regime shifts – issues that are particularly salient in emerging markets.

The second path dates bubbles from the time series itself. Pan (2020) applied the right-tailed sup augmented Dickey-Fuller (SADF) and the generalised supremum augmented Dickey-Fuller (GSADF) tests to 1978–2018 and detected two prominent US episodes (1985–October 1987 and 1995–2000). Phillips et al. (2012) also used SADF to search for explosive dynamics; however, SADF performs best when there is a single bubble. With multiple episodes – especially if a later one is shorter – SADF may miss onsets and becomes less reliable over long samples. To address these issues, Phillips et al. (2015) developed GSADF, which iterates SADF over rolling windows and improves both detection and date-stamping when several bubbles occur. This tradition has been widely adopted, including in emerging markets, because it yields explicit start/stop dates that facilitate subsequent policy analysis.

For Vietnam, numerous studies document the presence of stock market bubbles. Applying SADF/GSADF to 2006–2021, Lê and Nguyễn (2022) identify multiple episodes across the mid-2000s, mid-2010s, and the COVID-19 period. Complementing quantitative work, Nguyễn (2022) provided qualitative evidence that bubble conditions re-emerged during 2020–2022 in the wake of the COVID-19 shock. Together, these studies establish that exuberant phases do occur in Vietnam, though the role of specific policy instruments and their interaction with investor behaviour remain less explored within a unified design.

A related literature connects monetary policy to equity valuations and bubbles. At high frequency, Bernanke and Kuttner (2005) used an event-study design around policy decision days and showed that an unexpected 25-basis-point cut in the policy rate is associated with roughly a 1% rise in stock prices, indicating a powerful discount-rate channel. At lower frequencies, Caraiani and Călin (2020) employed a Bayesian VAR for OECD economies (1990–2017), modelling a price-index proxy for bubbles alongside CPI, policy rates, and dividends; they found that more financially developed markets exhibit stronger negative effects of policy tightening on bubbles. For Vietnam specifically, Nguyen (2013) analyzed transmission via the interest-rate, exchange-rate, asset-price, and credit channels using an LA-VAR, concluding that the stock-price channel was relatively ineffective – consistent with an emerging market with limited depth and frictions in arbitrage.

Building on the monetary-transmission evidence above, recent emerging-market studies highlight macro-financial channels shaping equity dynamics. Banerji and Shettima (2025) show oil – FX volatility spillovers in India. Dada et al. (2024) document nonlinear, asymmetric currency-equity effects in Nigeria. Kumar (2025) finds cross-market cointegration during COVID-19, indicating elevated comovement under uncertainty. Taken together, these findings imply that macro shocks and cross-market linkages can condition speculative pressure in settings comparable to Vietnam.

In parallel, adjacent literature points to institutional and ESG-related channels that shape valuations and expectations. In an exchange context, Abdelhaq et al. (2025) show that stronger governance aligns with higher intellectual-capital efficiency on the Palestine Stock Exchange. At the firm level, Tarigan et al. (2025) find ESG disclosure raises firm value, with innovation strengthening this effect, while macro-ESG backdrops differ markedly across China and India (Rao and Gaur, 2025). On the demand and innovation sides, young consumers' willingness to pay for green products (Pathak and Malakar, 2024) and evolving definitions of technology in university tech transfer (Townes, 2025) can fuel valuation themes. Finally, Taques et al. (2025) link patents/R&D to market value. Although indirect to monetary transmission, these studies offer expectation channels that may interact with bubble dynamics in emerging markets.

4 Methodology

4.1 Identifying stock market bubbles (SADF/GSADF)

We employ the right-tailed bubble tests of Phillips et al. (2012) – the supremum augmented Dickey-Fuller (SADF) test – and of Phillips et al. (2015) – the generalised SADF (GSADF) – on the VNI and HNX indices. Rejection of the null in these tests is taken as empirical evidence of explosive behaviour consistent with asset-price bubbles.

Critical values are obtained by Monte Carlo simulation, and the procedures also deliver date-stamping of bubble start and end points.

Both tests are based on the following reduced-form regression:

$$\Delta y_t = \mu + \delta y_{t-1} + \sum\nolimits_{i=1}^{p} \phi \Delta y_{t-1} + e_t \tag{1}$$

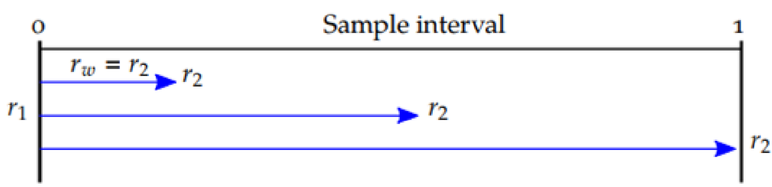
where y_t is the stock price at time t; μ is an intercept; p is the maximum lag length; ϕ_i (for $i = 1, \ldots, p$) are the autoregressive coefficients; and e_t is the error term. Testing for a bubble (explosive behaviour) uses a right-tail variant of the standard ADF unit-root test, with null and alternative hypotheses:

$$H_0 : \delta = 1, H_1 : \delta > 1$$

Let r_1 and r_2 denote the fractional start and end points of the estimation window, so the window size is $r_w = r_2 - r_1$. Let r_0 be the minimum initial window (chosen by the user). In the SADF test, the start point is fixed at the first observation, $r_1 = 0$, while the end point expands recursively $r_2 \in [r_0, 1]$ one observation at a time. Each regression yields an ADF statistic ADF_{r_2}. The SADF statistic is the supremum of ADF_{r_2} over $r_2 \in [r_0, 1]$:

$$SADF(r_0) = \sup\{ADF_{r_2}\}, \quad r_2 \in [r_0, 1] \tag{2}$$

Figure 1 Illustration of the SADF procedure (see online version for colours)



The GSADF generalises SADF by allowing both the start and end points to vary: $r_1 \in [0, r_2 - r_0]$ and $r_2 \in [r_0, 1]$. The GSADF statistic is then defined as the supremum over this two-dimensional recursion:

$$GSADF(r_0) = \sup\{ADF_{r_1}^{r_2}\}, \quad r_2 \in [r_0, 1] r_1 \in [0, r_2 - r_0] \tag{3}$$

Phillips et al. (2012) also set out a dating algorithm. Compare each element of the sequence ADF_{r_2} with the corresponding right-tail critical value to determine the bubble's origination time $T(r_e)$ (the first r_2 at which ADF_{r_2} crosses the critical value from below) and the termination time $T(r_f)$ [the first $r_2 > T(r_e)$ at which the statistic falls back below the critical value]. Formally:

$$r_e = \inf\{r_2 : ADF_{r_2} > cv_{r_2}^{\beta_T}\}, \quad r_2 \in [r_0, 1] \tag{4}$$

$$r_f = \inf\{r_2 : ADF_{r_2} > cv_{r_2}^{\beta_T}\}, \quad r_2 \in [r_e, 1] \tag{5}$$

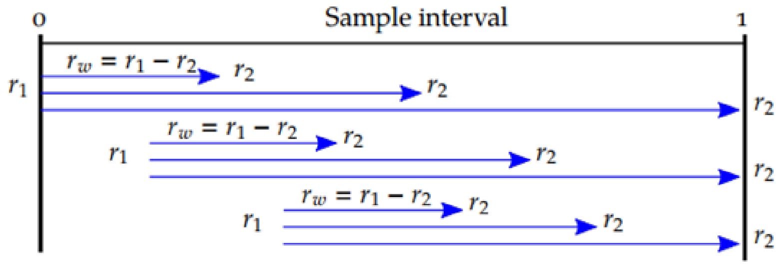
where $cv_{r_2}^{\beta_T}$ is the 100 (1 – β_T)% critical value of the ADF statistic based on $T(r_2)$ observations. For GSADF-based dating, the same logic applies using the backward SADF statistic $BSADF(r_0)$ for $r_2 \in [r_0, 1]$, which relates to GSADF as:

$$r_e = \inf\left\{ r_2 : BSADF_{r_2}(r_0) > cv_{r_2}^{\beta_{T r_2}} \right\}, \quad r_2 \in [r_0, 1] \tag{6}$$

$$r_f = \inf\left\{ r_2 : BSADF_{r_2}(r_0) > cv_{r_2}^{\beta_{T r_2}} \right\}, \quad r_2 \in [r_e, 1] \tag{7}$$

$$GSADF(r_0) = \sup\left\{ BSADF_{r_2}(r_0) \right\}, \quad r_2 \in [r_0, 1] \tag{8}$$

Figure 2 Illustration of the GSADF procedure (see online version for colours)



To examine the existence of stock-market bubbles in Vietnam, we use daily closing prices of the VNI (Ho Chi Minh City Stock Exchange) and HNX (Hanoi Stock Exchange) from 01/01/2010 to 31/12/2022. The sample is split into three subperiods – 2010–2013, 2014–2019, and 2020–2022 – to

1 isolate episodes of exuberance suggested by theory

2 test for suspected bubbles in 2021–2022, a period for which formal quantitative evidence has been limited.

4.2 Modelling monetary policy effects (VAR specification)

Building on Galí and Gambetti (2015), we estimate a vector autoregression (VAR) to assess how monetary policy relates to stock-market exuberance in Vietnam. The VAR includes the P/E ratio (exuberance proxy) and the policy/macro variables refinancing rate (IR), money supply (M2), consumer price index (CPI), industrial production index (IIP), and the overnight interbank rate (R). A VAR is appropriate because it handles monthly time-series interactions, is standard for policy analysis/forecasting, and allows Granger-causal links among variables.

The VAR is written as a system of six equations – one equation for each endogenous variable $y_t \in \{PE_t, IR_t, M2_t, CPI_t, IIP_t, R_t\}$ – with a constant and p lags of all variables on the right-hand side. Coefficients are indexed α_{ji}, β_{ji}, γ_{ji}, δ_{ji}, λ_{ji}, ϕ_{ji} for equation $j = 1, \ldots, 6$ and lag $i = 1, \ldots, p$; the error terms ε_{jt} are white noise and may be contemporaneously correlated (in matrix terms, rows correspond to the six equations – PE, IR, $M2$, CPI, IIP, R – while columns correspond to the stacked regressors PE_{t-i}, IR_{t-i}, $M2_{t-i}$, CPI_{t-i}, IIP_{t-i}, R_{t-i} for $i = 1, \ldots, p$; we use the subscript $t - i$ to denote the i^{th} lag).

$$\begin{aligned} PE_t &= c_{1t} \\ &+ \sum_{i=1}^{p} \left(\alpha_{1i} PE_{t-i} + \beta_{1i} IR_{t-i} + \gamma_{1i} M2_{t-i} + \delta_{1i} CPI_{t-i} + \lambda_{1i} IIP_{t-i} + \phi_{1i} R_{t-i} \right) + \varepsilon_{1t} \end{aligned} \tag{9}$$

$$IR_t = c_{2t} + \sum_{i=1}^{p} \left(\alpha_{2i} PE_{t-i} + \beta_{2i} IR_{t-i} + \gamma_{2i} M2_{t-i} + \delta_{2i} CPI_{t-i} + \lambda_{2i} IIP_{t-i} + \phi_{2i} R_{t-i} \right) + \varepsilon_{2t} \tag{10}$$

$$M2_t = c_{3t} + \sum_{i=1}^{p} \left(\alpha_{3i} PE_{t-i} + \beta_{3i} IR_{t-i} + \gamma_{3i} M2_{t-i} + \delta_{3i} CPI_{t-i} + \lambda_{3i} IIP_{t-i} + \phi_{3i} R_{t-i} \right) + \varepsilon_{3t} \tag{11}$$

$$CPI_t = c_{4t} + \sum_{i=1}^{p} \left(\alpha_{4i} PE_{t-i} + \beta_{4i} IR_{t-i} + \gamma_{4i} M2_{t-i} + \delta_{4i} CPI_{t-i} + \lambda_{4i} IIP_{t-i} + \phi_{4i} R_{t-i} \right) + \varepsilon_{4t} \tag{12}$$

$$IIP_t = c_{5t} + \sum_{i=1}^{p} \left(\alpha_{5i} PE_{t-i} + \beta_{5i} IR_{t-i} + \gamma_{5i} M2_{t-i} + \delta_{5i} CPI_{t-i} + \lambda_{5i} IIP_{t-i} + \phi_{5i} R_{t-i} \right) + \varepsilon_{5t} \tag{13}$$

$$R_t = c_{6t} + \sum_{i=1}^{p} \left(\alpha_{6i} PE_{t-i} + \beta_{6i} IR_{t-i} + \gamma_{6i} M2_{t-i} + \delta_{6i} CPI_{t-i} + \lambda_{6i} IIP_{t-i} + \phi_{6i} R_{t-i} \right) + \varepsilon_{6t} \tag{14}$$

The variables used in the model are defined as follows. The price-to-earnings (P/E) ratio for the Vietnamese stock market measures how much investors pay per unit of earnings. Because it tracks valuations relative to fundamentals, sustained run-ups followed by abrupt reversals are taken as indicative of bubble phases; we therefore use market-level P/E as our proxy for exuberance. The overnight interbank rate (R) is the cost of very short-term bank funding. Central-bank decisions pass through quickly to money-market rates and then along the curve (Isakova, 2008), altering financing conditions and near-term investor sentiment. In an accommodative stance, lower R tends to reallocate funds towards equities; tighter conditions temper equity demand. The refinancing rate (IR) – the SBV policy lending rate – regulates system-wide liquidity and the price of central-bank credit. Changes in IR affect discount rates and leverage in the banking system and, by extension, equity valuations. Money supply (M2) comprises currency, demand deposits, cheques, savings deposits and short-term time deposits, capturing broad liquidity available to households and firms. Where interest-rate tools face institutional constraints, money-supply management can be pivotal (Li and Liu, 2017; Berger et al., 2009); expansions in M2 generally coincide with easier conditions and stronger risk appetite. The consumer price index (CPI) is the standard indicator of inflation. Inflation influences equities through both cash-flow and discount-rate channels and is central to policy objectives; in Vietnam's context, near-term nominal and liquidity effects can raise the market P/E. The index of industrial production (IIP) tracks industrial output relative to a base period and signals real activity. Higher IIP is associated with stronger expected cash flows and higher equity values (Geske and Roll, 1983), whereas weak production points to softer revenues and profits.

We use monthly data for January 2010–December 2022. The P/E series is from Bloomberg; the remaining series are from the IMF (IFS). Variable names, definitions, units, expected signs, and sources are summarised in Appendix Table A1. Transformations (first differences for PE, R, IR, M2; levels for CPI and IIP) follow DF tests and are described in Subsection 5.2.1.

5 Results

5.1 Bubble detection results

Appendix Table A2 reports descriptive statistics for the daily closing levels of the VN-Index (VNI) and HNX-Index over 2010–2022. For each calendar year it lists the mean, range (min–max), standard deviation, and number of observations, providing a compact view of dispersion and sample coverage that frames the subsequent bubble-dating results.

We implemented Monte Carlo simulations with 1,000 iterations to obtain the finite-sample critical values for the SADF (Phillips et al., 2012) and GSADF (Phillips et al., 2015) tests. Stata 17 was used to select the initial window size and the optimal lag length automatically; the optimal lag was 1.

Table 1 SADF and GSADF test results for the VN-Index and HNX-Index, 2010–2022

2010–2013	*t-statistic VNI*	*Critical values*		*t-statistic HNX*	*Critical values*	
		99%	*95%*		*99%*	*95%*
SADF	–0.4951	2.0959	1.5696	0.5583	2.0959	–0.0458
GSADF	3.7203	2.7647	2.3050	2.2764	2.7647	2.3050
2014–2019	*t-statistic VNI*	*Critical values*		*t-statistic HNX*	*Critical values*	
		99%	*95%*		*99%*	*95%*
SADF	3.6408	2.1143	1.6190	1.5351	2.1139	1.6186
GSADF	4.6751	2.8436	2.3812	2.6523	2.3807	2.5434
2020–2022	*t-statistic VNI*	*Critical values*		*t-statistic HNX*	*Critical values*	
		99%	*95%*		*99%*	*95%*
SADF	1.0595	2.0836	1.5549	5.3661	2.0838	1.5552
GSADF	3.2859	2.6879	2.2718	5.5175	2.6886	2.2722

To determine the presence of stock-market bubbles, we applied the SADF and GSADF tests to both indices and flagged periods when the test statistics exceeded the 95% and 99% critical values (i.e., the 5% and 1% significance levels). The results show no evidence of bubbles for either index during 2010–2013. From 2014–2019, for the VN-Index both tests (SADF and GSADF) indicate a bubble at the 1% and 5% levels, whereas for the HNX-Index the SADF does not, but the GSADF – widely regarded as more precise for multiple episodes – does at the same levels. From 2020–2022, the SADF does not detect a bubble for the VN-Index, while the GSADF indicates one; the HNX-Index mirrors this pattern. We therefore conclude that bubbles occurred during 2017–2018 and 2020–2022, in line with existing claims. To pinpoint the timing more precisely, we conduct year-by-year tests within 2010–2022. The summary appears in Table 2.

Table 2 indicates bubbles in 2017, 2018, and 2020. In 2021–2022 the indices diverge: the VN-Index exhibits a bubble in 2022, while the HNX-Index shows one in 2021. Specifically, the detected windows are: for the HNX-Index, 06/2017–04/2018 and 11/2020–12/2021; for the VN-Index, 10/2017–03/2018, 02/2021–07/2021, and throughout 2022.

Table 2 Annual SADF and GSADF results for the VN-Index and HNX-Index, 2010–2022

VNI	*SADF*	*Critical values*		*GSADF*	*Critical values*	
		99%	*95%*		*99%*	*95%*
2010	–0.4951	1.9562	1.3928	3.0176	2.5564	2.1135
2011	–0.0997	1.9562	1.3928	1.2178	2.5564	2.1135
2012	0.3819	1.9562	1.3928	1.6553	2.5564	2.1135
2013	–2.2980	1.9562	1.3928	0.2540	2.5564	2.1135
2014	–0.9022	1.9562	1.3928	1.6195	2.5564	2.1135
2015	–0.4159	1.9562	1.3928	1.8582	2.5564	2.1135
2016	0.7488	1.9502	1.3928	0.9645	2.5438	2.0897
2017	3.4309	1.9562	1.3928	4.4587	2.5564	2.1135
2018	2.3909	1.9562	1.3928	2.9317	2.5564	2.1135
2019	1.3920	1.9562	1.3928	1.4448	2.5564	2.1135
2020	2.5115	1.9502	1.3928	2.5440	2.5438	2.0897
2021	–0.3808	1.9562	1.3928	1.6166	2.5564	2.1135
2022	1.6414	1.9562	1.3928	2.7803	2.5564	2.1135
HNX	*SADF*	*Critical values*		*GSADF*	*Critical values*	
		99%	*95%*		*99%*	*95%*
2010	0.5583	1.9562	1.3928	3.0805	2.5564	2.1135
2011	0.3565	1.9562	1.3928	2.5240	2.5564	2.1135
2012	1.7632	1.9562	1.3928	1.7632	2.5564	2.1135
2013	–1.7309	1.9562	1.3928	1.5636	2.5564	2.1135
2014	0.6952	1.9562	1.3928	1.6728	2.5564	2.1135
2015	0.2970	1.9562	1.3928	2.0213	2.5564	2.1135
2016	0.5484	1.9502	1.3928	1.3486	2.5438	2.0897
2017	2.0343	1.9562	1.3928	2.6792	2.5564	2.1135
2018	2.7893	1.9562	1.3928	2.7655	2.5564	2.1135
2019	–0.0761	1.9562	1.3928	0.8562	2.5564	2.1135
2020	3.8210	1.9502	1.3928	4.8504	2.5438	2.0897
2021	2.4148	1.9562	1.3928	3.4345	2.1135	2.5564
2022	0.4147	1.9502	1.3928	3.4463	2.5438	2.0897

5.2 VAR results

Appendix Table A3 summarises the monthly series used in the VAR – PE, R (refinancing rate), IR (interbank overnight rate), M2, CPI, and IIP – reporting observations, means, standard deviations, and extrema for 2010–2022. These statistics clarify scale differences across variables and complement the stationarity checks and transformations described below.

5.2.1 Stationarity test

The Dickey-Fuller (DF) test indicates non-stationarity for all variables except CPI and IIP. We therefore apply first-order differencing to PE, R, IR, and M2, yielding stationary series DPE, DR, DIR, and DM2; CPI and IIP remain in levels.

Table 3 Stationarity test results

Variable	*Original results*		*Results after first-order difference*	
	DF	*P-value*	*DF*	*P-value*
PE	–2.571	0.0992	–12.639	0.0000
R	–1.792	0.3844	–7.689	0.0000
IR	–0.826	0.8115	–7.183	0.0000
M2	2.339	0.9990	–13.076	0.0000
CPI	–4.160	0.0008	–7.016	0.0000
IIP	–16.170	0.0000	–23.573	0.0000

Lag length is chosen using VAR lag-order selection criteria (LR, FPE, AIC, SBIC, HQIC). Among these, FPE, HQIC, and SBIC jointly favour a one-lag specification; accordingly, we set $p = 1$ for VAR estimation and the Granger causality analysis.

Table 4 Optimal lag length for the VAR model

Sample: 2010m6–2022m12			*df*	*p*	*PPE*	*AIC*	*Number of obs. 151*	
Lag	*LL*	*LR*					*HQIC*	*SBIC*
0	–3,674.61	NA	NA	NA	6.0e+13	48.7498	48.7985	48.8697
1	–3,100.06	1149.100	36	0.0000	4.8e+10*	41.6166	41.9576*	42.4559*
2	–3,064.02	72.068	36	0.0000	4.8e+10	41.6162*	42.2494	43.1748
3	–3,037.03	53.973*	36	0.0028	5.4e+10	41.7356	42.6610	44.0135
4	–3,015.28	43.513	36	0.1820	6.6e+10	41.9242	43.1419	44.9215

5.2.2 Granger causality test

Results show bidirectional Granger causality between DPE and DR, and between DPE and DIR. With *p-value* < 0.05, we reject H_0 of no causality in both directions for these pairs. For one-way links, DPE Granger-causes DM2, while DM2 does not Granger-cause DPE. DPE does not Granger-cause CPI or IIP; in the reverse direction, both IIP and CPI Granger-cause DPE.

5.2.3 VAR estimation

Model diagnostics – stability, residual normality, and autocorrelation – support the VAR's adequacy at the 5% level. Estimation results (5% level) show that all variables except CPI significantly affect the bubble proxy. Specifically, DPE, DR, DIR, and IIP have negative effects, whereas DM2 and CPI have positive effects.

Table 5 Granger causality test results

Hypothesis H_0	*χ^2 (chi²)*	*P-value (prob. > chi²)*
DPE has no effect on DR	4.167	0.041
DR has no effect on DPE	2.474	0.016
DPE has no effect on DM2	4.790	0.029
DM2 has no effect on DPE	0.724	0.395
DPE has no effect on DIR	4.928	0.026
DIR has no effect on DME	5.287	0.021
DPE has no effect on CPI	0.267	0.605
CPI has no effect on DPE	3.758	0.043
DPE has no effect on IIP	0.350	0.554
IIP has no effect on DPE	4.134	0.042

Table 6 VAR estimation results

Vector autoregression						
Sample: 2010m3 thru 2022m12				Number of obs. = 154		
Log likelihood = –3157.616				AIC = 41.55346		
FPE = 4.48e+10				HQIC = 41.8899		
Det (Sigma_ml) = 2.60e+10				SBIC = 42.38172		
Equation	*Parms*	*RMSE*	*R-sq*	*Chi²*	*P > chi²*	
dpe	7	1.0012	0.1108	19.1901	0.0039	
dr	7	1.0057	0.2853	61.4877	0.0000	
dm2	7	97,844.5	0.0692	11.4523	0.0754	
dir	7	0.4311	0.2854	61.5180	0.0000	
cpi	7	0.7187	0.9989	134,406.6	0.0000	
iip	7	6.6728	0.1851	34.9708	0.0000	
	Coefficient	*Std. err.*	*z*	*P > |z|*	*[95% conf. interval]*	
			dpe			
dpe L1	–0.189469	0.085412	–2.22	0.027	–0.356873	–0.022066
dr L1	–0.153324	0.075115	–2.04	0.041	–0.300546	–0.006102
dm2 L1	0.000002	0.000001	2.19	0.029	0.000000	0.000003
dir L1	–0.366302	0.165004	–2.22	0.026	–0.689704	–0.042901
cpi L1	0.002034	0.003934	0.52	0.605	–0.005675	0.009744
iip L1	–0.022879	0.011252	–2.03	0.042	–0.044932	–0.000825
_cons	–0.352688	0.579595	–0.61	0.543	–1.488673	0.783297

Figure 3 Impulse response of the bubble to its own shock (see online version for colours)

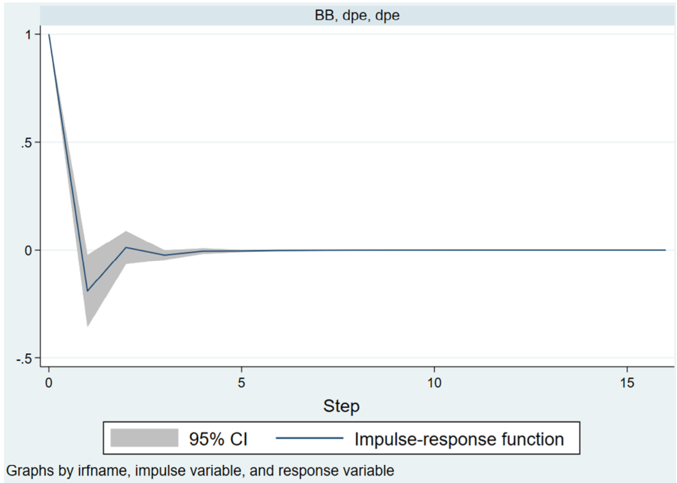
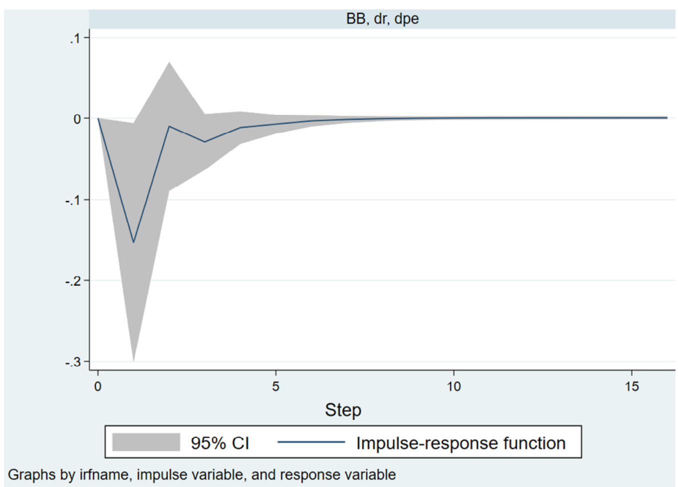


Figure 4 Impulse response of the bubble to an interbank overnight rate shock (see online version for colours)



5.2.4 Impulse response functions

We examine how the stock-market bubble (proxied by the P/E index) responds to shocks in key variables over a 16-month horizon. The bubble responds immediately to its own shock: early price surges attract investors and amplify the bubble initially, with the effect fading over time. In contrast, shocks to both interest rates produce a sharp initial decline in the bubble, followed by a gradual further decline. The bubble moves with M2 – expanding when money supply increases – before adjusting downward. Shocks to CPI and IIP have small direct effects, though they can signal conditions conducive to bubble formation given their sensitivity to monetary policy.

Figure 5 Impulse response of the bubble to a refinancing rate shock (see online version for colours)

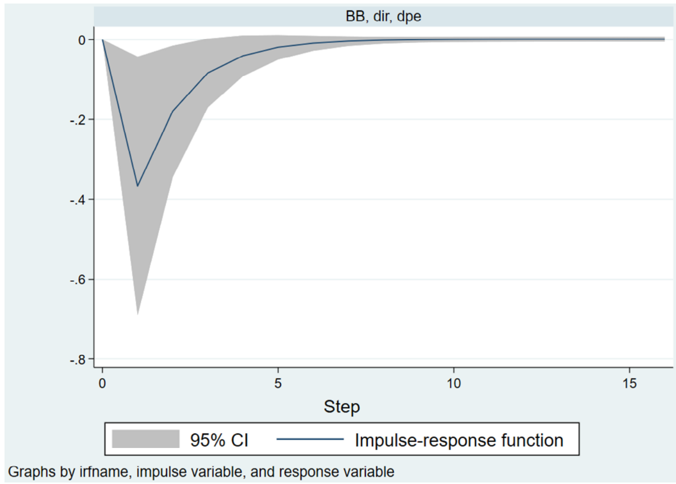
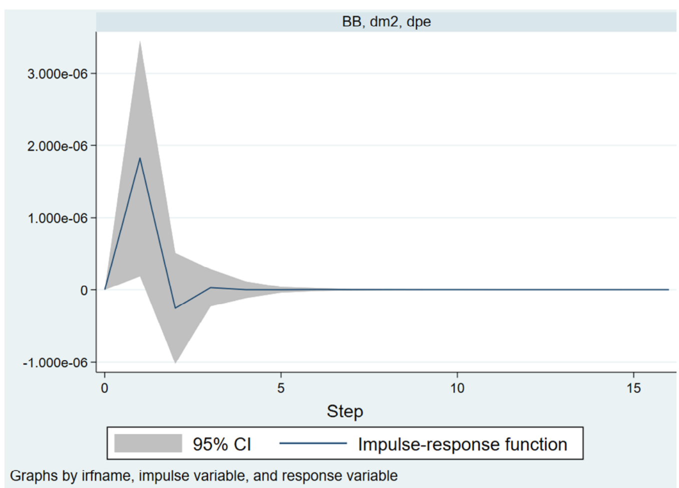


Figure 6 Impulse response of the bubble to a money supply (M2) shock (see online version for colours)



5.2.5 Variance decomposition

Using recursive (Cholesky) decomposition following Sims (1987), we quantify each factor's contribution to bubble fluctuations. The bubble's own shocks dominate, explaining about 89–90% of the variance. Other contributors are modest: the interbank overnight rate accounts for 1.8–1.9%, M2 for about 2.6%, the refinancing rate for 2.6–3.3%, and IIP for 1.9–2%, while CPI contributes less than 1%.

Figure 7 Impulse response of the bubble to an inflation (CPI) shock (see online version for colours)

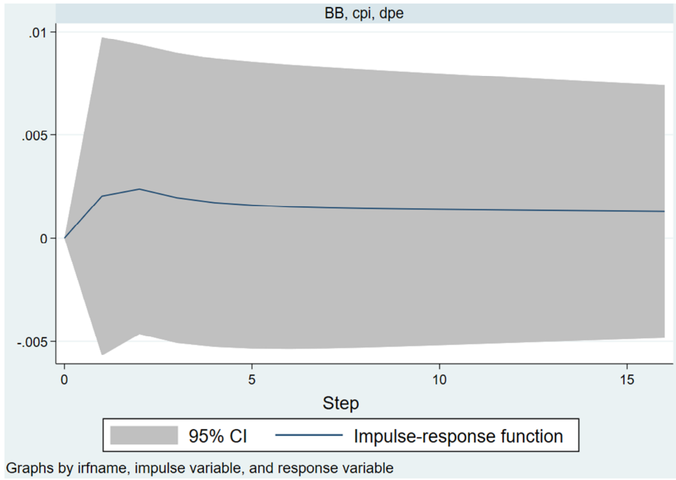
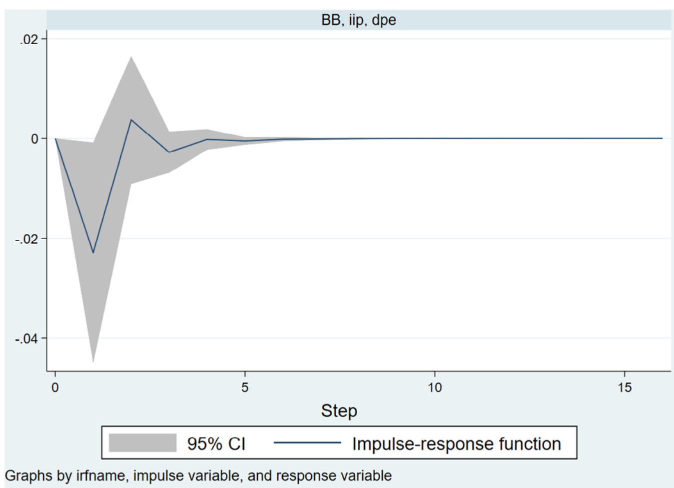


Figure 8 Impulse response of the P/E index to an IIP shock (see online version for colours)



6 Discussion

First, the SADF and GSADF tests indicate stock-market bubbles on the HNX during 06/2017–04/2018 and 11/2020–12/2021, and on the VNI during 10/2017–03/2018, 02/2021–07/2021, and throughout 2022. Notably, these windows often coincide with major macro-financial episodes. The 2000–2003 interval marked the market's inception and fragility; 2007–2008 saw a proliferation of new financial products alongside risks from the global financial crisis; 2017–2018 combined an IPO wave with strong growth and heightened international political uncertainty; and 2020–2022 coincided with the COVID-19 pandemic. Within 2020–2022, both tests point to the emergence of bubbles in

HNX and VNI in 2020, followed by an HNX bubble in 2021 and a VNI bubble in 2022. Several factors plausibly underpin these outcomes:

1 Vietnam's effective pandemic control limited economic damage and stabilised the recovery

2 supportive policies, including accommodative monetary conditions, faster credit growth, and lower interest rates

3 a shift by investors toward equities as a lucrative channel, reflected in record new account openings in 2021 and high liquidity.

Consistent with this narrative, during bubble phases stock prices rose much faster than corporate profits. As a result, the P/E ratio fluctuated and increased alongside the broad market index, reinforcing its use by practitioners and investors as a practical indicator of emerging bubbles.

Table 7 Variance decomposition results

Step	*DPE*	*DR*	*DM2*	*DIR*	*CPI*	*IIP*
0	0	0	0	0	0	0
1	1	0	0	0	0	0
2	0.90799	0.018461	0.026335	0.026742	0.000917	0.019556
3	0.902397	0.018585	0.026663	0.031493	0.000921	0.019941
4	0.900307	0.019311	0.026603	0.032668	0.000941	0.020171
5	0.899948	0.019432	0.026592	0.032922	0.000942	0.020161
6	0.899838	0.019477	0.026589	0.032981	0.000945	0.020171
7	0.899816	0.019486	0.026588	0.032993	0.000946	0.020171
8	0.899811	0.019488	0.026588	0.032995	0.000948	0.020171
9	0.89981	0.019488	0.026588	0.032995	0.000949	0.020171
10	0.899809	0.019488	0.026588	0.032995	0.00095	0.020171
11	0.899808	0.019488	0.026588	0.032995	0.00095	0.020171
12	0.899807	0.019488	0.026588	0.032995	0.000951	0.020171
13	0.899805	0.019488	0.026588	0.032995	0.000952	0.020171
14	0.899804	0.019488	0.026588	0.032996	0.000953	0.020171
15	0.899803	0.019488	0.026588	0.032996	0.000954	0.020171
16	0.899082	0.019489	0.026588	0.032997	0.000954	0.020171

Second, the VAR indicates that the interbank overnight rate (IR) and the refinancing rate (R) are inversely related to stock-market bubbles – when rates rise, bubbles tend to contract. From 2012 to 2014, cooling inflation allowed the State Bank of Vietnam to loosen policy: the refinancing rate fell from 8.5% to 6.5% per annum and the interbank overnight rate dropped from its 12.8% peak to around 1%–4%, translating into lower commercial lending rates. The market rebounded over this period, with the VNI up 41.3% in 2012, 21% in 2013, and reaching 640.8 points in 2014. From 2015 to 2018, the SBV further reduced the refinancing rate to 6.25% in 2017 and maintained it through end-2018, while the interbank overnight rate stayed low (≈1%–3%). Against a backdrop of looser policy and favourable macro conditions, the market overheated: the P/E ratio peaked at 21.7x – more than 20% above the Southeast Asian average – before a sharp correction began in 04/2018. External macro headwinds weighed on prices, but the proximate trigger was investor panic selling despite falling rates. The variance decomposition (FEVD) similarly shows that interest-rate factors account for a modest share of bubble dynamics – only about 5%. During 2019–2022, the COVID-19 pandemic prompted monetary easing globally, including in Vietnam, via liquidity injections and rate cuts to support the economy: the refinancing rate declined from 6.25% at the start of 2019 to 4% by end-2021. Under pressure from global policy shifts, Vietnam then raised the refinancing rate three times, to 6% by end-2022. In 2020–2022, the stock market nonetheless experienced a bubble, due in part to the low-rate environment.

Third, the VAR suggests a positive association between inflation and the bubble proxy and a negative association with industrial production. Both effects are economically small – the FEVD indicates that CPI and IIP jointly account for only about 2% of the bubble's forecast-error variance at the typical horizon – so they are better viewed as reflections of the macro-policy backdrop rather than primary drivers. The negative sign on IIP is consistent with several mechanisms. Because our proxy is the price-earnings (P/E) ratio, stronger real activity raises expected earnings (the denominator) faster than prices, compressing P/E and thus the bubble measure. IIP strength also typically coincides with anticipated policy tightening (via the refinancing rate and liquidity management), which lifts discount rates and dampens speculative demand. In addition, portfolio rebalancing toward real investment opportunities can drain flows from speculative equity segments; the sectoral composition of the Vietnamese indices may amplify this effect. The positive CPI sign is consistent with phases in which easier policy and rising prices lower real rates and support valuations, whereas during tightening episodes (e.g., 2022) CPI rises but the market falls – underscoring that CPI mainly captures the policy/macro backdrop rather than causing bubbles. The impulse responses (Figures 3–8) and the Granger results corroborate these channels: a positive IIP shock lowers the bubble proxy on impact, whereas CPI shocks have only small direct effects. While a full robustness exercise is beyond scope here, baseline diagnostics pass and these interpretations are consistent with both the IRFs and the FEVD.

Fourth, we find a weak positive effect of the money supply on stock-market bubbles (≈2.6% in the FEVD). The reverse channel – from bubbles to M2 – is not supported by our Granger tests, consistent with M2 being used primarily as a policy instrument rather than a proximate cause of bubbles. In practice, however, liquidity conditions are tightly linked to market dynamics. During the high-inflation phase of 2010–2012, M2 growth fell sharply; as policy loosened in 2014–2016, the market recovered. Although prices rose only moderately, M2 growth rose steeply and hovered around 17–18% in 2014–2016, then slowed to 14.97% in 2017 and 11.34% in 2018. Following the strong expansion in liquidity from 2016 to April 2018, equities surged by about 135% to a historical peak. Policy adjustments by the State Bank of Vietnam thus shaped both M2 and bubble dynamics. In 2020–2022, M2 growth edged up in 2020–2021 amid rate easing, while pandemic effects slowed real activity and stoked inflation pressures. By late 2022, amid a shift toward tighter policy, credit growth eased to ~9% and the bubble unwound. On balance, M2 and the equity market tend to move in the same direction, but in our model the marginal contribution of M2 to bubble variation is small, consistent with the FEVD.

Fifth, the dominant force behind bubbles is their own internal dynamics. Although the estimated own-lag coefficient is negative – consistent with mean reversion – this should not be read as evidence that bubbles mechanically self-correct; rather, the FEVD shows that shocks to the bubble itself explain about 90% of its variance, in line with behavioural feedback. Early price rises need not draw immediate participation, but sustained gains attract a ‘strong inflow of funds’, pushing valuations higher. Once overextension becomes salient, herding can flip to panic selling and a sharp correction. The 2017–2018 episode – buoyed by IPO activity and FTSE Russell developments – fits this pattern, as does 2020–2022, when idle cash and heightened retail participation produced record liquidity and new accounts before the unwind. In short, bubble formation and collapse are driven largely by investor psychology interacting with transient catalysts, a reading reinforced by the variance decomposition.

7 Conclusions

This paper develops a framework to examine how monetary policy relates to stock-market bubbles in Vietnam over 2010–2022. Bubble-dating tests identify several episodes – especially in 2017–2018 and 2020–2022 – while the VAR indicates that the refinancing rate and the interbank overnight rate lean against bubble formation, money supply is only weakly supportive, and CPI/IIP effects are modest. The impulse-response analysis aligns with these patterns, and variance decomposition suggests that boom–bust dynamics are driven largely by self-reinforcing feedback consistent with behavioural explanations.

For Vietnam’s policy design, the priority is to strengthen price-based transmission within a coherent market structure. Clarifying the operational role of the refinancing rate and its corridor, and linking open-market operations and liquidity management to that anchor, would improve pass-through from interbank rates to market pricing. Deepening collateralised funding and the government-bond market to anchor the yield curve would reduce reliance on administrative controls and let rate signals guide credit. Coordination with market supervisors can temper speculative phases via calibrated limits on securities credit and margin, greater transparency on broker funding and retail flows, and stronger enforcement against manipulation – complements, not substitutes, for the monetary stance. Beyond monetary tools, broader frameworks that widen participation and resilience can complement macro policy; inclusive-entrepreneurship programmes in Wallonia illustrate this potential (Lavison et al., 2023). A multi-dimensional approach – pairing rate-based transmission with investor education, inclusion, and market-integrity efforts – can enhance effectiveness and temper future cycles.

For investors, shifts in the refinancing-rate stance and liquidity conditions are risk-regime changes. Emphasising earnings resilience and market liquidity, using leverage prudently, hedging when breadth deteriorates, and applying pre-set de-risking rules around tightening can limit drawdowns as sentiment turns. Together, such adjustments by policymakers and market participants offer a pragmatic path to moderating bubbles while supporting healthier capital-market development.

This study uses a reduced-form VAR that traces dynamic associations rather than causal effects, so endogeneity among policy rates, liquidity (M2) and the bubble proxy may remain; diagnostics help but do not identify causality. Future work should adopt structural identification (e.g., an SVAR with sign/zero restrictions or external instruments) or narrative/high-frequency approaches using State Bank of Vietnam announcements and surprises – steps that are beyond this paper’s scope. For the same reason, a fuller robustness and sensitivity programme – testing alternative lag structures and orderings, generalised identification, real versus nominal quantities, subsample stability around regime shifts, and alternative operationalisations of the bubble variable (e.g., GSADF-dated in-bubble indicators, price-to-book, cyclically adjusted earnings) – is left to subsequent research, which would also probe counterintuitive signs such as the negative IIP relation using disaggregated evidence (sectoral indices, margin-credit flows).

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Appendix

Table A1 Variable definitions and data sources

Variable	*Definition*	*Unit*	*Expected sign*	*Source*
PE	Price-to-earnings ratio	Ratio	+	Bloomberg
IR	Refinancing rate (SBV policy lending rate)	% p.a.	−	IMF-IFS
R	Overnight interbank interest rate	% p.a.	−	IMF-IFS
M2	Money supply (M2)	Billions of VND	+	IMF-IFS
CPI	Consumer Price Index (headline)	Index	+	IMF-IFS
IIP	Index of Industrial Production	Index	+	IMF-IFS

Table A2 Descriptive statistics of daily closing prices: VNI and HNX, 2010–2022

Year	*Obs.*		*Mean*		*Std. dev.*		*Min.*		*Max.*	
	VNI	*HNX*	*VNI*	*HNX*	*VNI*	*HNX*	*VNI*	*HNX*	*VNI*	*HNX*
2010	250	250	486.05	146.04	30.00	24.72	423.89	97.40	549.50	187.20
2011	248	248	434.59	78.31	40.42	14.30	347.80	56.70	522.60	113.40
2012	250	250	412.83	64.96	30.49	9.56	336.73	50.66	488.07	83.79
2013	250	250	490.11	62.35	17.86	2.28	418.35	57.61	527.97	68.30
2014	247	247	579.95	82.17	29.67	6.04	504.51	67.93	640.75	92.99
2015	248	248	579.89	82.40	23.12	3.37	526.93	73.09	638.69	89.47
2016	251	251	625.91	81.40	47.66	2.90	521.88	73.06	688.89	88.16
2017	250	250	780.17	98.23	76.90	9.57	672.01	81.33	984.24	118.87
2018	250	250	1,008.87	115.20	81.58	10.27	888.69	96.39	1,204.33	138.02
2019	250	250	971.55	104.70	28.15	2.23	878.22	99.97	1,024.91	110.88
2020	252	252	889.47	122.19	90.03	21.64	659.21	92.64	1,103.87	203.12
2021	250	250	1,311.77	328.81	111.57	71.48	1,023.94	203.05	1,500.81	473.99
2022	250	250	1,252.87	315.17	178.87	89.19	911.90	175.78	1,528.57	493.84

Table A3 Descriptive statistics of model variables

Variable	*Obs.*	*Mean*	*Std. dev.*	*Min.*	*Max.*
PE	156	14.61	2.70	8.54	21.46
R	156	3.89	3.41	0.10	13.40
IR	156	7.09	2.71	4.00	15.00
M2	156	7,178,060.00	3,728,602.00	1,951,164.00	14,200,000.00
CPI	156	147.46	21.48	95.96	180.07
IIP	156	2.42	7.53	–22.30	27.60