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Macroeconomic Determinants of the Indonesian Composite Stock Price Index: An ARDL Approach (2016-2025)

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ABSTRACT
This study investigates the short and long run impacts of macroeconomic variables inflation, the BI Rate, the USD/IDR exchange rate, and world gold prices on the Indonesian Composite Stock Price Index (IHSG). Utilizing 120 monthly observations from January 2016 to December 2025, this period accounts for severe market turbulences, including the early 2025 trading halts. Given the mixed integration orders $I(0)$ and $I(1)$), an ARDL bounds testing framework and an Unrestricted Error Correction Model (UECM) with HC3 robust standard errors were applied. The ARDL bounds test reveals no stable long run cointegration (F-statistic $= 2.192$, below the lower critical bound), indicating that macroeconomic indicators do not establish a long term equilibrium with the IHSG. Conversely, short run dynamics demonstrate that contemporaneous rupiah depreciation significantly dampens the IHSG. Lagged inflation exhibits a positive effect, while gold prices stimulate the index concurrently but induce a negative effect in the subsequent month, reflecting a delayed safe haven rotation. The BI Rate shows no significant direct impact once exchange rate dynamics are controlled. These findings imply that the IHSG is primarily driven by immediate currency shocks and commodity price pressures rather than long term macroeconomic anchoring.

Keywords: IHSG, Exchange Rate, Inflation, BI Rate, ARDL.
JEL Code: E44, F31, F37, G15

I. Introduction

The sensitivity of the Indonesian stock market to macroeconomic shocks was prominently illustrated during the 2025–2026 period. In March and April 2025, the IHSG experienced dramatic declines, losing over 5% and 8% in single sessions respectively, which triggered unprecedented dual trading halts on the Indonesia Stock Exchange (IDX) amid a massive foreign capital outflow of approximately Rp22.21 trillion (Tempo, 2025; Kompas, 2025). This equity shock was synchronized with historical pressure on the domestic currency. By April 2025, the rupiah breached Rp16,970 per U.S. dollar surpassing its weakest intraday level during the 1998 Asian financial crisis (CNBC Indonesia, 2025) and continued to depreciate toward Rp18,141 per U.S. dollar by mid 2026 (ANTARA, 2026). In response to these out of cycle disruptions, Bank Indonesia aggressively raised the BI Rate to 5.50% to maintain exchange rate stability (Bank Indonesia, 2026). Concurrently, global financial uncertainty propelled world gold prices to a historic high of USD 4,357 per troy ounce by December 2025, capturing a widespread flight to safe haven assets. Even the exchange's own price stabilization instruments,

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such as the symmetric auto rejection policy, have been shown to alter investor behavior significantly during turbulent episodes on the IDX (Maudya & Yuliati, 2025). These concurrent shocks highlight a critical research problem regarding the nature of the relationship between macroeconomic indicators and equity pricing in an emerging economy. While Arbitrage Pricing Theory (Ross, 1976) and Chen et al. (1986) framework establish macroeconomic variables as systematic risk factors, the exact transmission mechanisms under extreme market stress remain theoretically ambiguous. Dornbusch and Fischer's (1980) flow-oriented model suggests that currency depreciation compresses corporate profits and drives stock prices down, whereas Fama's (1981) proxy hypothesis underscores how inflation shocks interact with real economic activity. Furthermore, the safe haven framework of Baur and Lucey (2010) suggests that portfolio rotations into gold during periods of opacity may drain liquidity from the equity market.

Despite these clear theoretical pathways, domestic empirical evidence remains highly fragmented. For instance, Fauzi and Wijoyo (2025) and Latif et al. (2024) document conflicting impacts of inflation and exchange rates on the IHSG, while Aminarta and Kurniawan (2021) identify macroeconomic influences exclusively in the long run. This empirical divergence stems from three distinct literature gaps. A heavy reliance on static multiple linear regressions that ignore time series nonstationarity, the omission of global uncertainty proxies such as world gold prices, and the lack of coverage extending to extreme market stress regimes. To resolve these inconsistencies, this study employs the ARDL bounds testing framework and the Unrestricted Error Correction Model (UECM) proposed by Pesaran et al. (2001). This dynamic approach is uniquely suited to evaluate both immediate short run variations and long run equilibrium properties within a single framework using monthly data from January 2016 to December 2025. This study provides three distinct empirical and methodological contributions that extend and challenge previous domestic literature. First, by strategically extending the observation window to encompass the dual trading-halt crisis of March–April 2025 and the historical milestone of the rupiah breaching its weakest intraday position, this analysis captures a high stress regime that is entirely absent from prior macroeconomic studies in Indonesia. Second, unlike earlier papers that rely on static models prone to spurious regression, this research introduces a rigorous dynamic framework by incorporating the world gold price as an explicit proxy for global uncertainty and portfolio substitution. Third, this study challenges the conventional consensus regarding stable macroeconomic integration in the Indonesian capital market. By documenting a total absence of long run cointegration alongside highly sensitive, immediate short run currency and commodity shocks, this analysis provides an alternative empirical perspective on emerging market vulnerabilities under structural transitions.

a. Does inflation have a significant short run dynamic effect or a stable long run relationship with Ind onesia's IHSG during the 2016–2025 period?
b. Does the BI Rate have a significant short run dynamic effect or a stable long run relationship with Indonesia's IHSG during the 2016–2025 period?
c. Does the USD/IDR exchange rate have a significant short run dynamic effect or a stable long run relationship with Indonesia's IHSG during the 2016–2025 period?
d. Does the world gold price have a significant short run dynamic effect or a stable long run relationship with Indonesia's IHSG during the 2016–2025 period?
e. Is there a stable long run cointegrating relationship linking inflation, the BI Rate, the USD/IDR exchange rate, world gold prices, and the IHSG jointly?

This study pursues three specific objectives. First, it estimates and distinguishes between the short run dynamic adjustments and long run effects of inflation, the BI Rate, the USD/IDR exchange rate, and the world gold price on the IHSG from January 2016 to December 2025. Second, it formally tests the existence of a stable long run cointegrating equilibrium among these five variables using the ARDL bounds testing framework.

II. Literature Review and Hypothesis Development

2.1. Theoretical Foundations

This study anchors its empirical framework on four interconnected financial and economic theories to explain the dynamic responses of the IHSG to macroeconomic shocks. The overarching architecture is built upon the Arbitrage Pricing Theory (APT) formulated by Ross (1976), which posits that an asset's returns are

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driven by a multi factor structure of systematic risks rather than a single market factor. Empirically extended by Chen et al. (1986), APT dictates that macroeconomic variables directly alter equity valuations through two primary transmission channels. Adjusting the expected future cash flows of listed firms and shifting the discount rates required by investors. Within this multi factor boundary, inflation, the BI Rate, the USD/IDR exchange rate, and global gold prices represent the priced systematic risk factors. Under the APT framework, any unanticipated innovations in these factors disrupt the short run market equilibrium, thereby justifying the formulation that macro indicators dynamically affect the stock price index.

To establish the specific direction of these systematic risks, three complementary transmission models are integrated into the APT baseline. First, the Flow Oriented Model of exchange rates introduced by Dornbusch and Fischer (1980) explains the currency equity nexus. This model assumes that exchange rate movements directly determine international competitiveness and corporate balance sheets. In an economy like Indonesia, a domestic currency depreciation (rupiah weakening) inflates the cost of imported raw materials and intensifies the foreign debt servicing burden of listed corporations. Consequently, compressed profit margins trigger capital flight by foreign institutional investors, driving the stock price index down. Recent studies between 2022 and 2023 strongly reinforce this flow oriented transmission, documenting that severe currency depreciation functions as an immediate negative shock to stock returns in developing Asian markets (Sreenu, 2023). This provides the theoretical foundation for expecting a significant short run negative impact of the USD/IDR exchange rate on the IHSG.

Second, Fama's (1981) Stock Proxy Hypothesis clarifies the dual transmission of inflation and interest rates. Fama asserts that the relationship between inflation and equity returns is proxying for the underlying strength of real economic activity. Under moderate demand pull conditions, inflation acts as a positive signal of macroeconomic expansion, which lifts equity returns, whereas severe cost push inflation destroys purchasing power and prompts aggressive monetary tightening. Recent literature published in 2022–2023 demonstrates that while interest rate hikes exert an immediate dampening effect on stock returns via the cost of capital channel, the direct short run impact of inflation remains mixed depending on whether it signals growth or stagflationary pressures (Sreenu, 2023). In Indonesia, previous findings remain fragmented. Fauzi and Wijoyo (2025) and Latif et al. (2024) document no significant inflation impact, whereas Hijrianti et al. (2024) and Purnamasari et al. (2025) observe a positive effect. These dynamics support the proposition that inflation and the BI Rate generate immediate short run variations in the IHSG.

Third, the Safe Haven Theory popularized by Baur and Lucey (2010) models the role of precious metals during periods of severe macroeconomic opacity. A safe haven is defined as an asset that holds a negative or zero correlation with equities during periods of extreme market stress. When systematic uncertainty escalates such as during the dual trading halts observed in the domestic market investors execute portfolio rotations out of risky equities and into gold to preserve capital. However, because the Indonesia Stock Exchange maintains a high capitalization weight in commodity based and mining corporations, a surge in global gold prices simultaneously yields windfall revenues for mining issuers. Literature from 2022–2023 suggests that this duality creates a complex lagged mechanism where gold shocks initially lift index performance through the corporate cash flow channel before the asset substitution effect takes dominance in subsequent periods (Chauhan et al., 2025; Sreenu, 2023). This justifies the dual directional formulation of the gold equity relationship.

Table 1. Summary of Previous Studies

Author(s)	Variables/Method	Main findings
Fauzi & Wijoyo (2025)	Inflation, exchange rate, interest rate, economic growth, regression	Exchange rate, interest rate, and economic growth are significant, inflation is not significant.
Aminarta & Kurniawan (2021)	Inflation, exchange rate, BI Rate, VECM	No short run effect is found, the exchange rate affects the index in the long run.
Latif et al. (2024)	Global indices, inflation, exchange rate, GARCH	DJIA and SSE have positive effects, the exchange rate has a significant negative effect, inflation is not significant.

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Author(s)	Variables/Method	Main findings
Hijrianti et al. (2024)	Inflation, interest rate, exchange rate, regression	Significant simultaneously, inflation is positive, interest rate and exchange rate are negative.
Purnamasari et al. (2025)	Inflation, exchange rate, interest rate, regression	Inflation is positive and significant, interest rate is negative and significant, exchange rate is not significant.
Siringoringo et al. (2025)	BI Rate, inflation, VAR	The dynamics of the effects vary across the 2019-2023 period.
Demir (2019)	Macroeconomics and BIST-100, ARDL	Growth, domestic currency, and portfolio flows raise the index, interest rates and oil depress it.
Sreenu (2023)	Inflation, exchange rate, market returns, ARDL-GARCH	A long run exchange rate return relationship is found, short run effects may be negative.
Chauhan et al. (2025)	Macroeconomic indicators and Indian sectoral indices, ARDL	Effects differ by short run, long run, and sector.
Peter et al. (2026)	Macroeconomics and Tanzanian stock index, SVAR	Stock market dynamics may be detached from some macroeconomic fundamentals in frontier markets.

Therefore, based on the theoretical frameworks and empirical literature gaps identified above, the hypotheses proposed in this study are formulated as follows:

H1: Inflation affects the IHSG in the short run.
H2: The BI Rate affects the IHSG in the short run.
H3: The USD/IDR exchange rate has a negative effect on the IHSG in the short run.
H4: The world gold price affects the IHSG in the short run.
H5: Inflation, the BI Rate, the USD/IDR exchange rate, gold prices, and the IHSG have a stable long-run cointegrating relationship.

III. Research Method

3.1. Research Design

This study adopts an explanatory quantitative approach based on monthly time series data, which is appropriate for its objective of explaining the dynamic relationship between macroeconomic variables and the IHSG from historical records. Monthly data allow these movements to be traced systematically over time. The analytical sequence follows standard time series econometric practice. The initial stage examines the characteristics of the data, particularly whether the series are stationary. Once the nature of the data has been established, the appropriate model specification is selected. Cointegration is then tested before any long run relationship is interpreted. Finally, diagnostic and stability tests ensure that the model is feasible, consistent, and reliable before conclusions are drawn (Pesaran et al., 2001; Brooks, 2019).

3.2. Data, Period, and Operational Definitions of Variables

The analysis relies on monthly secondary data covering January 2016-December 2025, totaling 120 observations, collected manually from official publications. Statistics Indonesia (2026) for inflation, Bank Indonesia (2026) for the BI Rate and the JISDOR reference exchange rate, Indonesia Stock Exchange (2026) for the IHSG, and World Gold Council (2026) for the spot gold price. A data audit confirmed that no variable contained missing values, and the operational definitions are presented in Table 2.

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Table 2. Operational Definitions of Variables

Variable	Notation	Definition/Transformation	Source
IHSG	Y / LNIHSG	Monthly closing value of the IHSG (points), ln(IHSG) transformation	Indonesia Stock Exchange (2026)
Inflation	X1 / INF	End of month annual inflation (yoy), in percent	Statistics Indonesia (2026)
BI Rate	X2 / BIRATE	Bank Indonesia policy rate at end of month, in percent	Bank Indonesia (2026)
USD/IDR exchange rate	X3 / LNKURS	Monthly average JISDOR rate (Rp/USD), ln(exchange rate) transformation	Bank Indonesia (2026)
World gold price	X4 / LNEMAS	End of month spot gold price (USD/troy oz), ln(gold) transformation	World Gold Council (2026)

Natural logarithm transformation is applied to large-scale variables IHSG, the exchange rate, and gold so that the coefficients can be interpreted as elasticities and the variance becomes more stable, inflation and the BI Rate remain in percentage form because both are already ratio variables.

3.3. Justification for the ARDL Bounds Testing Method

Macroeconomic series generally move over time, either upward or downward, and therefore are often nonstationary in raw form. This must be tested first so that the model does not generate misleading conclusions. Entering nonstationary data directly into an ordinary regression without a cointegration test risks producing spurious results. Statistically respectable on the surface, for example with a high R^2, but lacking valid economic meaning (Gujarati & Porter, 2021). Cointegration analysis offers several established routes. The two step Engle and Granger (1987) procedure, the Johansen (1988) test, and the ARDL bounds testing approach of Pesaran et al. (2001). Engle-Granger and Johansen generally require all variables to share the same order of integration, namely I(1). ARDL is more flexible because it can operate when variables combine I(0) and I(1), provided that none is I(2). ARDL is also suitable for small to moderate samples and allows each variable to have its own lag length. These qualities make it highly appropriate for the present study, especially because the preliminary tests indicate mixed integration orders among the variables. Before estimation, however, the analysis must ensure that no variable is I(2), a check conducted through unit root testing so that the model remains within the accepted boundaries of time series econometrics. To strengthen the cointegration decision for a moderate sample, this study reads the F-statistic against the more conservative critical values of Narayan (2005), alongside the asymptotic values of Pesaran et al. (2001).

3.4. Model Specification

The analytical backbone is the Autoregressive Distributed Lag (ARDL) model, selected because it can capture the relationship between the IHSG and macroeconomic variables in both the short and long run. In this model, the index responds not only to contemporaneous macroeconomic variables but also to its own past values and those of the explanatory variables. The general form is:

$$LNIHSG_t = \alpha_0 + \Sigma\varphi_i\, LNIHSG_{t-i} + \Sigma\beta_j\, INF_{t-j} + \Sigma\gamma_j\, BIRATE_{t-j} + \Sigma\delta_j\, LNKURS_{t-j} + \Sigma\theta_j\, LNEMAS_{t-j} + \varepsilon_t$$

To determine whether the variables share a long run relationship, this study conducts the ARDL bounds test through the Unrestricted Error Correction Model (UECM), a form particularly suitable for testing long run equilibrium. The UECM is written as:

$$\Delta LNIHSG_t = \alpha_0 + \lambda_1 LNIHSG_{t-1} + \lambda_2 INF_{t-1} + \lambda_3 BIRATE_{t-1} + \lambda_4 LNKURS_{t-1} + \lambda_5 LNEMAS_{t-1} + \Sigma\beta_j' \Delta X_{t-j} + \varepsilon_t$$

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Here, Δ denotes short run change, while the lagged level terms carry the test for the long-run relationship. The model therefore addresses two questions simultaneously. Short run dynamics and the possibility of a long run link among the IHSG, inflation, the BI Rate, the USD/IDR exchange rate, and the world gold price. The null hypothesis of the bounds test states that there is no long run relationship among the variables, formally written as:

$$H_0: \lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = 0$$

In other words, if all level coefficients equal zero, there is no long run relationship binding the IHSG to the macroeconomic variables in the model. The test uses the F-statistic under Case III, namely a model with an intercept and no trend, following Pesaran et al. (2001). The decision depends on where the F statistic falls relative to the I(0) lower bound and the I(1) upper bound. Above the I(1) upper bound, the variables are cointegrated, and the long run coefficients together with the error correction term may be interpreted. Below the I(0) lower bound, there is no long run relationship, and attention narrows to the short run. If the statistic falls between the two bounds, the result is considered inconclusive.

3.5. Data Analysis Procedures

The analysis proceeds sequentially. The opening stage is a data audit, examining the number of observations, sample range, completeness, and unit consistency of each variable, a step that secures data reliability before any analysis is performed. Next, descriptive statistics are used to describe the initial profile of each variable, namely the mean, minimum, maximum, and dispersion, while also identifying potential extreme values hidden in the IHSG, inflation, BI Rate, exchange rate, and world gold price series. To support full replicability, all estimations unit root tests, lag selection, the bounds test, the UECM, and every diagnostic were computed in Python 3 using the statsmodels library, with lag selection capped at four and all specification settings reported in this section. The IHSG, USD/IDR exchange rate, and world gold price are then converted into natural logarithms, which smooth the data scale and facilitate interpretation across variables. Inflation and the BI Rate remain in percentages because they are already expressed in that form.

Stationarity is examined next. The Augmented Dickey-Fuller (ADF) test is applied at level and first difference, while the KPSS test is introduced as a countercheck to confirm the order of integration. Combining the two is helpful because their null hypotheses differ. ADF takes nonstationarity as the null, whereas KPSS takes stationarity as the null. Read together, they reveal whether a variable is I(0), I(1), or must be ruled out as I(2) (Kwiatkowski et al., 1992). The lag length in the ARDL model is then determined using the Akaike Information Criterion (AIC), with a maximum of four lags. Each core variable is retained in the model because each maps directly onto the research hypotheses, H1 to H4. The bounds test then asks whether a long run relationship binds the variables. The test is conducted under Case III, a specification with an intercept and no trend, and the F statistic is evaluated against the critical values of Pesaran et al. (2001) and Narayan (2005), the latter serving as a comparison for small to moderate samples.

The study then estimates the Unrestricted Error Correction Model (UECM). When the bounds test fails to confirm cointegration, the long run coefficients are not interpreted, and the discussion remains focused on the short run relationship. Residual diagnostics follow to ensure that the model satisfies its econometric assumptions. the Jarque Bera normality test, the Breusch Pagan heteroskedasticity test, the Breusch Godfrey LM autocorrelation test at lags 2 and 12, the Ramsey RESET specification test, and the Durbin Watson statistic. If heteroskedasticity appears, HC3 robust standard errors keep the inference reliable (MacKinnon & White, 1985). The CUSUM parameter stability test closes the procedure, examining whether the model remains stable throughout the sample rather than breaking under structural change, a real concern given the major shocks included in this period, including the 2020 and 2025 market episodes.

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IV. Results and Discussion

4.1. Descriptive Statistics

Table 3. Descriptive Statistics (n = 120, January 2016-December 2025)

Statistic	Inflation (%)	BI Rate (%)	Exchange rate (Rp)	Gold (USD)	IHSG (points)
Mean	2.93	4.98	14,639.50	1,846.31	6,326.71
Median	2.99	5.00	14,403.65	1,755.70	6,314.01
Minimum	-0.09	3.50	13,017.24	1,116.40	4,538.93
Maximum	5.95	7.25	16,827.06	4,357.10	8,646.94
Standard deviation	1.12	0.95	1,018.29	687.35	862.03

The descriptive statistics clarify the extent of data shifts across the sample. Inflation ranged from deflation of -0.09% in February 2025 to 5.95% in September 2022, while the BI Rate also varied widely, from its pandemic era low of 3.50% to 7.25%. The rupiah also came under pressure throughout the period. It moved from Rp13,017 per U.S. dollar to the highest monthly average of Rp16,827 in April 2025, the same month in which it reached Rp16,970 per U.S. dollar intraday, providing evidence of substantial currency pressure, particularly during 2025. World gold prices rose sharply, from USD1,116 to USD4,357 per troy ounce, an increase that marked its growing appeal as a refuge whenever economic and financial conditions became uncertain. The IHSG, meanwhile, followed a volatile path. It reached a trough of 4,538.93 during the March 2020 pandemic panic before recovering to a record of 8,646.94 in December 2025. Along the way, it also fell to 6,270.60 in February 2025, just before the trading halt episode. This wide range indicates that the sample covers several distinct market states, from the pandemic and recovery to exchange rate pressure and broader financial uncertainty. This alone requires a method designed to handle the nonstationarity typical of time series data.

4.2. Stationarity Tests (ADF and KPSS)

Table 4. Results of the ADF Unit-Root Test and KPSS Test

Variable	ADF level	p-value	ADF diff-1	p-value	KPSS level
LNIHSG	-1.436	0.565	-9.786	0.000	1.281**
INF	-2.741	0.067	-4.455	0.000	0.248
BIRATE	-1.758	0.402	-6.856	0.000	0.224
LNKURS	-0.915	0.783	-10.564	0.000	1.616**
LNEMAS	1.785	0.998	-10.491	0.000	1.596**

The stationarity results show that all variables become stationary once first differenced. The Augmented Dickey Fuller (ADF) test clarifies this pattern: series that are nonstationary at level become stationary at first difference. As a confirmatory check, this study adds the KPSS test. According to this reading, LNIHSG, LNKURS, and LNEMAS are I(1), nonstationary at level but stationary after first differencing, a conclusion consistent with the ADF results that retain a unit root at level and the KPSS results that reject level stationarity. INF and BIRATE, by contrast, appear to be I(0) or borderline. For INF, the level ADF result is only marginally significant at 10%, and the KPSS results for both INF and BIRATE do not reject stationarity, allowing both to be treated as relatively stationary at level. The final result is a set of variables with mixed orders of integration, partly I(0) and partly I(1). This pattern is precisely the target of the ARDL bounds testing method because it accommodates a combination of I(0) and I(1) series, where the Engle-Granger and Johansen approaches are less suitable.Most

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importantly, no variable is I(2). With the main ARDL prerequisite satisfied, the model can proceed to cointegration testing through the bounds approach.

4.3. Optimal Lag Determination

Lag length is selected using the Akaike Information Criterion (AIC), with a maximum of four lags. The selected specification is ARDL(1, 2, 1, 3, 2) for the sequence LNIHSG, INF, BIRATE, LNKURS, and LNEMAS, meaning one lag for LNIHSG, two for INF, one for BIRATE, three for LNKURS, and two for LNEMAS. This combination records the lowest AIC among the candidate models examined. During selection, the BI Rate is retained with at least one lag so that the second hypothesis, namely its effect on the IHSG, can still be tested formally. The reasoning is straightforward and rests on the principle that a variable central to a research hypothesis should not be removed solely on statistical grounds. In addition, the lag structure remains sufficiently parsimonious. This is important with only 120 monthly observations because controlled lag length captures the dynamics linking the variables without exhausting too many degrees of freedom.

4.4. ARDL Bounds Test

Table 5. Results of the Bounds Cointegration Test (Case III, k = 4)

Description	I(0) lower bound	I(1) upper bound
10% critical value	2.45	3.52
5% critical value	2.86	4.01
1% critical value	3.74	5.06
Calculated F-statistic	2.192	

Asymptotic critical values are from Pesaran et al. (2001), Table CI(iii). The small sample values of Narayan (2005) are higher and therefore strengthen the conclusion. Source: research data, processed by the authors.

At 2.192, the F statistic lies below the I(0) lower bound at the 5% level (2.86), and even below the 10% lower bound. The null hypothesis of no level relationship therefore remains in force: cointegration among the IHSG, inflation, the BI Rate, the exchange rate, and gold is not confirmed (H5 is rejected). The methodological consequence is fully observed. Long run coefficients are not interpreted, and the analysis remains focused on short run dynamics in UECM form. This discipline avoids a common error, namely interpreting a long run relationship that the data do not support.

4.5. UECM Estimation with HC3 Robust Standard Errors

Because the Breusch Pagan test indicates heteroskedasticity (Section 4.6), all inference relies on HC3 robust standard errors. Table 6 presents the estimation results ($R^2 = 0.349$; adjusted $R^2 = 0.267$; model F test $p < 0.001$; effective n = 117).

Table 6. UECM Estimation Results for ARDL(1,2,1,3,2), Dependent Variable ΔLNIHSG

Variable	Coefficient	HC3 standard error	z	p-value
Constant	0.7456	1.3050	0.571	0.5690
LNIHSG(-1)	-0.1146	0.0455	-2.516	0.0134
INF(-1)	0.0020	0.0034	0.568	0.5712
BIRATE(-1)	0.0015	0.0052	0.300	0.7651
LNKURS(-1)	-0.0110	0.1704	-0.064	0.9488

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Variable	Coefficient	HC3 standard error	z	p-value
LNEMAS(-1)	0.0479	0.0305	1.570	0.1195
ΔINF	0.0060	0.0105	0.566	0.5723
ΔINF(-1)	0.0285	0.0078	3.663	0.0004
ΔBIRATE	0.0225	0.0244	0.920	0.3595
ΔLNKURS	-1.0933	0.5411	-2.020	0.0459
ΔLNKURS(-1)	0.0015	0.2196	0.007	0.9945
ΔLNKURS(-2)	-0.4274	0.2251	-1.899	0.0604
ΔLNEMAS	0.1880	0.0903	2.083	0.0398
ΔLNEMAS(-1)	-0.2236	0.1131	-1.977	0.0507

Four key findings emerge from the estimation. The first concerns the exchange rate. Its contemporaneous change imposes a negative and significant burden on the IHSG, with a ΔLNKURS coefficient of -1.0933 (significant at 5%). In other words, a 1% rupiah depreciation in the same month tends to reduce the IHSG by about 1.09% in that month. The burden also persists, reappearing at the second lag with a coefficient of -0.4274, significant at 10%, indicating that exchange rate pressure is felt not only immediately but also in subsequent months. Second, the first lag of changes in inflation lifts the IHSG strongly. The ΔINF(-1) coefficient of 0.0285 implies that a one percentage point increase in the change in inflation in the previous month is followed by an approximately 2.85% increase in the IHSG in the following month. This may reflect that, across the sample, certain inflation increases were still read as signs of strengthening activity and therefore elicited a positive market response. Third, gold prices behave differently immediately and after a short delay. In the current month, changes in gold prices raise the IHSG, with a coefficient of 0.1880, significant at 5%. At the first lag, the sign reverses to negative, with a coefficient of -0.2236, nearly significant at 5%. This pattern suggests that gold and the index may initially rise together before portfolio rotation from equities to safer assets takes over one month later. Fourth, the BI Rate does not record a significant effect on the IHSG, either in lagged level form or as a contemporaneous change. In this model, therefore, changes in the BI Rate are not shown to move the index directly in the short run. The LNIHSG(-1) coefficient of 0.1146, significant at 5%, points to a dynamic correction mechanism of about 11.5% per month. Because the cointegration test does not confirm a long run relationship, however, this coefficient is not read as the speed of adjustment toward long run equilibrium. It is more appropriately understood as persistence, or a short run adjustment pattern, in the IHSG.

4.6. Diagnostic and Model Stability Tests

Table 7. Results of Diagnostic and Stability Tests

Test	Statistic	p-value	Conclusion
Jarque Bera (normality)	20.708	0.0000	Residuals are nonnormal, inference relies on asymptotic properties and robust standard errors
Breusch Pagan (heteroskedasticity)	24.646	0.0257	Heteroskedasticity detected. HC3 is used
Breusch Godfrey LM lag 2	4.012	0.1345	No autocorrelation
Breusch Godfrey LM lag 12	14.708	0.2578	No autocorrelation up to lag 12
Ramsey RESET	4.470	0.0369	Indication of limitations in the linear functional form

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Test	Statistic	p-value	Conclusion
Durbin Watson	2.046	—	Close to 2, consistent with BG
CUSUM (parameter stability)	0.726	0.6678	Parameters are stable across the sample period

The diagnostics show that the model is free from autocorrelation up to lag 12, an important point because it means the residuals do not contain serial correlation patterns capable of undermining the estimation. The CUSUM parameter stability test also finds that the model is stable throughout the full period, with a p-value of 0.668. This is reassuring given the major shocks embedded in the sample, including the 2020 pandemic and the 2025 market pressure. There is no sign of structural change that would erode the estimation. Residual nonnormality is understandable because the data contain several extreme episodes, namely trading halts and very sharp rupiah depreciation, which may pull the residual distribution away from normality. This concern is mitigated, however, by 117 effective observations, which are sufficiently adequate for inference to rely on asymptotic arguments. Heteroskedasticity is addressed using HC3 robust standard errors, which keep coefficient testing robust and reliable even when residual variances differ. The Ramsey RESET test, for its part, suggests possible nonlinearity; for example, the IHSG response to rupiah depreciation may not mirror its response to appreciation. This does not invalidate the main results. It simply marks a direction for future research. A natural next step is the NARDL approach, which can test asymmetric effects among the variables. Therefore, subsequent research may consider the NARDL approach to test possible asymmetric effects among the variables.

4.7. Discussion

The USD/IDR exchange rate, namely the rupiah against the U.S. dollar, is shown to have a negative and significant effect on the IHSG, therefore, H3 is accepted. This is among the strongest findings of the study and speaks directly to the phenomenon motivating the analysis. In March-April 2025, the rupiah's decline to the intraday level of Rp16,970 per U.S. dollar coincided with sharp pressure in the stock market, including two trading halts and substantial foreign net selling. The UECM estimation indicates that the link between rupiah weakening and IHSG decline was not a one off event. it persisted systematically throughout 2016-2025. Theoretically, this result aligns with the flow oriented view of Dornbusch and Fischer (1980). Rupiah depreciation can raise import costs and foreign debt burdens, especially for issuers with foreign currency liabilities, reducing expected profits and sharpening risk perception. Foreign investors may then withdraw, dragging prices down, while those outflows can further deepen pressure on the currency, creating feedback between rupiah weakness and IHSG decline. This finding is consistent with Latif et al. (2024) and Sreenu (2023) and reinforces Aminarta and Kurniawan's (2021) point that the exchange rate is among the most influential macroeconomic variables for the IHSG.

Inflation also moves the IHSG, so H1 is accepted, with this study finding a positive effect of changes in inflation at a one month lag. Average inflation during the period remained low to moderate, around 2.93%, and in such conditions an increase was not always read negatively: controlled inflation can signal strengthening demand and activity, thereby eliciting a positive market response. This reading is consistent with Fama's (1981) proxy hypothesis as well as the findings of Purnamasari et al. (2025) and Hijrianti et al. (2024). The February 2025 deflation of -0.09%, which coincided with the IHSG falling to 6,270, sharpens the point that inflation pushed too low can be recorded as weakening demand. The world gold price also affects the IHSG, hence, H4 is accepted, and it operates in two directions. Positive in the current month and negative in the following month. The contemporaneous boost works through improved performance of commodity and precious metal shares on the IDX when gold rises, raising profit expectations for issuers linked to the metal.

The negative effect one month later, however, indicates a different mechanism. Over time, higher gold prices may persuade investors to move out of equities and into gold as a safer asset, consistent with the safe haven logic of Baur and Lucey (2010). The effect of gold on the IHSG is therefore not unidirectional. It depends

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on the exchange's sectoral composition and the horizon examined, echoing Demir (2019) and Chauhan et al. (2025), who link the commodity index relationship to market structure and time frame.

The BI Rate does not show a significant direct effect on the IHSG, so H2 is rejected. However, this does not mean that Bank Indonesia's interest rate policy is irrelevant for the market. The result is better read as evidence that interest rates affect the index indirectly, primarily through the rupiah. When Bank Indonesia raises interest rates, the usual objective is to stabilize the currency and ease external pressure. If the effect on equities flows mainly through the exchange rate, then the direct BI Rate coefficient in a model already including the exchange rate may become insignificant. This finding is consistent with Aminarta and Kurniawan (2021) and Siringoringo et al. (2025), who show that the BI Rate's effect on the market shifts across periods. The results also leave long run cointegration unconfirmed. Therefore, H5 is rejected. During 2016-2025, in other words, the IHSG does not have a robust long run equilibrium with inflation, the BI Rate, the USD/IDR exchange rate, and the world gold price jointly. This result is unsurprising because the index responds to far more than these four series, including issuer earnings performance, foreign fund flows, global equity market conditions, energy prices, fiscal policy, and changes in the sectoral structure of the Indonesia Stock Exchange. Firm level evidence from the IDX reinforces this point. Stock prices and returns of listed issuers are strongly shaped by fundamentals such as profitability, liquidity, and solvency rather than by macroeconomic aggregates alone (Prabowo & Vidada, 2026; Widianto & Bagana, 2025). Read in this way, the effect of macroeconomic variables on the IHSG in this study is best understood as a dynamic short run relationship rather than a long run equilibrium. This is consistent with the behavior of the Indonesian market across the sample, especially during major shocks, namely the pandemic, exchange rate pressure, trading halts, and broader uncertainty in 2025.

4.8. Linking the Findings to the 2025-2026 Phenomenon

These findings offer a reasonably coherent explanation for recent market events. The March-April 2025 episode, when the IHSG weakened and triggered two trading halts, occurred as the rupiah sank to a very weak level, an extreme manifestation of the negative exchange rate IHSG relationship found significant by the model. Currency pressure continuing into 2026 confirms that the exchange rate remains a major source of risk for Indonesia's capital market. Bank Indonesia data place JISDOR at Rp18,141 per U.S. dollar on 9 June 2026, the same day that Bank Indonesia again raised the BI Rate by 25 basis points to 5.50% in an effort to stabilize the rupiah. Throughout the first half of 2026, the rupiah successively breached Rp17,000 and Rp17,308 per U.S. dollar before reaching Rp18,141 (CNBC Indonesia, 2026; Tempo, 2026). For investors, the key implication is that exchange rate movements and world gold prices deserve close attention when assessing the near term direction of the IHSG. Both often provide earlier warnings of changing market risk than the BI Rate level alone, because the policy rate operates largely through indirect channels, especially its influence on the exchange rate and capital flows. For monetary authorities, the result reaffirms the value of policies directed at exchange rate stability. A stable rupiah matters not only for the external sector but also for maintaining investor confidence in equities, because the exchange rate and the index both respond to foreign flows, risk perception, and confidence in domestic economic conditions.

V. Conclusion

Built on 120 monthly observations from January 2016 to December 2025 drawn from official sources, this study first screened the data through stationarity tests. The ADF results, confirmed by KPSS, show that the variables combine I(0) and I(1) orders with none at I(2), establishing ARDL bounds testing as an appropriate tool because one of its main requirements is satisfied. The bounds test does not reveal long run cointegration among the IHSG, inflation, the BI Rate, the USD/IDR exchange rate, and the world gold price. With an F statistic of 2.192, below the I(0) lower bound at 5%, the long run relationship is not confirmed and long run coefficients are not interpreted. On this basis, H5, the long run hypothesis, is rejected. Without a long run relationship to rely on, the discussion centers on short run dynamics. The HC3 estimation shows that contemporaneous USD/IDR depreciation places a negative and significant burden on the IHSG, so H3 is accepted, meaning that rupiah weakening tends to be followed by an IHSG decline in the same period. The first lag of changes in inflation has a significant positive effect, so H1 is accepted, indicating that prior period inflation changes may still elicit a positive market response when inflation remains sufficiently controlled.

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The world gold price also affects the IHSG, so H4 is accepted, although its effect changes over time. In the current month, gold raises the index, in the following period, it pushes it downward. This points to a very short run positive effect, operating mainly through commodity and precious metal stocks, which then gives way to later fund rotation into safer assets. The BI Rate, by contrast, does not show a significant direct effect on the IHSG, so H2 is rejected, suggesting that interest rates reach the index more indirectly, particularly through the rupiah. Diagnostics report a model free from autocorrelation up to lag 12 and passing the CUSUM stability test, so the estimation is sufficiently stable and not the product of structural breaks, even though the sample spans major shocks such as the 2020 pandemic and the 2025 trading-halt episode. Overall, this study finds that IHSG movements during 2016-2025 are more responsive to rupiah pressure and world gold price dynamics than to the direct effect of the BI Rate. The evidence marks exchange rate pressure and gold movements as key variables for reading the short-run direction of the index, and it aligns with the IHSG decline and rupiah depreciation observed in 2025-2026, although the 2026 segment is more appropriately treated as a post sample illustration than as part of the core estimation.

For investors, the message is not to rely on a single indicator. A multi signal approach is more prudent, with special weight on the USD/IDR exchange rate and the world gold price, both of which can provide early readings of pressure building on the IHSG. Particular caution is needed when the rupiah weakens rapidly because this study links rupiah softness to IHSG declines in the same month over the historical record. For monetary and fiscal authorities, these results underscore the centrality of rupiah stability for a resilient capital market. The dominant role of the exchange rate indicates that a weaker rupiah can raise market risk through foreign flows, investor expectations, and perceptions of domestic risk. Policies aimed at keeping the rupiah stable, including pro stability interest rate policy, may therefore serve as an important lever for easing market pressure. Nevertheless, because the sample ends in December 2025, Bank Indonesia's 2026 measures are best framed as post sample illustrations rather than part of the core estimation. Clear and credible policy communication is also important so that the market does not overreact to exchange rate movements or shifts in monetary policy. Good communication can help prevent feedback between a weakening rupiah and declining stock prices. For academics, this study is a reminder to move through time series analysis stages with discipline. Claims about long run relationships are justified only when cointegration tests actually find them, without such evidence, the discussion should remain within short run dynamics. Estimation also requires diagnostic support so that the resulting conclusions are valid and accountable.

This study carries several limitations that should be noted. First, the sample ends in December 2025, so anything after that lies outside the estimation. The rupiah's escalation to Rp18,141 per U.S. dollar and successive BI Rate increases in 2026 are therefore not modeled directly. Updating the data once the 2026 record is complete will be important to determine whether these findings hold under stronger exchange rate pressure. The model also relies on only four macroeconomic variables inflation, the BI Rate, the USD/IDR exchange rate, and the world gold price whereas the IHSG certainly responds to more factors, including foreign capital flows, global indices, world oil prices, the money supply, bond yields, and one off events such as the pandemic and trading halts. Omitting these variables may help explain the imperfect specification indicated by the Ramsey RESET result, so adding them is a reasonable way to make future models more complete. Another limitation lies in the residuals, which the tests find to be nonnormal and heteroskedastic. Asymptotic inference and HC3 robust standard errors blunt this issue, but the results still require cautious interpretation. The conclusions are therefore best understood as short-run empirical relationships rather than rigid causal claims. In view of these limitations, future work should widen the methodological net. The NARDL model can test whether rupiah depreciation and appreciation affect the IHSG asymmetrically, a real possibility signalled by the RESET test. GARCH or DCC GARCH models can examine volatility dynamics and intervariable linkages across a turbulent period. Cointegration tests that allow structural breaks, such as Gregory Hansen, would also be suitable for a sample that crosses several market regimes, including the pandemic, exchange rate pressure, and trading halt episodes. Volatility based studies of the IHSG indicate that such second moment dynamics are empirically substantial (Mukmin & Firmansyah, 2020).

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