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Estimation of the Impact of Devaluation on Indonesian Aggregate Trade Performance

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Abstract

This study evaluates the relationship between the real exchange rate and Indonesia's aggregate trade performance between 1975 and 1990. I estimate structural import and export functions that are consistent with traditional empirical trade models, in which a rupiah devaluation leads rupiah import and export prices to increase relative to domestic prices. The study provides the first Indonesian import demand elasticity estimates, and it offers more precisely estimated non-oil export supply elasticities than previous work. Results add to the body of evidence that supports devaluation's use in dampening imports and stimulating exports.

I. Introduction

As an oil exporter, Indonesia gained large windfall oil earnings after the oil price quadrupled in 1973. The price hike led Indonesia to rely on oil and gas exports for at least 65% of total export earnings until the early 1980s to

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finance debt repayments and capital imports. However, declining oil revenues after 1982 and rising debt service costs required diversification of the oil-based economy towards non-oil exports and increased domestic savings. Strong non-oil export growth and increased foreign aid during the 1980s helped Indonesia to cushion these international shocks and maintain on-time foreign debt repayments.[1] In 1986 when the oil price bottomed, non-oil exports surpassed the oil sector as the primary contributor to total export earnings.

Indonesia's international trade patterns raise several interesting questions. In particular, what role did exchange rate devaluation play in the rapid growth of non-oil exports, and did devaluation help to curb import growth? In 1978, Indonesia's government began to use active exchange rate management to encourage domestic production of tradeable goods. Further, the oil price crash after 1982 led the Indonesian government to promote a structural overhaul of its microeconomic trade, industrial, and financial sector policies, while it continued sensible macroeconomic policies from earlier years to improve the trade and investment environment. The policy reforms contributed to relatively strong income growth, which helped fuel import demand. While annual real GDP growth in other OPEC members and heavily indebted countries averaged less than 2% in the 1980s, Indonesia averaged 5.5% real GDP growth ⟨Rodgers [1993, p. 1]⟩.

This study evaluates the relationship between the real exchange rate and Indonesia's aggregate trade performance between 1975 and 1990. I estimate structural import and export functions that are consistent with traditional empirical trade models, in which a rupiah devaluation leads rupiah import and export prices to increase relative to domestic prices. I exclude oil and gas exports in order to highlight the determinants of Indonesia's phenomenal non-oil export growth. This study provides the first Indonesian *import demand* elasticity estimates in the empirical trade literature; my results fall within the range of estimates for other countries with similar characteristics. Through more accurate measurement of the aggregate non-oil export

1. Negative net capital flows during the oil boom years became increasingly more positive during the 1980s. Net capital inflows during the late 1980s averaged approximately $4 billion a year, or 5% of GDP.

price and quantity, and not just a larger sample, this study also offers more precisely estimated non-oil *export supply* elasticities than Arize [1990], the only previous empirical study of Indonesia's aggregate exports.

My results indicate a large and statistically significant relationship between the real exchange rate and aggregate trade. With a 1.0% increase in the relative rupiah price of imports, total import demand falls between 1.2% and 1.7% depending on the specification. With a 1.0% rise in the relative rupiah price of non-oil exports, non-oil export supply increases between 0.7% and 1.3%. Tests for structural change in the elasticity estimates indicate that regardless of estimation in levels or first differences, import demand's responsiveness to prices becomes smaller in magnitude during the 1986-1990 "steady reform" period. A strong feedback effect from non-oil exporters' needs for imported inputs and machinery, particularly in textiles and garments, may have lessened devaluation's ability to control import growth. However, stable non-oil export supply coefficient estimates provide a useful rule of thumb to forecast devaluation's impact on exports.

II. Data Description and Previous Work

This study uses an extensive data set of Indonesia's imports and exports, by sector and quarter from 1975 to 1990, which I constructed from data provided by the Indonesian Bureau of Statistics (BPS).[2] The original data (coded from Indonesian customs forms) cover individual import and export transactions from 1975 to 1990. The aggregation and classification process resulted in quarterly and annual trade data, sectoral and aggregate, by value and metric tons. This unique data set has several advantages over published trade data sources, since it offers more detailed and higher frequency information.[3] The Appendix records all data sources and variables.[4]

Panel A of Table 1 highlights three empirical studies of small country

2. Most data processing and aggregation took place at the Harvard Institute for International Development project on Customs and Economic Management in Jakarta, Indonesia. I worked under the supervision of Jeffrey Lewis and Michael Roemer.
3. See the International Monetary Fund, International Financial Statistics, and the United Nations, International Trade Statistics Yearbook. The IMF publishes only aggregate, oil, and rubber trade figures. The UN publishes annual aggregate values, sec-

Table 1
Comparison of Trade Elasticities with Previous Estimates

Panel A: Import Demand				
Study	Country	Sample	Long-Run Price Elasticity	Long-Run Income Elasticity
Gafar	Trinidad & Tobago	1967-1984	-0.53^{*} (0.29)	3.01^{***} (0.28)
Salehi-Isfahani	Nigeria	1963-1982	-1.01^{***} to -1.24^{***} (0.14) (0.22)	0.36 to 0.69^{***} (0.22) (0.16)
Tegene	African Sample	1973-1985	-0.13^{**} to -2.74^{***} (0.06) (0.71)	-0.17 to 2.61^{***} (0.63) (0.84)
Rodgers	Indonesia	1975-1990	-1.69^{***} (0.27)	2.37^{***} (0.45)

Panel B: Export Supply						
Study	Country	Sample	Short-Run Price Elasticity	Long-Run Price Elasticity	Short-Run Income Elasticity	Long-Run Income Elasticity
Arize	Indonesia	1973-1985	0.57^{*} (0.31)	2.15	1.06^{***} (0.27)	4.02
Rodgers	Indonesia	1975-1985	0.38^{***} (0.11)	0.94^{**} (0.35)	0.92^{***} (0.24)	2.27^{**} (0.97)
Rodgers	Indonesia	1975-1990	0.40^{***} (0.10)	1.31^{**} (0.53)	0.43^{**} (0.16)	1.40 (0.86)

Notes: Significance levels are *** = 1%, ** = 5%, * = 10% (two tail tests). The parentheses contain standard errors. Arize [1990] does not provide standard errors for long-run estimates. The Rodgers export sample begins with 1975 and cannot completely replicate the Arize sample period. All results are from regressions in levels.

toral values, and some sectoral weights.

4. I interpolate some annual figures into quarters following Goldstein and Khan [1976]. The choice of domestic price deflator follows suggestions in Devarajan, Lewis, and Robinson [1991].

importers with similar characteristics to Indonesia, that serve as useful bench-marks against which to compare this study's estimates.[5] Precisely estimated long-run price elasticities range from –0.13 (0.06) to –2.74 (0.71). These studies each suggest that import demand responds to price incentives to some extent. Like the others, I assume that the import price is exogenously determined on the world market. Indonesia's low shares in world imports across sectors support the small country assumption.

In the only previous empirical export study to examine Indonesia, Arize [1990] estimates aggregate export supply and demand determinants from 1973 to 1985.[6] Panel B reports Arize's estimates. Arize does not present standard errors for the long-run estimates or sufficient information with which to calculate them, weakening any policy conclusions. The present study does present long-run standard errors and makes two major changes.

First, to avoid the problems associated with Arize's aggregate export unit value and weight variables, I construct Divisia weighted non-oil export price and quantity indices.[7] I use IMF data on world commodity prices to measure Indonesia's commodity export prices; UN data on developing country manufactured exports to measure Indonesia's aggregate manufactured export price; and BPS data on each category's value share in Indonesia's non-oil exports.[8] Hence the Indonesian Divisia export price variable is based on actual price data, and quantity is derived as the residual from total values and prices. The procedure accounts for structural change in non-oil export composition and avoids the high volatility in unit values.

Second, with the Divisia constructions I effectively assume Indonesia is a

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5. For studies not cited elsewhere that estimate LDC aggregate trade behavior, see Khan [1974], Bond [1985], and Marquez and McNeilly [1988].
6. Moran [1988] estimates structural supply and demand equations for manufactured exports, from 1965 to 1983, of a pooled developing country sample that includes Indonesia. The study does not pinpoint Indonesian elasticities.
7. See Jorgenson and Griliches [1971]. The Divisia method is not feasible for imports because over three quarters of Indonesia's imports consist of capital goods and manufactured intermediates, for which world prices are not readily available.
8. The International Monetary Fund, International Financial Statistics, has quarterly world price indices of Indonesia's commodity exports. The United Nations International Trade Statistics Yearbook 1989, Special Table C, has data on the index of unit values in US$ for manufactured goods exports from developing countries.

price taker for its aggregate non-oil exports, while Arize uses 2SLS to estimate supply and demand equations.[9] Indonesia may have some market power for plywood and rubber; however, these two commodities together only account for an average 25% share in Indonesian non-oil exports from 1975 to 1990. This lends credibility to the assumption that in aggregate, Indonesia takes the world price of its non-oil exports as given. Also, Bound, Jaeger, and Baker [1993] indicate that instrumental variable estimates can have large inconsistencies if the instruments are only weakly correlated with the endogenous explanatory variable. I performed similar tests in the first stage regression and found low R^2 and F statistics.[10] Because the instruments (trading partner and competitor export prices and income) are only weakly correlated with the endogenous variable (the dollar export price), export supply elasticity estimates using demand side instruments may be inconsistent. Assuming Indonesia takes the world price as given avoids the problem. An interesting question is whether calculation of long-run standard errors, change in price and quantity measures, and use of OLS will yield precisely estimated long-run elasticities.

III. Methodological Approach

The structural import demand function is consistent with traditional empirical trade models.[11] I assume that Indonesian producers and state owned enterprises, rather than consumers, purchase imports, and they maximize profits. Their import demand responds to the rupiah import price relative to domestic prices, and to income. Desired aggregate import demand follows:

$$\log M_t^d = \alpha_0 + \alpha_1 \log(PM_t * ER_t / PD_t) + \alpha_2 \log Y_t + \varepsilon_t.$$

9. Estimation with 2SLS yielded similar unprecisely estimated export demand coefficients as Arize [1990].
10. Also, Arize constructs one demand side instrument, the world price, as a weighted index of competitor country and trading partner aggregate export prices. However, this construction includes many products in which Indonesia does not trade.
11. Goldstein and Khan [1985] provides an extensive review of supply and demand specifications and estimation issues.

The notation M_t^d denotes quantity demand of Indonesian total imports in period t; PM_t denotes the dollar price of imports; ER_t denotes the nominal exchange rate (rupiah/US$); PD_t measures domestic prices in rupiah; and Y_t denotes real domestic GDP. The relative demand price measures aggregate profitability of importing compared to purchasing domestically produced goods. A rupiah devaluation increases the rupiah cost of imports relative to domestic prices. I do not control for the impact of devaluation on domestic prices. Import demand should have a negative price elasticity, α_1, and a positive income elasticity, α_2. The disturbance term's distribution has the usual i.i.d. assumptions.

The export supply function has a similar construction. I assume that Indonesian exporters produce for the world and domestic markets, and they maximize profits. Their export supply responds to the rupiah price of exports relative to domestic prices, and to domestic income. Desired aggregate export supply follows:

$$\log X_t^s = \beta_0 + \beta_1 \log(PX_t * ER_t / PD_t) + \beta_2 \log Y_t + v_t. \quad (2)$$

The notation X_t^s denotes quantity supply of Indonesian non-oil exports in period t; PX_t denotes the dollar price of non-oil exports; ER_t denotes the nominal exchange rate (rupiah/US$); PD_t measures domestic prices in rupiah; and Y_t denotes real domestic income. The relative supply price measures profitability of producing for the world market. A rupiah devaluation increases the rupiah export price relative to domestic prices. Export supply should have a positive price elasticity, β_1, and a positive income elasticity, β_2.

The actual import and export levels may respond with a lag to desired demand and supply due to transactions costs, contracts, and delivery delays.[12] Also, use of quarterly data to test the model makes it more likely that adjustment within the period will not hold. Using a partial adjustment approach, the short-run import demand and export supply equations become:

$$\log M_t = \gamma\alpha_0 + \gamma\alpha_1 \log(PM_t * ER_t / PD_t) + \gamma\alpha_2 \log Y_t + (1-\gamma)\log M_{t-1} + \gamma\varepsilon_t. \quad (3)$$

$$\log X_t = \lambda\beta_0 + \lambda\beta_1 \log(PX_t * ER_t / PD_t) + \lambda\beta_2 \log Y_t + (1-\lambda)\log X_{t-1} + \lambda v_t. \quad (4)$$

12. The treatment of lag structures varies widely. See Wilson and Takacs [1979], Thursby and Thursby [1984], Bahmani-Oskooee [1986], and O'Neill and Ross [1991].

To construct the structural coefficients α_1, α_2, β_1, and β_2, I divide the estimated coefficients by γ and λ, calculated from the coefficient on lagged imports and exports, respectively. I calculate standard errors following the method in Kendall and Stuart [1977].[13]

I add several other variables. In the import equation, the stock of foreign exchange reserves lagged one period proxies for the level of tariffs and non-tariff barriers, for which direct data are not available. Reserves capture the government's likelihood to impose trade restrictions, and the lag allows for the time required for trade policy to formulate. The foreign exchange variable is expected to have a positive coefficient, assuming the government restricts imports when foreign exchange levels are low. The export equation includes the oil price relative to the price of non-oil exports, which is expected to have a negative coefficient. Dutch Disease and government policies tended to push non-oil exports in the opposite direction of the relative oil price.

To control for seasonal effects, I include quarterly dummy variables in all estimation equations. For example, seasonality could arise from observed trade surges recorded in the month of December. Accumulated backlogs in data processing from previous months and pressure from higher level government officials to complete trade statistic figures for the calendar year are possible reasons for higher value entries in December. In addition, I interact prices and income with a dummy variable for the 1986-1990 steady reform period to test for structural change in the import and export elasticities. Finally, I estimate the equations with constrained and unconstrained relative price terms, and test whether the separated price coefficients significantly differ from each other.

IV. Results

Table 2 reports the import demand results from regressions using data in levels.[14] Because the separate coefficient estimates for the rupiah import

13. Kendall and Stuart [1977, p. 247] calculates the variance of a ratio of two random variables as follows:

$$Var\left(\frac{a_1}{a_2}\right)=\left(\frac{E(a_1)}{E(a_2)}\right)^2*\left(\frac{Var(a_1)}{E^2(a_1)}+\frac{Var(a_2)}{E^2(a_2)}-\frac{2Cov(a_1, a_2)}{E(a_1)E(a_2)}\right)$$

14. Results for non-oil imports are similar.

Table 2
Aggregate Import Demand Results, 1975-1990

Panel A: Short-Run Elasticity Estimates		
PM_t*ER_t/PD_t	−1.073*** (0.151)	–
PM_t*ER_t	–	1.085*** (0.146)
PD_t	–	1.551*** (0.348)
Y_t	1.509*** (0.161)	0.680 (0.618)
M_{t-1}	0.364*** (0.075)	0.350*** (0.072)
FX_{t-1}	−0.186*** (0.048)	−0.189*** (0.048)
Constant	1.797*** (0.417)	−1.392 (1.382)
Sample Size	63	63
Adj R^2	0.830	0.832
F	44.283	39.469
Durbin h	1.263	1.174
Panel B: Long-Run Price and Income Elasticity Calculations		
PM_t*ER_t/PD_t	−1.687*** (0.273)	–
PM_t*ER_t	–	1.670*** (0.271)
PD_t	–	2.389*** (0.627)
Y_t	2.373*** (0.449)	1.047 (0.950)
Coefficient Adjustment	0.636*** (0.075)	0.650*** (0.072)

Panel C: Structural Change in Short-Run Elasticities			
	1975-1985	Interaction Term	1986-1990
PM_t*ER_t/PD_t	−1.179*** (0.170)	0.615** (0.267)	−0.563*** (0.195)
Y_t	1.924*** (0.262)	−0.537 (0.424)	1.387*** (0.342)

Notes: Significance levels: *** = 1%, ** = 5%, * = 10% (two tail tests). The parentheses contain standard errors. I correct for seasonality with quarterly dummies. No equation has a significantly large Durbin h. I calculate heteroskedastic consistent standard errors following White [1980]. The coefficients on PM_t*ER_t and PD_t do not significantly differ from each other in absolute value (t = −1.13). In Panel *C*, price and income slope dummy variables for 1986-1990 are added to the constrained price short-run import demand equations. The equations are all estimated in levels.

price and domestic price do not significantly differ from each other in absolute value, I will limit discussion to the constrained price results. Durbin *h* tests find no evidence of serial correlation ⟨Johnston [1984]⟩. Panel A shows that with a rise of 1.0% in the relative rupiah price of imports, Indonesian producers decrease their purchase of imports by 1.1% in the short run. Short-run import demand is also income elastic. The estimates' magnitudes and precision are robust to alternative measures of economic activity and import restriction proxies.[15] The negative coefficient estimate on foreign exchange reserves, robust to estimation in levels and first differencing, is surprising. One interpretation is that the government actually increased import restrictions as reserves grew from booming oil earnings in the 1970s into the early 1980s, and it loosened import barriers in the mid-1980s as foreign exchange earnings fell in order to provide non-oil exporters with lower input costs.

Panel B reports long-run price and income elasticities. Import demand drops 1.7% with a 1.0% increase in the relative import price after a mean adjustment lag of less than one quarter. The result indicates that Indonesia's import demand has a strong relationship with the real exchange rate. The long-run estimates' magnitudes and precision are robust to using the Almon lag as an alternative lag structure.[16]

Panel C reports test results for structural change in import demand elasticities from 1986 to 1990. Price and income slope dummy variables for 1986-1990 are added to the constrained price short-run import demand equations. Column 1 reports estimated 1975-1985 price and income coefficients; Column 2 reports estimated interaction term coefficients; and Column 3 reports calculated 1986-1990 price and income coefficients. The relative price interaction term has a significant coefficient estimate, suggesting that imports become less responsive to price signals after 1986. A strong feedback effect from non-oil exporters' needs for imported inputs and machinery, particularly in the textile and garments industries, may have reduced

15. Alternatives include quarterly GDP's deviation from trend to measure cyclical income; and debt service/exports, debt service/GDP, and the oil price to proxy for trade restrictions. Also, variations in the number of lags and choice of deflator on the foreign exchange reserve variable do not significantly change the import results.
16. Almon lag estimation results for imports and exports are available upon request.

Table 3
Aggregate Non-Oil Export Supply Results, 1975-1990

Panel A: Short-Run Elasticity Estimates		
PX_t*ER_t/PD_t	0.402*** (0.096)	–
PX_t*ER_t	–	0.390*** (0.115)
PD_t	–	–0.346 (0.282)
Y_t	0.429** (0.162)	0.353 (0.421)
X_{t-1}	0.693*** (0.077)	0.693*** (0.077)
$POIL_t$	–0.112** (0.044)	–0.114*** (0.041)
Constant	–2.015*** (0.561)	–0.008 (0.828)
Sample Size	63	63
Adj R^2	0.977	0.977
F	374.8	322.1
Durbin h	–1.199	–1.214
Panel B: Long-Run Price and Income Elasticity Calculations		
PX_t*ER_t/PD_t	1.309** (0.531)	–
PX_t*ER_t	–	1.271** (0.589)
PD_t	–	–1.127 (1.059)
Y_t	1.396 (0.861)	1.151 (1.558)
Coefficient Adjustment	0.307*** (0.077)	0.307*** (0.077)

Panel C: Structural Change in Short-Run Elasticities			
	1975-1985	Interaction Term	1986-1990
PX_t*ER_t/PD_t	0.456*** (0.101)	–0.098 (0.260)	0.358 (0.241)
Y_t	0.621*** (0.221)	0.273 (0.212)	0.894*** (0.325)

Notes: Significance levels: *** = 1%, ** = 5%, * = 10% (two tail tests). The parentheses contain standard errors. I correct for seasonality with quarterly dummies. No equation has a significantly large Durbin *h*. I calculate heteroskedastic consistent standard errors following White [1980]. The coefficients on PX_t*ER_t and PD_t do not significantly differ from each other in absolute value (t = –0.08). In Panel C, price and income slope dummy variables for 1986-1990 are added to the constrained price short-run export supply equations. The equations are all estimated in levels.

import demand's responsiveness to prices after 1986. The value of imported textile yarns, fabrics, and machinery jumped from 1.6% of total imports before the steady reform period to 9.3% by 1990. Structural change in the import price elasticity is robust to estimation with first differenced data.

Table 3 reports non-oil export supply elasticities from regressions using data in levels. Again, a strong relationship between Indonesian trade and price incentives constitutes a highlight of the table. Because the separate coefficient estimates for the rupiah export price and domestic price do not significantly differ from each other in absolute value, I will limit discussion to the constrained price results. Durbin h tests find no evidence of serial correlation. Panel A shows that with a 1.0% increase in the relative rupiah price of exports, non-oil export supply increases 0.4% in the short run. Export supply is also income inelastic. The estimates' magnitudes and precision are robust to alternative constructions of the Divisia indices and of trade policy proxies.[17] The negative oil price coefficient suggests that non-oil exports fell during the 1970s and early 1980s' oil boom in a Dutch Disease phenomenon, and they rose after 1982 as the oil price collapsed, due to the active government policy to promote non-oil exports.

Panel B shows that long-run export supply moves positively and more than proportionately with relative prices, after a longer mean adjustment lag compared to imports of almost 3 quarters. Long-run export supply rises 1.3% with a 1.0% increase in relative prices. Using the Almon lag as an alternative lag structure yields a trade-off: both the price and income estimates gain precision at the cost of a large decline in overall explanatory power of the regression. Panel C reports test results for structural change in non-oil export supply elasticities from 1986 to 1990. Price and income slope dummy variables for 1986-1990 are added to the constrained price short-run export supply equations. Results indicate that the price and income elasticity estimates do not exhibit structural change.

Finally, Table 4 reports both import and export results from regressions

17. Alternatives include replacing the United Nations measure of the LDC aggregate manufactured export price with World Bank and Hong Kong data. Also, including various measures of the duty drawback facility and the stock of foreign exchange reserves does not change the price results and yields insignificant coefficient estimates on the trade policy proxies.

Table 4
Sensitivity Analysis: Results Based on First Differenced Series

Panel A: Long-Run Aggregate Import Estimates		
$PM_t{}^*ER_t/PD_t$	−1.189*** (0.127)	–
$PM_t{}^*ER_t$	–	−1.183*** (0.130)
PD_t	–	1.717* (0.920)
Y_t	1.294 (1.029)	1.235 (1.060)
FX_{t-1}	−0.062 (0.085)	−0.063 (0.085)
Constant	−0.082*** (0.027)	−0.096** (0.041)
Sample Size	62	62
Adj R^2	0.724	0.720
F	27.615	23.384
DW	3.043	3.070
Panel B: Long-Run Aggregate Non-Oil Export Estimates		
$PX_t{}^*ER_t/PD_t$	0.718*** (0.166)	–
$PX_t{}^*ER_t$	–	0.666*** (0.164)
PD_t	–	0.185 (0.632)
Y_t	0.596 (0.860)	0.509 (0.863)
$POIL_t$	−0.076 (0.083)	−0.088 (0.084)
Constant	−0.080*** (0.022)	−0.101*** (0.027)
Sample Size	63	63
Adj R^2	0.428	0.432
F	8.722	7.729
DW	2.310	2.397

Notes: Significance levels: *** = 1%, ** = 5%, * = 10% (two tail tests). The parentheses contain standard errors. I correct for seasonality with quarterly dummies. I calculate heteroskedastic consistent standard errors following White [1980]. The coefficients on $PM_t{}^*ER_t$ and PD_t, and on $PX_t{}^*ER_t$ and PD_t, do not significantly differ from each other in absolute value ($t = 0.59, 0.67$).

using first differenced series. Tests for unit roots indicate the import price and quantity series are stationary, but the export and domestic variables are integrated of order one. None of the series are cointegrated. Results from first differenced regressions are consistent in sign but smaller in magnitude for both imports and exports. The long-run import price elasticity drops from 1.7 to 1.2, and the export price elasticity falls from 1.3 to 0.7. Hence first differencing places a lower bound on my estimates of the relationship between the real exchange rate and Indonesia's aggregate trade.

V. Conclusion

The study has evaluated Indonesian import and non-oil export performance from 1975 to 1990, and found a large and statistically significant relationship between the real exchange rate and Indonesia's aggregate trade. While estimation with first differenced data yields coefficients that are smaller in magnitude, my conclusions do not change. These results add to the body of evidence that supports devaluation's use in dampening imports and stimulating exports. The Indonesian estimates are consistent with those for OECD countries summarized in Goldstein and Khan [1985]. In addition, Table 1 shows that the import demand results fall within the range of estimates for other countries with similar characteristics. The table also shows my export supply results for 1975-1990 and for 1975-1985, the sample which most closely replicates the Arize sample. While my supply elasticities for both samples are smaller in magnitude than Arize, my short-run price elasticities are more precisely estimated. Arize's exclusion of long-run standard errors prevents comparison of the long-run estimates. My tests for structural change during deregulation find no evidence of significant changes in the non-oil export estimates. Hence this study's stable and precisely estimated price elasticities provide a useful tool to forecast Indonesian non-oil export responses to changes in the exchange rate.

Appendix

Data Sources and Variable Construction[1]

Name	Meaning	Data Units or Construction	Source
PX	Export Price	Divisia price index	BPS, IMF, UN
X	Export Quantity	Divisia quantity index	BPS, IMF, UN
PM	Import Price	Unit value (US$)	BPS
M	Import Quantity	Weight (metric tons)	BPS
Y	Domestic Income	Real quarterly GDP by construction	IMF
PD	Domestic Price	Domestic absorption GDP deflator by construction	WB
ER	Exchange Rate	Nominal exchange rate (rp/$), period average	IMF
FX	Foreign Exchange	Central Bank foreign exchange reserves, minus gold, deflated by the dollar import price	IMF, BPS
POIL	Oil Price	Unit value of Indonesian oil exports, relative to *PX*	BPS

Note: 1: The study uses quarterly data and converts all data into indices with base year=1985.

Sources: BPS = Indonesian Bureau of Statistics, trade data tapes; IMF = International Monetary Fund, International Financial Statistics; UN = United Nations, International Trade Statistics Yearbook; WB = World Bank, World Tables.

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