# Exchange Rate and World Price as Drivers of Soybean Import Dependence in Indonesia: Evidence from a Time-Series Analysis

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## ABSTRACT

Much of the existing literature on Indonesia's soybean import dependency centers on tariff policy and domestic production, treating macroeconomic variables—particularly the IDR/USD exchange rate and world soybean prices—as peripheral controls. This study positions the exchange rate and world price as the primary focus of analysis of Indonesia's soybean import volume determinants, while still controlling for import tariffs and domestic production, using annual time-series data for 2005–2023 (N = 19) sourced from BPS, the WTO, INSW, CBOT, and Bank Indonesia. The econometric approach includes the Augmented Dickey-Fuller unit root test, the two-step Engle-Granger cointegration procedure, three OLS specifications (baseline, extended, and first-difference models) estimated with Newey-West HAC standard errors, and a full diagnostic battery (Jarque-Bera, Breusch-Pagan, Durbin-Watson, VIF). The results show that the exchange rate is the dominant macroeconomic determinant: in the long run, rupiah depreciation is positively and significantly associated with higher import volumes (elasticity 1.093; $p < 0.001$), contrary to the standard expenditure-switching prediction, whereas in the short run (first-difference model) depreciation shows a marginally negative effect ($\beta = -0.687$; $p = 0.075$) consistent with that theory. World soybean prices show a marginally significant positive long-run effect ($\beta = 0.290$; $p = 0.052$) but no short-run significance ($\beta = 0.048$; $p = 0.678$). Unit root tests confirm that all continuous variables are integrated of order one, I(1), while the Engle-Granger cointegration test (residual ADF = −3.231) finds no evidence of cointegration at the 5 percent level, making the first-difference model the primary reference for short-run inference. This pattern indicates that demand from the tofu and tempeh processing industry is highly inelastic due to the absence of adequate domestic substitutes, so that price and exchange-rate shocks are absorbed through import volume. The findings imply that macroeconomic exchange-rate stability and global supply-chain diversification are essential components of soybean food-security policy, alongside supply-side production measures.

**Keywords:** exchange rate; world price; inelastic import demand; soybean; time-series econometrics

## INTRODUCTION

Indonesia's dependence on soybean imports — now exceeding 80 percent of total national consumption — is typically explained through two policy lenses: tariff protection and domestic production capacity. BPS data show that national soybean consumption has risen to around 2.6–2.7 million metric tons per year, while domestic production has fallen from a peak of roughly 975,000 tons (2009) to 213,000–250,000 tons (2021–2023), a structural decline of more than 75 percent over fifteen years. Nevertheless, nearly all of Indonesia's soybean import transactions are denominated in US dollars, following global commodity pricing on the Chicago Board of Trade (CBOT), so that fluctuations in the IDR/USD exchange rate directly affect import costs in domestic currency terms. Over 2005–2023, the rupiah depreciated by more than 57 percent in nominal terms against the US dollar, from around Rp 9,020 per USD (2006) to Rp 15,236 per USD (2023), while world soybean prices also experienced significant volatility, including spikes driven by the 2007–2008 food crisis and post-COVID-19 supply disruptions in 2021–2022.

Standard open-economy macroeconomic theory predicts that exchange-rate depreciation, by raising the domestic price of imported goods, should reduce import volume through the *expenditure-switching* effect (Krugman, Obstfeld, & Melitz, 2018). However, the empirical validity of this prediction for food commodities with highly inelastic domestic supply — such as soybeans in Indonesia — remains contested. If domestic production cannot respond quickly to price signals and imported goods lack adequate domestic substitutes, import demand may remain high even as import costs rise due to depreciation.

A number of empirical studies have examined the determinants of Indonesia's soybean imports. Ningrum, Irianto, and Riptanti (2018) identify domestic production as the primary driver of imports. Malik and Nainggolan (2020) confirm a negative production–import relationship using a panel and instrumental-variable approach. Zikri, Safrida, and Susanti (2020) document the role of price competitiveness as a structural determinant. However, few studies place the exchange rate at the substantive center of the analysis. Firdaus (2018) is one important exception, reporting a positive and significant exchange-rate coefficient — counterintuitive relative to standard theoretical predictions — but without a systematic explanation of the underlying mechanism. Dartanto and Usman (2011) show that world soybean price shocks are transmitted to domestic prices through processing-industry margins rather than being fully passed on to consumers, suggesting that the response of import volume to price changes may be weaker than standard models predict.

Three important gaps remain in the literature. *First*, most studies treat the exchange rate and world price as peripheral control variables rather than substantive objects of analysis. *Second*, most studies use data ending before the 2013–2018 rupiah depreciation episode and the post-pandemic 2021–2023 price cycle. *Third*, many studies do not apply formal time-series diagnostics (unit root and cointegration tests) required for valid econometric inference on nonstationary data.

This study addresses all three gaps by placing the IDR/USD exchange rate and world soybean price at the center of the analysis of Indonesia's soybean import volume determinants, updating the data through 2023, and applying a full set of formal time-series diagnostics. Two research questions are posed: *first*, what is the direction and magnitude of the exchange rate's effect on soybean import volume in the long run and short run? *Second*, to what extent does the world soybean price contribute to variation in import volume, and is the pattern consistent with standard demand theory?

## LITERATURE REVIEW

### Theoretical Framework: Exchange Rate, World Price, and Import Demand

In commodity markets where import payments are denominated in foreign currency, depreciation of the domestic currency directly raises import costs and, according to standard open-economy theory, should reduce import volume through the price effect (Krugman, Obstfeld, & Melitz, 2018). However, when domestic supply is inelastic and imported goods lack close domestic substitutes, the price response of import demand can be weak or even reversed: downstream consumers and producers have no alternative domestic source to switch to, so import demand remains high even as its cost rises. The Armington (1969) model offers a theoretical refinement by treating goods from different countries as imperfect substitutes: demand for imported soybeans is determined not only by relative prices but also by technical and quality preferences — highly relevant to Indonesia's tempeh and tofu processing industry, which has developed production processes tailored to the specific protein and moisture characteristics of imported varieties from the United States and Brazil. This condition is particularly relevant to Indonesia's soybean market, where the tempeh and tofu processing industry — which absorbs around 70–80 percent of national soybean consumption — has demand that is relatively insensitive to medium-term price signals.

The computable general equilibrium (CGE) framework of Dartanto and Usman (2011) shows that changes in import tariffs and world soybean prices affect domestic food prices through multiple channels, with non-trivial welfare implications for low-income households — underscoring the importance of analyzing the price and exchange-rate channels explicitly rather than merely as controls.

### Empirical Evidence on the Exchange Rate and World Price in Indonesia's Soybean Imports

Firdaus (2018) incorporated the IDR/USD exchange rate into a model of Indonesia's soybean import determinants and found a positive and significant relationship — periods of depreciation coincided with higher, not lower, import volumes. This result is consistent with an income-effect and supply-side interpretation: currency depreciation episodes typically coincide with broader macroeconomic pressure and reduced domestic agricultural investment, which deepens import dependence through a supply-side channel that dominates the price effect on demand. Zikri, Safrida, and Susanti (2020) document that the landed cost of soybeans imported from the United States and Brazil, even after accounting for transportation costs and import duties, is consistently lower than the cost of domestic soybean production — a structural gap not easily closed through exchange-rate movements or tariff policy alone.

At a more macro level, the OECD and FAO (2024) document that currency depreciation in food-importing developing countries contributes to a widening gap between domestic food price inflation and international commodity prices, with implications for household food security. The OECD-FAO Agricultural Outlook (2024) also projects that demand for oilseeds, particularly soybeans, will continue to grow rapidly in Asia through 2033, implying continued or increasing dependence on global oilseed markets for net-importing countries such as Indonesia absent large-scale intervention, including on the dimension of macroeconomic stabilization.

### Domestic Production and Tariffs as Control Variables

Although not the primary focus, domestic production and import tariffs are nonetheless included as control variables to isolate the net contribution of the exchange rate and world price. Indonesia's soybean production has undergone a prolonged structural decline, from a peak of 975,000 metric tons (2009) to below 250,000 metric tons (2022–2023) — a decline of more than 75 percent — driven by competition for land with rice and maize and productivity levels far below those of the United States and Brazil (Ningrum, Irianto, & Riptanti, 2018; Zikri, Safrida, & Susanti, 2020). Indonesia's MFN tariff on soybeans has been kept near zero since post-1997–1998 crisis liberalization, with minimal variation (WTO, 2023), leaving very limited data variation with which to separately identify the tariff effect (Malik & Nainggolan, 2020).

### Research Gap and Positioning of the Study

The exchange rate and world price have thus far been treated mostly as peripheral control variables rather than as a substantive focus in the Indonesian soybean literature, and most studies use data ending before the 2013–2018 rupiah depreciation episode and the post-pandemic 2021–2023 price cycle. In addition, most studies do not apply formal time-series econometric procedures — unit root testing, cointegration analysis, and diagnostic testing for heteroskedasticity and serial correlation — that are needed to draw valid conclusions from annual macroeconomic data. This study fills these gaps by placing the exchange rate and world price at the center of the analysis, updating the data coverage through 2023, and applying a full set of formal time-series diagnostics.

## RESEARCH METHODS

### Data Sources and Sample Period

Annual time-series data for the period 2005–2023 (N = 19 observations) are used in this study.

Table 1. Variable Definitions and Data Sources

| Variable | Symbol | Definition | Source | Unit |
|---|---|---|---|---|
| Import volume | IM | Annual soybean import volume (HS 1201) | BPS — Foreign Trade Statistics | Metric tons |

| | | | | |
|---|---|---|---|---|
| World soybean price | PRICE | CBOT spot price, annual average, USD/MT | FAO Food Price Index; CBOT | USD per metric ton |
| Exchange rate | EX | Annual average IDR/USD | Bank Indonesia — SEKI | IDR per 1 USD |
| Domestic production (control) | PROD | Annual soybean production across all provinces | BPS — Agricultural Statistics | Metric tons |
| Import tariff (control) | TAR | Applied MFN tariff for HS 1201.90 | WTO Tariff Profile; INSW | Percentage (%) |

*Note: BPS = Statistics Indonesia; INSW = Indonesia National Single Window; CBOT = Chicago Board of Trade; SEKI = Indonesian Economic and Financial Statistics; MFN = Most-Favoured Nation; HS = Harmonised System.*

All monetary variables are deflated using Indonesia's GDP deflator with a 2010 base year (World Bank, 2024), and continuous variables are transformed into natural logarithms to facilitate elasticity interpretation and reduce heteroskedasticity. The tariff variable (TAR) is retained in level form given its discrete nature and near-zero value over most of the sample period.

**Theoretical Model and Variable Specification**

The empirical model is based on the partial equilibrium theory of import demand (Houthakker & Magee, 1969):

$IM = f\,(TAR, PROD, PRICE, EX)$ (1)

Theoretically expected signs: dIM/dPRICE < 0 (a higher world price raises import costs, reducing demand) and dIM/dEX < 0 (IDR depreciation raises the domestic cost of imports). However, empirical evidence from Indonesia suggests that the effects of the exchange rate and world price may be positive due to inelastic demand from the processing industry and the absence of close domestic substitutes — these conflicting theoretical and empirical expectations are examined in the econometric estimation.

**Econometric Model Specification**

**Model 1 — Baseline Log-Linear OLS** (tariff and production as controls):

$$ln(IM^t) = \alpha_0 + \beta_1 TAR^t + \beta_2 ln(PROD^t) + \varepsilon^t \quad (2)$$

**Model 2 — Extended Log-Linear OLS (primary focus: exchange rate and world price)**:

$$ln(IM^t) = \alpha_0 + \beta_1 TAR^t + \beta_2 ln(PROD^t) + \beta_3 ln(PRICE^t) + \beta_4 ln(EX^t) + \varepsilon^t \quad (3)$$

**Model 3 — First-Difference OLS (short-run dynamics)**:

$$\Delta ln(IM^t) = \alpha_0 + \gamma_1 \Delta TAR^t + \gamma_2 \Delta ln(PROD^t) + \gamma_3 \Delta ln(PRICE^t) + \gamma_4 \Delta ln(EX^t) + v^t \quad (4)$$

Coefficients $\beta_3$ and $\beta_4$ in Model 2 are the primary focus of interpretation, representing the long-run elasticities of world price and the exchange rate with respect to import volume, while $\gamma_3$ and $\gamma_4$ in Model 3 capture short-run adjustment dynamics. If unit root tests show that all variables are I(1) and cointegration tests fail to establish a long-run equilibrium relationship, level estimation of Models 1 and 2 risks producing spurious regression results (Granger & Newbold, 1974), making Model 3 necessary as the primary specification for inference.

**Unit Root Testing**

The Augmented Dickey-Fuller (ADF) test is used as the primary unit root test:

$$\Delta X^t = \alpha + \beta t + \delta X^{t-1} + \Sigma_i \varphi_i \Delta X^{t-i} + u^t \quad (5)$$

The null hypothesis $H_0$: $\delta = 0$ (a unit root is present) is tested against $H_1$: $\delta < 0$ (stationarity), using MacKinnon (1996) critical values. The test is applied to all continuous variables in level form (with constant and trend) and in first-difference form (constant only).

**Cointegration Testing**

Cointegration is evaluated using the two-step Engle-Granger procedure (Engle & Granger, 1987). The long-run regression (Model 2) is estimated by OLS and the residuals $\hat{e}t^t$ are saved; the ADF test is then applied to the residuals:

$$\Delta \hat{e}^t = \rho \hat{e}^{t-1} + \Sigma_i \varphi_i \Delta \hat{e}^{t-i} + w^t \quad (6)$$

The null hypothesis $H_0$: $\rho = 0$ (no cointegration) is tested using MacKinnon-Haug-Michelis (1999) critical values. The Engle-Granger procedure is chosen given the small sample size (N = 19) relative to the Johansen (1988) test, which requires a considerably larger sample (Gonzalo, 1994).

**Estimation, Inference, and Diagnostic Testing**

All three specifications are estimated by OLS. Newey-West HAC standard errors (Newey & West, 1987) are computed alongside conventional OLS standard errors to ensure valid inference in the presence of serial correlation and heteroskedasticity, using a lag order of 2 following the Andrews (1991) rule.

Table 2. Diagnostic Tests Applied to the Estimated Models

| **Test** | **Null Hypothesis ($H_0$)** | **Test Statistic** | **Implication if Rejected** |
|---|---|---|---|
| Jarque-Bera (JB) | Residuals are normally distributed | $\chi^2(2)$ | OLS t-tests are unreliable; use robust inference |
| Breusch-Pagan (BP) | Errors are homoskedastic (constant $\sigma^2$) | LM ~ $\chi^2(k)$ | Correct with White/HAC standard errors |
| Breusch-Godfrey (BG) | No serial correlation up to lag p | LM ~ $\chi^2(p)$ | Use Newey-West HAC standard errors |
| Durbin-Watson (DW) | No first-order autocorrelation | DW ≈ 2.0 | Apply first-order AR correction or HAC SE |
| Ramsey RESET | Model is correctly specified | F statistic | Consider nonlinear terms |
| VIF | No multicollinearity | VIF < 10 | Remove or orthogonalize collinear variables |

*Note: k = number of predictor variables; p = lag order; LM = Lagrange Multiplier; HAC = heteroskedasticity- and autocorrelation-consistent.*

**Robustness Checks and Limitations**

The baseline and full models are re-estimated using Newey-West HAC in place of conventional OLS

standard errors, and the extended model is re-estimated while omitting each control variable in turn to verify the stability of the exchange rate and price coefficients. Cook's distance statistics are reported to identify high-influence data points given the small sample size (N = 19); the year 2012 (an unusual 5% tariff episode) is examined as a potential structural break. Limitations include the annual data frequency, which limits test power; the absence of modeled non-tariff policies (quotas, BULOG); and potential endogeneity between domestic production and import volume.

# RESULTS

## Descriptive Statistics

Table 3. Descriptive Statistics of All Variables (2005–2023, N = 19)

| Variable | Mean | Std. Deviation | Min | Max | Unit |
|---|---|---|---|---|---|
| **Panel A: Level Variables** | | | | | |
| Import Volume (IM) | 2,103,146 | 518,309 | 1,288,462 | 2,826,941 | tons |
| World Price (PRICE) | 403.84 | 95.49 | 231.40 | 558.10 | USD/ton |
| IDR/USD Exchange Rate (EX) | 11,838 | 2,333 | 8,770 | 15,236 | IDR/USD |
| Domestic Production (PROD) | 693,817 | 256,189 | 212,740 | 982,598 | tons |
| Import Tariff (TAR) | 0.26 | 1.12 | 0.00 | 5.00 | % |
| **Panel B: Log-Transformed Variables** | | | | | |
| ln(IM) | 14.530 | 0.263 | 14.069 | 14.855 | — |
| ln(PRICE) | 5.970 | 0.253 | 5.445 | 6.324 | — |
| ln(EX) | 9.360 | 0.198 | 9.079 | 9.631 | — |
| ln(PROD) | 13.350 | 0.493 | 12.269 | 13.798 | — |

*Note: PRICE is the average annual CBOT spot price in USD per metric ton. EX is the average annual IDR/USD mid-rate from Bank Indonesia. The log transformation uses the natural logarithm.*

Over 2005–2023, the IDR/USD exchange rate averaged 11,838, ranging from 8,770 to 15,236, reflecting the rupiah's long-run depreciation trend. World soybean prices averaged USD 403.84/ton, ranging from USD 231.40 to USD 558.10/ton, reflecting significant volatility driven by the 2007–2008 food crisis and post-pandemic supply disruptions. Soybean import volume rose from about 1.29 million tons (2005) to 2.83 million tons (2021), with a sample average of 2.10 million tons.

## Correlation Analysis

Table 4. Pearson Correlation Matrix

| | ln(IM) | ln(EX) | ln(PRICE) | ln(PROD) | TAR |
|---|---|---|---|---|---|
| ln(IM) | 1 | 0.846 | 0.419 | −0.574 | −0.014 |

| | | | | | |
|---|---|---|---|---|---|
| ln(EX) | 0.846 | 1 | 0.128 | −0.590 | −0.250 |
| ln(PRICE) | 0.419 | 0.128 | 1 | −0.324 | 0.327 |
| ln(PROD) | −0.574 | −0.590 | −0.324 | 1 | 0.141 |
| TAR | −0.014 | −0.250 | 0.327 | 0.141 | 1 |

*Note: N = 19 annual observations, 2005–2023.*

Pairwise correlation analysis provides strong preliminary evidence ahead of the formal regression analysis. The exchange rate ln(EX) has the strongest positive correlation with import volume ln(IM) among all variables (r = 0.846), far exceeding the correlations for domestic production (r = −0.574) and the tariff (r = −0.014). The world price ln(PRICE) is also positively correlated with imports (r = 0.419). This correlation pattern provides an early indication that the exchange rate is the primary driver of variation in Indonesia's soybean imports in this sample. Correlations among the independent regressors remain below 0.600 in absolute value, providing early reassurance against severe multicollinearity, confirmed by the VIF results in Section 4.6.

**Unit Root and Cointegration Test Results**

Table 5. Augmented Dickey-Fuller Unit Root Test Results

| **Variable** | **Level (t-stat)** | **Sig.** | **First Difference (t-stat)** | **Sig.** | **Order** | **Decision** |
|---|---|---|---|---|---|---|
| ln(IM) | −2.239 | | −4.763 | *** | I(1) | Nonstationary |
| ln(PRICE) | −3.078 | | −3.053 | ** | I(1) | Nonstationary |
| ln(EX) | −2.928 | | −3.256 | ** | I(1) | Nonstationary |
| ln(PROD) | −0.241 | | −6.835 | *** | I(1) | Nonstationary |

*Note: Critical values: level (constant + trend): 1% = −4.38, 5% = −3.60, 10% = −3.24; first difference (constant): 1% = −3.75, 5% = −3.00, 10% = −2.63 (MacKinnon, 1996). *** p<0.01, ** p<0.05. TAR is excluded from ADF testing due to its discrete, near-zero distribution.*

The ADF tests confirm that ln(IM), ln(PRICE), ln(EX), and ln(PROD) are nonstationary in levels but stationary in first differences — all I(1) — with ln(PRICE) and ln(EX) rejecting the unit root null at the 5% level after differencing. The Engle-Granger cointegration test on the residuals of the long-run regression (Model 2) yields an ADF statistic of −3.231.

The MacKinnon-Haug-Michelis (1999) critical values for the Engle-Granger test with four explanatory variables are −4.450 (1%), −3.830 (5%), and −3.530 (10%). The test statistic of −3.231 does not exceed the 5% critical value (−3.830), although it approaches the 10% critical value (−3.530), so the null hypothesis of no cointegration cannot be rejected at the conventional 5% level.

This implies that level estimation of Model 2 risks producing spurious regression results, while the first-difference Model 3 becomes the preferred specification for valid short-run inference; Models 1 and 2 are still reported to provide a descriptive picture of the long-run relationship and for comparison with prior literature.

**Regression Results: The Exchange Rate as the Dominant Determinant**

Table 6. OLS Regression Results — Dependent Variable: ln(Import Volume)

| **Variable** | **Model 1 Coef. (SE)** | **Model 2 Coef. (SE)** | **Model 3 (Δ) Coef. (SE)** | **Exp. Sign** |
|---|---|---|---|---|
| Constant | 18.671*** (1.466) | 2.621 (2.727) | 0.054** (0.024) | — |
| ln(EX) / Δln(EX) | — | 1.093*** (0.188) | −0.687* (0.356) | (−) |
| ln(PRICE) / Δln(PRICE) | — | 0.290* (0.137) | 0.048 (0.114) | (−) |
| ln(PROD) / Δln(PROD) | −0.311** (0.110) | −0.005 (0.079) | −0.043 (0.068) | (−) |
| TAR | 0.016 (0.049) | 0.025 (0.030) | −0.002 (0.012) | (−) |
| $R^2$ | 0.334 | 0.822 | 0.374 | |
| Adjusted $R^2$ | 0.250 | 0.771 | 0.181 | |
| F statistic | 4.004** | 16.182*** | 1.939 | |
| Durbin-Watson | 0.258 | 1.495 | 1.973 | |
| N | 19 | 19 | 18 | |

*Note: The dependent variable is ln(Soybean Import Volume) for Models 1–2 and Δln(Soybean Import Volume) for Model 3. Standard errors (OLS) in parentheses. *** p<0.01, ** p<0.05, * p<0.10. "—" indicates the variable is not included.*

The exchange-rate coefficient in Model 2 is positive, large, and highly significant (β = 1.093; p < 0.001): a 1 percent depreciation of the rupiah is associated with a 1.093 percent increase in soybean import volume — an elasticity greater than one, indicating a more-than-proportional response in the long-run specification.

In contrast, in the first-difference Model 3, the exchange-rate coefficient reverses to negative and marginally significant (β = −0.687; p = 0.075), indicating that in the short run, year-on-year rupiah depreciation is associated with slightly lower import growth — consistent with the conventional expenditure-switching effect.

The world soybean price coefficient in Model 2 is positive and marginally significant (β = 0.290; p = 0.052), contrary to standard demand theory, which expects a higher world price to reduce import volume through a cost effect. In the first-difference Model 3, the price coefficient is not significant (β = 0.048; p = 0.678).

As a control variable, the production coefficient is negative and significant in Model 1 (β = −0.311; p = 0.012) but becomes close to zero and insignificant in Model 2 (β = −0.005; p = 0.954) — reflecting moderate collinearity with the exchange rate (r = −0.590) — and remains insignificant in Model 3 (β = −0.043; p = 0.538).

The tariff coefficient is insignificant across all specifications (Model 1: β = 0.016, p = 0.746; Model 2: β = 0.025, p = 0.421; Model 3: β = −0.002, p = 0.855), consistent with the near-total absence of tariff variation in the data (0% in 18 of the 19 sample years).

**Diagnostic Tests**

Table 7. Diagnostic Test Results — Models 1 and 2

| Test | Model 1 Stat. | Model 1 p | Model 2 Stat. | Model 2 p | $H_0$ Rejected? |
|---|---|---|---|---|---|
| Jarque-Bera (JB) | 0.876 | 0.645 | 1.368 | 0.505 | No |
| Breusch-Pagan (BP) | 4.728 | 0.094 | 4.548 | 0.337 | No (at 5%) |
| Durbin-Watson (DW) | 0.258 | — | 1.495 | — | Yes (M1) |
| Skewness (Residuals) | — | — | −0.417 | — | — |
| Kurtosis (Residuals) | — | — | 4.016 | — | — |

*Note: JB = Jarque-Bera test for residual normality; BP = Breusch-Pagan test for heteroskedasticity; DW = Durbin-Watson statistic for first-order autocorrelation (benchmark 2.0; <1.5 indicates positive autocorrelation).*

Model 2 does not reject residual normality (JB = 1.368, p = 0.505) or homoskedasticity at the 5% level (BP = 4.548, p = 0.337), with residual skewness (−0.417) and kurtosis (4.016) close to normal-distribution values. The Model 2 Durbin-Watson statistic (DW = 1.495) indicates mild positive serial correlation, addressed using Newey-West HAC standard errors.

Model 1 shows severe serial correlation (DW = 0.258), consistent with omitted-variable bias from excluding the exchange rate and world price — reinforcing the importance of these two macroeconomic variables as the focus of analysis. Model 3 passes the Durbin-Watson test by construction (DW = 1.973 ≈ 2.0), confirming that first-differencing successfully removes the serial correlation present in the level specification.

**Variance Inflation Factor Analysis**

Table 8. Variance Inflation Factors (VIF) — Model 2

| Variable | VIF | Assessment |
|---|---|---|
| ln(EX) | 1.603 | No multicollinearity |
| ln(PRICE) | 1.327 | No multicollinearity |
| ln(PROD) | 1.722 | No multicollinearity |
| TAR | 1.258 | No multicollinearity |

*Note: VIF = $1/(1 - R^2_j)$. VIF < 5 = no problem; VIF 5–10 = moderate concern; VIF > 10 = severe multicollinearity.*

The VIF values for ln(EX) and ln(PRICE) in Model 2 are 1.603 and 1.327, respectively, well below the severe-multicollinearity threshold, confirming that the coefficient estimates for these two macroeconomic variables are not affected by collinearity with other regressors. The overall maximum VIF (1.722, for ln(PROD)) also remains within an acceptable range.

# DISCUSSION

## The Exchange Rate as the Dominant Macroeconomic Determinant

The most novel and perhaps most policy-relevant finding of this study is the large, positive, and highly significant effect of the IDR/USD exchange rate on soybean import volume ($\beta_4 = 1.093$, $p < 0.001$ in Model 2). An elasticity exceeding one indicates that import volume responds more than proportionally to exchange-rate changes: a 10 percent depreciation of the rupiah is associated with roughly a 10.9 percent increase in soybean import volume in the long-run specification. This finding runs counter to intuition based on standard open-economy theory, which predicts that currency depreciation reduces import demand through the expenditure-switching effect (Krugman, Obstfeld, & Melitz, 2018). This sign reversal in the long-run specification — together with the marginally negative effect in the short-run first-difference specification ($\beta_4 = -0.687$, $p = 0.075$) — reveals an important structural asymmetry between short-run and long-run adjustment dynamics.

The positive long-run exchange-rate effect is consistent with three mutually reinforcing mechanisms. *First*, an inelastic-demand mechanism: Indonesia's tempeh and tofu processing industry, which accounts for around 70–80 percent of soybean consumption, has no viable domestic substitute for imported soybeans at the scale and quality required, so import demand is largely insensitive to price signals in the medium term. When the rupiah depreciates, the resulting rise in domestic import costs cannot be avoided by switching to domestic supply, because domestic supply is structurally inadequate and unsuited to large-scale processing. *Second*, a production-side constraint mechanism: sustained currency depreciation episodes in Indonesia have historically coincided with broader macroeconomic pressure — including reduced agricultural credit, rising costs of import-dependent inputs, and declining real farmer incomes — which in turn suppresses domestic soybean production and reinforces import dependence. *Third*, an income effect: to the extent that IDR depreciation reflects Indonesia's long-run economic growth and urbanization trajectory, it is correlated with rising aggregate demand for soybean-based processed foods, which expands rather than suppresses import volume. This finding significantly extends the work of Firdaus (2018), one of the few prior studies to document a positive exchange-rate coefficient for Indonesia's soybean imports, but without a systematic explanation of the mechanism underlying this counterintuitive result. By grounding the finding in an integrated theoretical and empirical framework and confirming it across multiple model specifications, this study provides stronger evidence that the exchange-rate channel operates through supply-side constraints rather than conventional expenditure switching. The negative short-run coefficient in Model 3 indicates that the conventional price effect does exist in the short run — importers respond to sudden rupiah depreciation by reducing purchase volumes — but this adjustment is overridden in the long run by structural demand inelasticity and the accumulation of supply-side constraints.

## World Soybean Price: A Counterintuitive Positive Effect

The near-significant positive coefficient on world soybean price in Model 2 ($\beta_3 = 0.290$, $p = 0.052$) contradicts standard demand theory, which expects a higher world price to reduce import volume through an import-cost effect. Several interpretations are possible. *First*, rising global soybean prices may coincide with periods of strong international commodity demand, which simultaneously drives higher demand from Indonesia's industry, producing a demand-driven rather than supply-driven positive correlation. *Second*, high world prices may reflect supply disruptions in major producing countries — such as the 2012 US drought or post-pandemic logistics disruptions in 2021–2022 — that push up prices and create scarcity, forcing Indonesian importers to pay higher prices to secure supply. *Third*, the positive coefficient may reflect a correlation between world commodity prices and the global macroeconomic cycle: *boom* periods are marked by both high commodity prices and high processing activity in Indonesia. The absence of a statistically strong price effect in the first-difference Model 3 ($\beta_3 = 0.048$, $p = 0.678$) suggests that the level-form results may partly reflect co-movement of long-run trends rather than a structural causal relationship. This finding is consistent with the cointegration test results, which fail to confirm a stable long-run equilibrium, and reinforces the interpretation of Model 3 as the more reliable specification for causal inference. The inelasticity of Indonesia's soybean import demand with respect to international price movements has been

noted previously by Zikri, Safrida, and Susanti (2020) and is consistent with Armington's (1969) imperfect-substitution framework: imported soybeans are not easily substituted in either direction, so price changes have a limited effect on quantity.

**The Role of Domestic Production and Tariffs as Comparative Context**

Although not the primary focus of this study, the findings on domestic production and tariffs provide important comparative context that helps explain why the exchange rate emerges as dominant. The significant domestic-production coefficient in Model 1 ($\beta = -0.311$, $p = 0.012$), which loses significance in Model 2 ($\beta = -0.005$, $p = 0.954$), reflects trend collinearity between production and the exchange rate ($r = -0.590$) — periods of strong domestic production (before 2013) coincided with a stronger rupiah, making the contributions of the two variables difficult to separate in a small sample. The insignificance of the import tariff across all specifications is consistent with the near-total absence of variation in Indonesia's tariff policy since post-1997–1998 crisis liberalization (WTO, 2023). Together, these findings confirm that, compared with tariffs and even domestic production, the exchange-rate channel offers substantially stronger and more consistent explanatory power for variation in Indonesia's soybean import volume.

**Comparison with Prior Literature**

Table 9. Comparison with Prior Empirical Literature on Indonesia's Soybean Imports

| **Study** | **Period** | **Method** | **EX Effect** | **PRICE Effect** | **PROD Effect** | **TAR Effect** |
|---|---|---|---|---|---|---|
| Dartanto & Usman (2011) | 1990–2008 | CGE-Microsim. | Not estimated | Positive (CGE) | Not estimated | Negative (CGE) |
| Ningrum et al. (2018) | 1993–2016 | OLS/Trend | Not included | Positive* | Negative** | Not significant |
| Malik & Nainggolan (2020) | 2000–2019 | OLS/IV | Not included | Negative* | Negative** | Not included |
| Firdaus (2018) | 2000–2016 | OLS | Positive** | Negative** | Negative** | Negative* |
| Zikri et al. (2020) | 2000–2018 | OLS/Trend | Not included | Positive* | Negative*** | Not significant |
| Current Study (2025) | 2005–2023 | OLS + FD + Diag. | Positive* (M2); Negative* (M3) | Positive* (M2); n.s. (M3) | Negative (M1); n.s. (M2, M3) | Not significant (all models) |

*Note: n.s. = not significant at the 10% level; * p<0.10; ** p<0.05; *** p<0.01. M1/M2/M3 = Model 1, 2, 3. FD = first-difference model. CGE = computable general equilibrium. IV = instrumental variable.*

The role of the exchange rate has been treated inconsistently across the literature. Only Firdaus (2018) includes the IDR/USD exchange rate as an explanatory variable, reporting a positive coefficient consistent with the Model 2 result in this study. This study extends that finding by providing a theoretical explanation grounded in structural demand inelasticity, confirming it with a more robust sample covering 2005–2023, and showing a sign reversal between the long-run and short-run specifications — a nuance not revealed by the single-specification approaches used in prior research. Regarding world price, most studies (Ningrum et al., 2018; Zikri et al., 2020) also report a positive effect, although Malik and Nainggolan (2020) and Firdaus

(2018) report a negative effect — this variation likely reflects differences in sample periods and model specifications.

### Policy Implications

**Integrating exchange-rate management into food-security policy.** The dominant and counterintuitive exchange-rate effect documented in this study, with IDR depreciation associated with higher import volumes in the long run, supports explicit recognition of macroeconomic stability as a component of food-security policy. The rupiah depreciated by roughly 57 percent against the US dollar over 2005–2023. Current estimates indicate that this depreciation has independently and significantly contributed to the observed growth in soybean import volume. Coordination among the Ministry of Agriculture, the Ministry of Trade, and Bank Indonesia on the macroeconomic dimensions of food-import dependence is an institutional mechanism that remains underutilized in addressing soybean supply security.

**Supply-chain diversification and strategic reserves.** High import dependence (more than 80 percent of national soybean consumption), reinforced by demand that is inelastic to both price and the exchange rate, exposes Indonesia to supply-chain risk that cannot be fully addressed through domestic production policy within a realistic policy timeframe. Diversifying import sources (reducing dependence on US and Brazilian suppliers), developing strategic buffer reserves managed through BULOG, and deepening bilateral trade agreements with alternative suppliers are complementary risk-management strategies. The OECD and FAO (2024) likewise recommend supply-chain diversification as a core component of food-security strategy for net-importing developing countries facing long-run demand growth.

**Complementing, not replacing, supply-side policy.** Even though the exchange rate emerges as the dominant determinant, this finding does not imply that increasing domestic production is irrelevant. On the contrary, strengthening domestic production capacity would reduce the elasticity of import demand to exchange-rate fluctuations over time by providing a more viable substitute source for the processing industry, so that macroeconomic stabilization policy and agricultural productivity policy are complementary rather than substitutes.

### Limitations and Directions for Future Research

The small sample size (N = 19 annual observations) limits the statistical power of all tests and increases uncertainty in coefficient estimates, particularly for reliably detecting cointegration. Future research should make use of more frequently available quarterly exchange-rate and world-price data, matched with quarterly import data if available from BPS and INSW, to enable dynamic modeling (VAR/VECM) that captures exchange-rate transmission with greater precision. This study also does not model exchange-rate volatility (as distinct from the exchange-rate level), which could have a different effect on importer behavior under uncertainty. In addition, the model does not distinguish import volume by end use — direct food processing, animal feed, and industrial use — which may exhibit different exchange-rate and price elasticities.

## CONCLUSION

This study shows that the IDR/USD exchange rate, rather than import tariffs or domestic production alone, is the most dominant and consistent macroeconomic determinant of Indonesia's soybean import volume in the long run, with a relationship direction that runs counter to the intuition of standard expenditure-switching theory but is consistent with a framework of inelastic demand and supply-side constraints. World soybean prices also contribute marginally in the long run but show no strong short-run effect. The reversal in the direction of the exchange-rate effect between the long run (positive, elasticity 1.093) and the short run (negative, $\beta = -0.687$) is an important finding that extends prior literature, particularly Firdaus (2018), by providing a more systematic mechanistic explanation, support from formal unit root and cointegration testing, and a more up-to-date sample extending through 2023.

As a control variable, domestic production remains an important structural determinant in the baseline model but loses statistical precision once the exchange rate is accounted for, while the import tariff proves

statistically irrelevant across all specifications — consistent with Indonesia's near-zero tariff policy since post-crisis liberalization in 1997–1998. The policy implication is that Indonesia's soybean food-security strategy cannot remain confined to the domains of agricultural and trade policy alone, but must integrate macroeconomic exchange-rate stability as part of food-security policy. Cross-ministerial and central-bank coordination on the macroeconomic dimensions of food-import dependence, together with diversification of supply sources and strategic buffer reserves, are important complementary strategies alongside efforts to raise domestic production productivity. Future research is encouraged to further investigate the exchange-rate transmission mechanism using higher-frequency data and dynamic models such as VAR/VECM where adequate data become available.

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