Journal of Agricultural and Resource Economics 42(3):372–385 ISSN 1068-5502

Quality, Sourcing, and Asymmetric Exchange-Rate Pass-Through into U.S. Coffee Imports

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Few studies to date have investigated the extent of linkages between long-run asymmetries in bilateral trade and fluctuations in real exchange rates and importer demand in non-oil commodity markets. This paper generates estimates of trade elasticities in U.S. raw coffee imports, applying a nonlinear autoregressive distributed lag model and explicitly testing the extent to which nonlinearities matter to U.S. commodity sourcing in the short and long run. Models with asymmetries in both exchange rates and U.S. income point to the critical role that asymmetric pass-through plays in explaining long-run dynamics in U.S. import trade for a major commodity supply chain.

Key words: asymmetric cointegration, coffee trade, import demand, NARDL, real exchange rates, trade elasticities

Introduction

Since the collapse of the fixed exchange-rate regime, the uncertainties of volatile exchange rates and the resulting risks to international trade have attracted the attention of both theoretical and empirical trade economists. While the theoretical literature disagrees about the direction of the effect of exchange-rate volatility on trade flows, a rapidly growing empirical literature has benefited from progressive time-series econometric techniques and more available macro- and micro-level data in the quest for answers (Bahmani-Oskooee and Hegerty, 2007).

Past empirical studies largely neglected the possibility of a nonlinear and/or asymmetric reaction of prices or trade to external shocks (McKenzie, 1999). More recent literature, however, has shown that asymmetries in trade are often structural in nature and thus should be modeled in terms of short- and long-run responses in order to overcome assumptions about the uniform pass-through of macroeconomic shocks. Understanding the dynamics in which currency fluctuations are passed through major commodity supply chains has become of paramount importance to policy-making, trade-risk management, and the welfare of economic agents at large. A case in point is the increasing evidence of asymmetric and nonlinear pass-through of exchange rates and crude oil prices to the price of gasoline (e.g., Kilian, 2008; Atil, Lahiani, and Nguyen, 2014; Bagnai and Mongeau Ospina, 2015). Delatte and López-Villavicencio (2012) and Fedoseeva and Werner (2016) investigate exchange-rate pass-through in the context of exporter pricing-to-market behavior, while Verheyen (2013) studies European-U.S. export trade at large. Fedoseeva and Zeidan (2016) analyze European industry-level exports to the BRIC countries and Fedoseeva (2014, 2016) conducts the only published studies to test for long-run asymmetries and nonlinearities in agricultural commodity trade. Independent of their specific focuses, these studies conclude that the assumptions of symmetry and linearity in trade elasticities are likely to lead to biased coefficient estimates and misleading policy implications unable to represent the true underlying trade behaviors.

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Review coordinated by Tian Xia.

Yet in spite of its strong relevance for trade in non-oil commodities, research on patterns of pass-through or its determinants has received little attention so far (Verheyen, 2013). In particular, few studies have addressed issues of exchange rate and price pass-through in export or import demand equations for agricultural commoditiesâĂŢwhere the price volatility of recent years (Ganneval, 2016) and exchange-rate interdependence (Hatzenbuehler, Abbott, and Foster, 2016) challenge the standard assumption of linear and symmetric market reactions.

We focus on the international trade in a high-value commodity, raw coffee, a market that features many of the current trade economics topics discussed above. With the liberalization of the international coffee market in 1989 (Lindsey, 2003; Russell, Mohan, and Banerjee, 2012), strategic sourcing behavior on the part of U.S. roasters (International Trade Centre, 2011; Karp and Perloff, 1993) and shifts in consumer preferences and demand across Asia following the "latte revolution" (Chu, 2013; Ponte, 2002) have added to the complexity of trade elasticities in commodity supply (Soderbery, 2015). Yet despite the strong relevance of coffee trade to many export-oriented developing and emerging economies, little empirical research is available.

This paper generates estimates of trade elasticities for the U.S. raw coffee import trade to detect trade developments and their implications for policy. As a classic "dollar commodity," the vast bulk of coffee trade exposes exporters to U.S. dollar (USD) exchange-rate risks, adversely affecting export revenues. Major movements in local currency unit (LCU)–USD exchange rates directly affect the relative pricing of different coffee qualities, which are often linked to specific origins (e.g., Brazilian vs. Colombian Arabica beans), and the relative prices of raw coffee are a key competitive factor in a highly concentrated U.S. coffee roasting industry (Lindsey, 2003). In order to satisfy consumers' quality expectations, U.S. roasters' sourcing strategies present a complex repeated game of minimizing input costs while maintaining output quality under seasonal availability constraints, which leads to slow and lagged responses in trade flows to movements in important macroeconomic factors. "Today... macroeconomic moves in wealth create changes in currencies on a daily basis, resulting in a situation where for many exporters currency fluctuations now have a much greater impact on domestic coffee prices than ever before." (International Trade Centre, 2011, p. 164).

Our particular focus therefore lies on relaxing the standard assumptions of symmetry in the U.S. coffee import trade to fluctuations in exchange rates and trends in U.S. demand previously neglected in many aggregate analyses of the effects of exchange-rate movements on commodity trade (e.g., Bahmani-Oskooee, Harvey, and Hegerty, 2015). In other words, under volatile world market conditions, coffee exporters (and U.S. importers) can no longer be certain of the real prices of their goods. This increased risk is likely to affect expected profits and thus may influence exporters' decision to trade. While exchange-rate uncertainty has been shown to have affected bilateral commodity trade, the empirical evidence regarding the direction in which exchange-rate movements affect trade flows is mixed (Bahmani-Oskooee, Harvey, and Hegerty, 2015). Moreover, possible asymmetries in traders' responses to fluctuations in an importer's domestic demand conditions have not been addressed in previous literature.

Our empirical approach implements recent advances in modeling underlying structural asymmetries by employing Shin, Yu, and Greenwood-Nimmo's (2014) nonlinear autoregressive distributed lag (NARDL) model, which is capable of testing the extent to which nonlinearities matter in short- and long-run adjustment behaviors. Our analysis does not strive to fully characterize U.S. coffee trade or its suppliers. Instead, we draw attention to how critical relaxing assumptions of symmetric and linear exchange-rate and income effects is to the estimation of reliable trade elasticities. And what difference does this added flexibility make to implications derived from such analysis? Our main empirical purpose is to test this hypothesis using raw coffee trade-flows between major suppliers and the United States, the world's single largest importer.

Raw coffee (green coffee beans) is the single most-traded non-oil commodity in the world. Almost 112 million 60 kg bags (6.72 million tons) of raw coffee were exported in 2015 (International Coffee Organization, 2016). Annual coffee export revenues exceed the GDPs of countries like El Salvador, Estonia, or Afghanistan. On the production side, the geo-location of coffee plantations (e.g., latitude), soil conditions, management (e.g., trees per acre), and specific climatic factors interact to determine both the quantity and the quality of beans harvested, with different optimal conditions for Arabica and Robusta bean varieties. A major factor in the export supply of raw coffee is stockholding in what is otherwise an inelastic tree-crop production system. Once harvested and processed into dried raw beans, coffee becomes a storable commodity, allowing exporters to respond to varying world market conditions and demand fluctuations in destinations markets. Significant dependence on coffee production and exports (e.g., Honduras 20%, Guatemala 13%, Colombia 6% of GDP) makes the trade in raw coffee a vital contributor to employment, tax income, foreign trade earnings, and agriculture-led development in many producing economies (International Trade Centre, 2011).

Despite its significance as a tradable commodity of importance to both producers and major importers around the world, most empirical analyses of the determinants of coffee trade predate the demise of the International Coffee Agreement in 1989 (International Coffee Organization, 2016). Formerly highly regulated by quota systems and controlled by state trading agencies, the liberalization of the coffee trade has triggered dramatic changes in the international market. Early studies of U.S. import demand—relying on simple reduced-form models to estimate trade elasticities—characterized coffee as an inferior good with declining consumption as U.S. household incomes rose (Daly, 1958; Hughes, 1969; Parikh, 1973). Goddard and Akiyama (1989) and Houston, Santillan, and Marlow (2003) estimated U.S. import demand in light of shifting consumption preferences. Subsequently, however, economists' attention has been channeled away from coffee trade and toward case studies attempting to quantify the impacts of coffee trade liberalization on producer welfare (e.g., Hussien, 2015; Bacon et al., 2014), roaster/trader market power on production and trade (e.g., Igami, 2015), and growing consumer demands for differentiated origins and fair trade coffee in mature markets as well as, increasingly, among Asian middle-class consumers (e.g., Ponte, 2002; Hainmueller, Hiscox, and Sequeira, 2015). Finally, several studies have investigated price formation and vertical transmission in different retail markets (e.g., Bettendorf and Verboven, 2000; Mehta and Chavas, 2008; Nakamura and Zerom, 2010; Lee and Gómez, 2013).

This paper contributes to the literature in a number of areas. We estimate trade elasticities for the most traded agricultural commodity and its largest import market and update empirics that are now more than twenty-five years old. The analysis applies Shin, Yu, and Greenwood-Nimmo's (2014) nonlinear autoregressive distributed lag (NARDL) model to explicitly test for nonlinearities in trade equations with regard to exchange-rate and income variations in the short and long-run. Our analysis adds to a small literature on the effects of incomplete pass-through in markets for non-oil commodities, a topic of particular interest to developing countries oriented toward commodity exports. Reliable estimates of the degree of nonlinearities in the coffee trade are increasingly important to understanding trader behavior in a market of highly origin-differentiated qualities under seasonal and availability constraints, where clearly not all coffee is the same.

Methodology

In line with Bahmani-Oskooee and Ardalani (2006) or Verheyen (2014), we start with a classic reduced-form import-demand function for raw coffee in the U.S. market. Equation (1) expresses the long-run relationship between the logarithm of the value of U.S. raw coffee imports (in USD), y_t, for Arabica and Robusta bean qualities, to the logarithm of the USD/LCU real exchange rate, e_t, and the logarithm of U.S. domestic per capita GDP (or income), i_t, in a linear model:

$$y_t = a + \beta_1 e_t + \beta_2 i_t + u_t. \tag{1}$$

If all variables in equation (1) share the same order of integration and the residuals, u_t, are stationary, then β_1 and $\beta_2 \sim$ represent long-term trade elasticities and the error-correction model (ECM) in equation (2) can be estimated to assess short-term dynamics:

$$\Delta y_t = \lambda u_{t-1} + \sum_{n=1}^{\psi} \gamma_n \Delta y_{t-n} + \sum_{n=0}^{\psi} (\rho_{1,n} \Delta e_{t-n} + \rho_{2,n} \Delta i_{t-n}) + \varepsilon_t, \tag{2}$$

where λ is the error-correction term expected to belong to the interval $(-1;0)$. If $\lambda = -1$ holds, then the adjustment to long-run equilibrium is immediate. If λ is equal to 0, no long-run relationship between variables exists and only short-term dynamics can be estimated. Estimates of λ outside the interval suggest no cointegration relationship between variables. Parameters γ and ρ represent short-term dynamics, where $\rho_{1,0}$ is the impact elasticity of exchange rates on trade volumes and $\rho_{2,0}$ is the contemporaneous reaction of trade volumes to changes in U.S. import demand.

The main shortcomings of the two-stage ECM include deferring possible short-term dynamics into the residual term of the long-run relationship equation and the weak power of unit root tests, especially in short data samples (Banerjee et al., 1993; Banerjee, Dolado, and Mestre, 1998). The single equation ECM approach overcomes these issues and allows the econometrician to take advantage of Pesaran et al.'s (2001) bounds-testing approach to test for cointegration between variables of different degrees of integration: I(0), I(1), or mixed. Moreover, bounds testing has been shown to possess superior short-sample properties (e.g., Tang, 2001) and can also be applied in the presence of structural breaks (Belke and Polleit, 2006).

Following Pesaran and Shin (1999), the above two-stage error-correction model can be replaced by a single-step auto-regressive distributed lag model (ARDL):

$$\Delta y_t = a + \lambda y_{t-1} + \phi_1 e_{t-1} + \phi_2 i_{t-1} + \sum_{n=1}^{\psi} \gamma \Delta y_{t-n} + \sum_{n=0}^{\psi} (\rho_{1,n} \Delta e_{t-n} + \rho_{2,n} \Delta i_{t-n}) + \xi_t, \tag{3}$$

where λ, γ, and ρ maintain their above interpretation. Long-term elasticity parameters can be obtained from

$$\beta_1 = -\frac{\phi_1}{\lambda}; \sim \beta_2 = -\frac{\phi_2}{\lambda}. \tag{4}$$

The main advantage of the ARDL model, however, lies in its flexibility. To investigate whether there are nonlinearities with respect to the impact of USD/LCU exchange-rate fluctuations and/or fluctuations in U.S. demand on coffee trade, a nonlinear ARDL approach (NARDL)—a generalization of the ARDL bounds-testing approach (Shin, Yu, and Greenwood-Nimmo, 2014)—allows the asymmetric long-run as well as short-run coefficients to be directly estimated in a cointegration framework.

Recent studies based on Shin, Yu, and Greenwood-Nimmo's (2014) NARDL approach indicate asymmetric pass-through of exchange rates to export volumes (Verheyen, 2013). Fedoseeva and Werner (2016) conclude that exporters with endogenously determined markups adjust their prices conditional on the magnitude and direction of exchange rate and/or cost shocks. The main benefits of assessing asymmetries in cointegration relationships include the possibility of addressing the existence of underlying asymmetries in both the short and long runs, including testing for hidden cointegration, which standard cointegration techniques are incapable of. As such, Shin, Yu, and Greenwood-Nimmo (2014) advance Houck's (1977) idea of decomposing explanatory variables into partial-sum processes and nest them into a time-series ECM approach, which converts an ARDL into the NARDL:

$$x_t = x_0 + x_t^+ + x_t^-, \tag{5}$$

where x_0 is the value of x at time t_0 and x_t^+ and x_t^- are positive and negative partial sum processes of changes of x:

$$x_t^+ = \sum_{n=1}^{t} \Delta x_n^+ = \sum_{n=1}^{t} \max(\Delta x_n, \sim 0), \tag{6}$$

$$x_t^- = \sum_{n=1}^{t} \Delta x_n^- = \sum_{n=1}^{t} \min(\Delta x_n, \sim 0). \tag{7}$$

Substitution of the USD/LCU exchange rate and foreign demand (U.S. per capita GDP) variables in equation (3) by their partial sum processes results in

$$
(8)\quad \begin{aligned}
\Delta y_t &= \beta_0 + \lambda y_{t-1} + \phi_1^+ e_{t-1}^+ + \phi_1^- e_{t-1}^- + \phi_2^+ i_{t-1} + \phi_2^- i_{t-1} + \sum_{n=1}^{\psi} \gamma \Delta y_{t-n} \\
&\quad + \sum_{n=0}^{\psi} \left(\rho_{1,n}^+ \Delta e_{t-n}^+ + \rho_{1,n}^- \Delta e_{t-n}^- + \rho_{2,n}^+ \Delta i_{t-n}^+ + \rho_{2,n}^- \Delta i_{t-n}^- \right) + \xi_t .
\end{aligned}
$$

Equation (8) introduces $\beta_0 = a + e_0 + i_0$, and the lag order ψ can be determined by BIC information criterion.

The asymmetric long-run coefficients in equation (8) can be calculated as

$$
(9)\quad \beta_1^+ = -\frac{\phi_1^+}{\lambda}; \beta_1^- = -\frac{\phi_1^-}{\lambda} \sim \text{ or } \beta_2^+ = -\frac{\phi_2^+}{\lambda}; \sim \beta_2^- = -\frac{\phi_2^-}{\lambda}
$$

for import elasticities with respect to exchange rate and U.S. demand, respectively. It is also possible, however, to specify a NARDL model with only one decomposed explanatory variable at time. Equations (10) and (11) present NARDL model specifications with asymmetry in USD/LCU exchange rates and U.S. import demand, respectively:

$$
(10)\quad \begin{aligned}
\Delta y_t &= \beta_0 + \lambda y_{t-1} + \phi_1^+ e_{t-1}^+ + \phi_1^- e_{t-1}^- + \phi_2 i_{t-1} \\
&\quad + \sum_{n=1}^{\psi} \gamma \Delta y_{t-n} + \sum_{n=0}^{\psi} \left(\rho_{1,n}^+ \Delta e_{t-n}^+ + \rho_{1,n}^- \Delta e_{t-n}^- + \rho_{2,n} \Delta i_{t-n} \right) + \xi_t ,
\end{aligned}
$$

$$
(11)\quad \begin{aligned}
\Delta y_t &= \beta_0 + \lambda y_{t-1} + \phi_1 e_{t-1} + \phi_2^+ i_{t-1} + \phi_2^- i_{t-1} \\
&\quad + \sum_{n=1}^{\psi} \gamma \Delta y_{t-n} + \sum_{n=0}^{\psi} \left(\rho_{1,n} \Delta e_{t-n} + \rho_{2,n}^+ \Delta i_{t-n}^+ + \rho_{2,n}^- \Delta i_{t-n}^- \right) + \xi_t .
\end{aligned}
$$

The presence of cointegrating relationships defined between positive and negative components of the underlying explanatory variables in the above NARDL framework—"hidden cointegration"—can be tested using the bounds-testing approach. Critical values for bounds testing are tabulated in Pesaran, Shin, and Smith (2001); we pursue a more conservative approach by relying on higher critical values by considering decomposed variables as one (which implies $k = 2$ in our model). Testing for the degree of pass-through of exchange rates and income shocks to trade volumes, the NARDL framework allows us to distinguish two types of asymmetric adjustments: (i) long-run or reaction asymmetry resulting from an inequality between positive and negative partial sums of long-run coefficients and (ii) short-run or impact asymmetry resulting from an inequality between coefficients due to contemporaneous effects. Table 1 summarizes possible combinations of NARDL model specifications of symmetric and asymmetric adjustment processes in the short and long run. Asymmetry between long-run coefficients referring to positive or negative partial sums can be assessed by testing $H_0 :\sim \phi_1^+ = \phi_1^-$ and $\phi_2^+ = \phi_2^-$ for exchange rates and income, respectively. Similarly, the existence of short-term asymmetry can be assessed by testing: $H_0 :\sim \rho_{1,n}^+ = \rho_{1,n}^-$ and $\rho_{2,n}^+ = \rho_{2,n}^-$.

Data

The empirical analysis of asymmetry in pass-through patterns in non-oil commodity trade employ data on U.S. quarterly USD-valued imports of raw coffee, in particular Arabica and Robusta beans not decaffeinated (HS-901110010, HS-901110090).[1] The data were obtained from the USDA

[1] Robusta is used in this paper for the green coffee beans from the group 901110090 of the USDA FAS GATS database that are referred there to as "Other."

Table 1. NARDL Model Specifications of Possible Adjustment Processes

Model	Reaction (A)Symmetry	Impact (A)Symmetry
1: Symmetry in exchange-rate and income effects (equation 2)	$\phi_1^+ = \phi_1^-$ and $\phi_2^+ = \phi_2^-$	$\rho_{1,0}^+ = \rho_{1,0}^+$ and $\rho_{2,0}^+ = \rho_{2,0}^+$
2: Asymmetry in exchange-rate effects, symmetric income in effects (equation 10)	$\phi_1^+ \neq \phi_1^-$ and $\phi_2^+ = \phi_2^-$	$\rho_{1,0}^+ \neq \rho_{1,0}^+$ and $\rho_{2,0}^+ = \rho_{2,0}^+$
3: Symmetry in exchange-rate effects, asymmetric income effects (equation 11)	$\phi_1^+ = \phi_1^-$ and $\phi_2^+ \neq \phi_2^-$	$\rho_{1,0}^+ = \rho_{1,0}^+$ and $\rho_{2,0}^+ \neq \rho_{2,0}^+$
4: Asymmetry in exchange-rate and income effects (equation 8)	$\phi_1^+ \neq \phi_1^-$ and $\phi_2^+ \neq \phi_2^-$	$\rho_{1,0}^+ \neq \rho_{1,0}^+$ and $\rho_{2,0}^+ \neq \rho_{2,0}^+$

(United States Department of Agriculture, Foreign Agricultural Servive, Global Agricultural Trade System, 2016) for the period 1989–2013 to capture dynamics in coffee trade in the aftermath of the ICA in 1989 and include nine major U.S. suppliers (Brazil, Colombia, Costa Rica, El Salvador, Guatemala, Honduras, Indonesia, Mexico,[2] and Peru), which together account for roughly 80% of total U.S. imports over the sample period. Import values were adjusted by a U.S. GDP deflator (2009=100, Federal Reserve Bank of St. Louis). To acknowledge seasonality in coffee harvests across regions and to allow for delays between harvesting and export decisions, the trade flow data were seasonally adjusted, which has the benefit of avoiding additional variables in the empirical analysis.

Data for real USD/LCU exchange rates were obtained from the USDA (U.S. Deparment of Agriculture, Economic Research Service, 2016) database; real GDP figures for the United States were obtained from the Federal Reserve Bank of St. Louis (Federal Reserve Bank of St. Louis, 2016) and adjusted by population size to obtain per capita GDP. All model variables were logarithmized. Descriptive statistics of model variables are presented in Online Supplement A.

Figure 1 plots U.S. origin-specific import trade of raw coffee (by value) and ICO composite indicator price as a benchmark across all origins and varieties. Despite considerable variability in U.S. imports, the leading suppliers (Brazil and Colombia) are able to maintain their leading U.S. market share position, while Mexican exports experience a pronounced decline after NAFTA in 1994. NAFTA appears to have benefited a larger number of African specialty coffee producers ,summarized in the ROW category. Due to large irregularities in trade at the extensive margin, U.S. imports from many ROW (African) suppliers cannot be accommodated in the NARDL framework. The same restriction applies to Vietnam, which emerged as a major supplier of Robusta coffee to the U.S. market only in the early 2000s but accounts for roughly 40% of the ROW share of Robusta imports after 2011. In a paper focused on estimating market power in non-oil commodity trade, Igami (2015) provides a more detailed discussion of the role of Vietnam in the U.S. coffee trade.

Despite its commodity nature, raw coffee increasingly has to be treated as a quality-differentiated good, where variety and country of origin are significant determinants of price (Ponte, 2002). While coffee qualities are commonly classified into four groups (International Trade Centre, 2011)—Colombian Milds, other Milds, Brazilian Naturals, and Robustas—official U.S. trade statistics only differentiate between aggregate Arabica and Robusta varieties (United States Department of Agriculture, Foreign Agricultural Servive, Global Agricultural Trade System, 2016). Robusta beans account for roughly 40% of world supply and a large share of Arabica beans are of nonpremium quality and are often used in blended ground-coffee products (e.g., Folger's). Since most exporters to the United States (e.g., Brazil) tend to produce and export a combination of Arabica and Robusta varieties, our classification of raw coffee along the main varieties goes beyond previous studies (e.g., Igami, 2015) and allows us to produce sets of reliable and robust estimates of trade elasticities within the given limitations of available U.S. trade data.

Value-over-quantity ratios as crude indicators of value confirm the existence of source-differentiated quality differences in U.S. imports. Arabica beans are traded at a premium, with

[2] Data for Mexico are only available beginning in 1994.

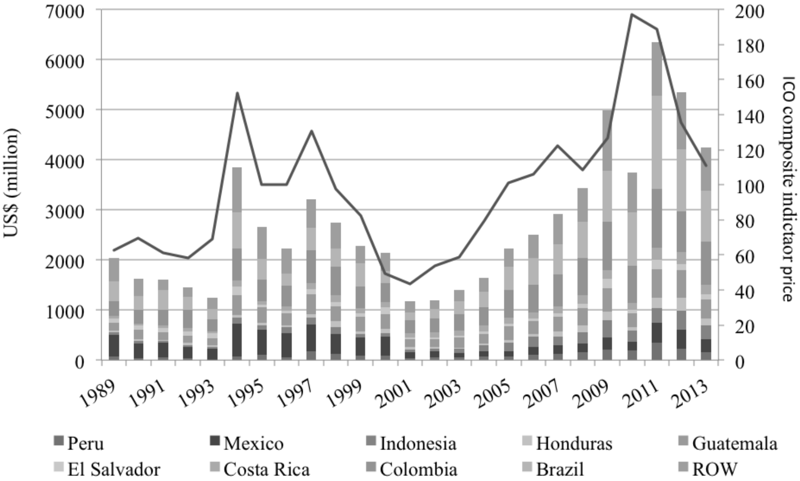

Figure 1. U.S. Raw Coffee Imports by Origin (USD million), 1989–2013

Notes: Based on data from United States Department of Agriculture, Foreign Agricultural Servive, Global Agricultural Trade System (2016) and International Coffee Organization (2016).

an average price of \$1.04/kg (Indonesia \$1.25/kg), versus Robusta coffee, with an average price of \$0.99/kg (Vietnam \$0.65/kg). Our sample of the leading nine suppliers is able to capture these differences by covering the majority of all U.S. Arabica imports and about 50% of Robusta imports, resulting in 80% coverage of quality-specific trade for the sample period.

Empirical Results and Discussion

We present estimates of long-run trade elasticity across four model specifications, Arabica and Robusta coffee qualities, and nine major U.S. import trade relationships. Models (1) impose symmetry in both exchange-rate and income effects on trade (equation 2) and models (2) allow for asymmetric adjustments to exchange-rate movements while income effects remain symmetric (equation 10). Models (3) reverse this assumption and impose symmetric exchange-rate effects while allowing for asymmetric trade responses to negative or positive income shocks over time (equation 11). Finally, models (4) allow for possible nonlinear adjustments in both exchange rates and income effects in U.S. coffee import trade (equation 8). Complete NARDL model estimation results, including diagnostic tests and results of cointegration tests, can be found in Online Supplements B and C.

Independent of model specification, the overall model fit is reasonable. The mean adjusted-R^2 value across models is 0.25 for trade in Arabica and 0.20 for trade in Robusta beans. Overall, higher R^2 values for trade in higher-quality Arabica coffee give a first indication of varietal quality differences and their likely role in determining pass-through rates of exchange rates and demand factors into U.S. sourcing decisions. In turn, lower fit values for U.S. Robusta import trade are in line with estimates reported by Parikh (1973). The majority of models pass diagnostic tests, which we report at the bottom of each results table in Online Supplements B and C. Models that do not perform well in stability tests or residual diagnostics are those for which we could not reject the H0 of no cointegration by means of bounds testing. The evidence in favor of cointegration is strong, with 15 of 36 models for Arabica and only 6 of 36 models for Robusta rejecting a cointegration relationship. Considering hidden cointegration proves to be important since the symmetric model specification (1) rejects cointegration for some trade functions, whereas introducing nonlinearities for the same relationship in models (2)–(4) reverses the rejection.

The evidence produced so far suggests that relying on linear cointegration may be too restrictive for explaining the dynamics in U.S. sourcing of raw coffee. In fact, relying on a symmetric model of U.S. coffee trade would produce no cointegration relationship for Arabica coffee from Costa Rica or Guatemala, suggesting no pass-through relationship from bilateral exchange rates or U.S. income into trade flows. Since we could not reject the null hypothesis of no cointegration in the symmetric model specification (see column 1 in table 2) for Costa Rica or Guatemala, respectively, we would conclude that U.S. imports from these origins are unaffected by income or exchange-rate movements if we only considered a simple symmetric trade model. Nonlinear specifications, however, prove that such relationships do indeed exist. In 51 of 72 trade relationships, the hypothesis of no long-run relationship between level variables cannot be rejected. Additionally, 27 of 54 estimated models with at least one threshold in place show at last one significant nonlinear cointegrating relationship. In models where cointegration does hold—irrespective of whether asymmetries are considered—allowing for asymmetric pass-through tends to increase their explanatory power. In such cases the estimates of symmetric specifications reveal trade elasticities that tend to provide some average description of a more complex case. Recent empirical research on firm-level exporting has emphasized the role of exporter heterogeneity and product quality in shaping the patterns of bilateral trade flows (Bastos and Silva, 2010).

Table 2 provides complementary evidence, showing that differences in origin-based quality matter to trade. Sets of coefficient estimates that are not bold are those for which no cointegration relationship by means of bounds testing could be established. For example, Brazilian Robusta or Colombian Arabica imports, leading suppliers to the U.S. market, respond very selectively to only positive or negative movements in both of the independent variables. Looking at the coefficient estimates in table 2 in more detail, the USD/LCU exchange-rate elasticities of imports are in line with theory. U.S. imports decline as the USD depreciates and expand again when the USD appreciates, enabling producers to maximize LCU returns in USD-denominated coffee exports to the United States.

Overall, we find U.S. import demand to be mostly exchange-rate elastic (e.g., Colombia, Arabica model (1), ER: -1.74), suggesting that roasters' origin-specific sourcing decisions respond to price signals emanating from currency fluctuations (International Trade Centre, 2011). However, the magnitudes of exchange-rate effects differ substantially across U.S. suppliers and coffee varieties. Exchange-rate elasticities for trade in Arabica beans tend to be more elastic than for Robusta (e.g., Colombia, Robusta model (3), ER: -0.54). This result confirms previous research that U.S. roasters' sourcing strategies focus on minimizing higher-priced Arabica varieties. This is commonly achieved by either substituting between alternative suppliers of similar-quality beans and/or through advanced roasting technologies that allow larger shares of cheaper Robusta beans in roast coffee blends without compromising taste (International Trade Centre, 2011; Lewin, Giovannucci, and Varangis, 2004; Ponte, 2002). For instance, U.S. imports of Arabica beans from Colombia, the leading Arabica exporter in the U.S. market (figure 1), feature a less elastic adjustment compared to comparable qualities from Guatemala, Honduras, or Peru. For lower-quality imports of Robusta beans, trade with Brazil is exchange-rate inelastic, pointing to the country-leading supplier position for Robusta coffee, compared to its competing export suppliers.

The evidence also exhibits considerable difference in elasticities between suppliers of different sizes (of coffee sector) and geo-locations. For instance, exchange-rate appreciations have a smaller negative effect on Colombian and Peruvian Arabicas (e.g., ER(+) 3.95 vs. ER($-$) 3.09) or Robusta imports from Honduras, than the positive import-increasing effect due to exchange-rate depreciations. In contrast, U.S. Arabica imports from Guatemala and Indonesia and Robusta imports from Peru, Mexico, and Indonesia increase less due to exchange-rate depreciation than they decrease due to appreciations.

For the case of Mexican Arabica, we were not able to reject the null hypothesis of no cointegration between variables in any of the model specifications. As a result, we cannot discuss the results for Mexican Arabica imports in detail. Despite being a steady supplier of raw coffee to the

Table 2. Long-Run U.S. Coffee Import Trade Elasticities

		Arabica Imports				Robusta Imports			
		(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Brazil (BR)	ER	-1.48^{**}		-1.22^{***}		$\mathbf{-0.77}^{**}$		$\mathbf{-0.72}^{*}$	
	ER(−)		1.61^{***}		1.43^{**}		$\mathbf{0.77}^{**}$		$\mathbf{0.88}$
	ER(+)		-1.06^{**}		-1.10^{*}		$\mathbf{-0.66}$		$\mathbf{-0.61}$
	GDP	6.85^{***}	3.32			$\mathbf{-1.35}$	$\mathbf{-2.02}$		
	GDP(−)			1.28	-0.97			$\mathbf{2.16}$	$\mathbf{0.42}$
	GDP(+)			4.38^{**}	3.51			$\mathbf{-1.53}$	$\mathbf{-2.31}$
Colombia (CO)	ER	$\mathbf{-1.74}^{***}$		$\mathbf{-1.62}^{***}$		$\mathbf{-0.54}$		$\mathbf{-1.68}^{***}$	
	ER(−)		$\mathbf{1.73}^{***}$		$\mathbf{1.65}^{***}$		$\mathbf{0.71}$		$\mathbf{2.77}^{***}$
	ER(+)		$\mathbf{-1.58}^{***}$		$\mathbf{-1.57}^{**}$		$\mathbf{-2.41}^{*}$		$\mathbf{-0.49}$
	GDP	$\mathbf{2.42}^{***}$	$\mathbf{1.68}$			$\mathbf{-3.45}^{**}$	$\mathbf{3.07}$		
	GDP(−)			$\mathbf{-0.90}$	$\mathbf{-0.89}$			$\mathbf{-13.33}$	$\mathbf{-23.81}^{***}$
	GDP(+)			$\mathbf{1.98}^{**}$	$\mathbf{1.65}$			$\mathbf{0.62}$	$\mathbf{-3.44}^{*}$
Costa Rica (CR)	ER	-0.68		$\mathbf{-0.10}$		$\mathbf{1.25}$		$\mathbf{-2.57}^{*}$	
	ER(−)		0.19		2.96		$\mathbf{0.22}$		$\mathbf{3.32}^{***}$
	ER(+)		2.05		7.07		$\mathbf{-5.90}$		$\mathbf{-0.71}$
	GDP	5.82^{***}	3.17			$\mathbf{-4.57}^{**}$	$\mathbf{1.73}$		
	GDP(−)			$\mathbf{-2.18}$	-19.39^{*}			$\mathbf{-15.38}$	$\mathbf{-19.99}^{**}$
	GDP(+)			$\mathbf{4.70}^{**}$	0.37			$\mathbf{0.10}$	$\mathbf{-1.00}$
El Salvador (SV)	ER	0.67		1.19		-0.87		-1.13	
	ER(−)		$\mathbf{1.19}$		$\mathbf{0.62}$		0.45		0.95
	ER(+)		$\mathbf{3.63}$		$\mathbf{3.33}$		-2.00^{**}		-1.72
	GDP	2.93	$\mathbf{-1.71}$			-5.65^{**}	-5.19		
	GDP(−)			6.98	$\mathbf{5.17}$			1.16	1.91
	GDP(+)			0.38	$\mathbf{-2.01}$			-4.57	-4.86
Guatemala (GT)	ER	-2.48^{**}		$\mathbf{-2.94}^{***}$		$\mathbf{-0.19}$		$\mathbf{-3.01}^{***}$	
	ER(−)		$\mathbf{2.62}^{***}$		$\mathbf{1.12}$		$\mathbf{0.27}$		$\mathbf{1.89}$
	ER(+)		$\mathbf{-3.85}^{**}$		$\mathbf{-5.02}^{*}$		$\mathbf{-5.52}$		$\mathbf{-4.30}$
	GDP	-0.25	$\mathbf{0.81}$			$\mathbf{-4.70}^{**}$	$\mathbf{1.67}$		
	GDP(−)			$\mathbf{-1.54}$	$\mathbf{4.34}$			$\mathbf{-11.04}^{**}$	$\mathbf{-7.40}$
	GDP(+)			$\mathbf{-0.23}$	$\mathbf{3.03}$			$\mathbf{-2.90}^{**}$	$\mathbf{-0.74}$
Honduras (HN)	ER	$\mathbf{-3.02}^{*}$		$\mathbf{0.41}$		$\mathbf{-1.15}$		$\mathbf{-1.79}^{*}$	
	ER(−)		$\mathbf{4.06}^{**}$		$\mathbf{-5.52}^{*}$		$\mathbf{0.59}$		$\mathbf{9.20}^{**}$
	ER(+)		$\mathbf{-0.98}$		$\mathbf{0.71}$		$\mathbf{-1.09}$		$\mathbf{-2.41}^{***}$
	GDP	$\mathbf{0.19}$	$\mathbf{-3.69}^{**}$			$\mathbf{-3.31}^{**}$	$\mathbf{-1.44}$		
	GDP(−)			$\mathbf{17.55}^{***}$	$\mathbf{22.99}^{***}$			-6.65	$\mathbf{-17.49}^{**}$
	GDP(+)			$\mathbf{-1.49}$	$\mathbf{3.84}$			-0.92	$\mathbf{-8.06}^{*}$
Indonesia (ID)	ER	$\mathbf{-1.05}^{***}$		-0.90^{**}		$\mathbf{-0.76}$		-0.79	
	ER(−)		$\mathbf{1.41}^{***}$		$\mathbf{0.12}$		$\mathbf{0.61}$		$\mathbf{-0.49}$
	ER(+)		$\mathbf{-1.86}^{***}$		$\mathbf{-1.28}^{***}$		$\mathbf{-1.81}^{*}$		$\mathbf{-1.64}^{*}$
	GDP	$\mathbf{8.25}^{***}$	$\mathbf{11.27}^{***}$			$\mathbf{0.96}$	$\mathbf{8.80}^{*}$		
	GDP(−)			-3.00	$\mathbf{-1.76}$			-1.28	$\mathbf{-1.17}$
	GDP(+)			6.35^{**}	$\mathbf{11.47}^{***}$			0.96	$\mathbf{10.62}^{**}$
Mexico (MX)	ER	6.86^{**}		6.68^{**}		$\mathbf{-2.06}$		$\mathbf{-2.88}^{**}$	
	ER(−)		-3.29		2.07		$\mathbf{2.73}^{**}$		$\mathbf{0.84}$
	ER(+)		4.05		4.71^{*}		$\mathbf{-4.45}^{***}$		$\mathbf{-4.66}^{***}$
	GDP	5.14^{*}	1.95			$\mathbf{-10.33}^{**}$	$\mathbf{-2.61}$		
	GDP(−)			-2.18	-14.75			$\mathbf{3.06}$	$\mathbf{7.22}$
	GDP(+)			4.55	-12.22			$\mathbf{-7.77}^{***}$	$\mathbf{2.36}$
Peru (PE)	ER	$\mathbf{-3.19}^{**}$		$\mathbf{-3.90}^{**}$		$\mathbf{1.06}$		$\mathbf{-2.43}^{**}$	
	ER(−)		$\mathbf{3.95}^{***}$		$\mathbf{5.11}^{***}$		$\mathbf{-1.42}$		$\mathbf{2.15}^{*}$
	ER(+)		$\mathbf{-3.09}^{**}$		$\mathbf{-3.65}^{**}$		$\mathbf{0.47}$		$\mathbf{-2.45}^{**}$
	GDP	$\mathbf{6.81}^{***}$	$\mathbf{5.05}^{**}$			$\mathbf{-2.23}^{*}$	$\mathbf{2.15}$		
	GDP(−)			$\mathbf{-6.97}$	$\mathbf{-11.64}$			$\mathbf{-23.88}^{***}$	$\mathbf{-22.91}^{***}$
	GDP(+)			$\mathbf{7.16}^{**}$	$\mathbf{6.20}^{**}$			$\mathbf{5.60}^{***}$	$\mathbf{6.08}^{***}$

Notes: Single, double, and triple asterisks (*,**,***) indicate significance at the 10%, 5%, and 1% level. Coefficients in **bold** refer to models for which the H_0 of no cointegration was rejected by means of the bounds-testing approach. Note that negative partial sums of exchange rates and GDP series contain only negative values. The respective coefficient estimates were multiplied by (−1) for interpretation purposes.

U.S. market, Mexico does not feature a large coffee-growing sector when compared to Colombia, Brazil, or Vietnam, producers dominating U.S. import trade. However, the estimated ER coefficients across models (1) and (3) suggest that the Mexican export trade is highly exchange-rate elastic, which confirms Mexico's role as a "niche" supplier (International Trade Centre, 2011). In other words, Arabica beans produced by Mexican growers can rather easily be substituted with beans from Latin American growers in Guatemala, Honduras, Costa Rica, and Colombia.

The question of how changes in demand affect import trade has existed for decades (Bahmani-Oskooee and Hegerty, 2007). Verheyen (2013) proxied U.S. import demand for industrial imports using the index of industrial production. As coffee imports predominantly serve the U.S. consumer market, we use U.S. per capita GDP to estimate the degree to which changing market preferences (Ponte, 2002) are passed through into short- and long-run imports. The GDP coefficient estimates in table 2 suggest that U.S. demand affects exporter supply predominantly positively for Arabicas (e.g., Arabica model (4) Guatemala, GDP(+) 3.03) and rather negatively for Robustas (e.g., Robusta model (3) Guatemala, GDP(−) -0.74) and in most cases over-proportionally. However, both magnitudes and direction of pass-through differ considerably across export origins. For instance, Arabica exports from Honduras increase when U.S. GDP declines (−1.49), indicating its lower quality. This underscores the role that market preferences (through willingness-to-pay) play in roaster sourcing strategies over time (Lewin, Giovannucci, and Varangis, 2004) and the fact that coffee clearly cannot be treated as a homogeneous commodity.

However, it is important to remember that all coefficient estimates are expressed as elasticities when considering the magnitude of effects. For instance, with the exemptions of the leading Robusta supplier, Indonesia, models with asymmetric GDP effects suggest that imports of lower quality Robusta beans decline with higher U.S. per capita GDP. Imports of sought-after Arabica beans, however, are less responsive to decline in U.S. demand. Colombia is a case in point. Models (3) and (4) indicate elastic and significantly asymmetric responses in long-run sourcing due to changing demand in the U.S. market.

Although we mainly focus on long-run asymmetries to explain the dynamics of U.S. sourcing of raw coffee, asymmetries are also apparent in short-run estimates of exchange-rate and demand effects on trade flows. Online Supplement D summarizes the results of formal symmetry tests. For example, short-term asymmetric responses to exchange-rate variability affect trade in Arabicas from Honduras and Indonesia (niche suppliers in this segment), allowing U.S. importers to take advantage of favorable short-term relative price effects (see Online Supplement tables B1–B4). In the case of Honduras, this short-term import response reverts and overshoots the initial trade effect in the third quarter. We find similar effects for trade in Honduran and Peruvian Robusta beans.

These finding further substantiate the critical role that asymmetries in trader behavior play in explaining the dynamics of pass-through patterns in major commodity supply chains, topics that have become of great importance to policy decision-making, trade risk management, and the welfare of economic agents. To this effect, our results add new and detailed insights into the questions of how trade flows should respond to fluctuations in exchange rates and/or importer income, effects not unequivocally accepted in the literature (Bahmani-Oskooee and Hegerty, 2007). Adding evidence to the debate on the expected (negative or positive) effect of exchange-rate variability on trade flows (McKenzie, 1999; Bahmani-Oskooee and Hegerty, 2007), our estimates show that some U.S. coffee imports do expand irrespective of adverse exchange-rate and/or domestic income shocks. Cases in point are U.S. income effects for Arabica imports from Honduras or Indonesia. The steady expansion of Indonesian Arabica and Robusta coffee can be attributed to the steady quality improvements by producers that to go hand in hand with the popularity of Indonesian coffee (e.g., Java and Sumatra) in a growing U.S. market for café-style espresso (International Trade Centre, 2011). This is an interesting finding given that Goddard and Akiyama (1989) isolated Indonesian Robusta exports as a separate category because of their poor quality.

The evidence produced in this analysis broadly suggests that differential pass-through rates both in magnitude and direction emerge when the nonlinear impacts of exogenous shifters are considered.

The effect of asymmetries on trade is particularly pronounced for estimates of income pass-through where symmetric specifications consistently underestimate the magnitudes of elasticities (e.g., Indonesian Arabica trade).

Several sets of results (e.g., Honduras) clearly demonstrate that under-imposed symmetry model estimates are converging on average effects that in most cases provide some information but are unsuitable to inform trade policy or sector strategy (International Trade Centre, 2011). Hence, premature interpretation of symmetric trade elasticities, especially as inputs in further modeling exercises, should be treated with caution.

Conclusions

The analysis presented in this paper has been motivated by a number of reasons, including a renewed interest in modeling the impact of fluctuations in macroeconomic factors—particularly exchange rates and incomes—on bilateral trade flows in the long run and more generally the neglect of nonlinear pass-through patterns in non-oil commodity trade. In light of the drastic structural changes in the international market for raw coffee since the demise of the International Coffee Agreement, this paper brings the importance of nonlinear modeling to the forefront of the discussion over commodity trade dynamics that have gained the attention of policy-makers and in trade risk management, particularly for commodity-export-dependent economies.

While matters of asymmetric pass-through of crude oil prices and exchange rates are prominent in the literature, this papers adds to a small but growing literature that directly tests for nonlinearities in traders' adjustment behaviors over the short and long run. Building on classic trade theory and the latest econometric NARDL framework by Shin, Yu, and Greenwood-Nimmo (2014), we model import trade for the most traded non-oil commodity, raw coffee beans, and its largest import market, the United States, a topic of relevance to large number of export-oriented developing and emerging economies.

Our estimates of trade elasticities point to the critical role played by asymmetric and nonlinearities in the pass-through of exchange-rate and income effects in explaining long-run dynamics in U.S. commodity trade and coffee trade in particular. We find that imports exhibit mostly elastic responses in long-run and short-run sourcing due to fluctuations in exchange rates and U.S. demand. However, significant asymmetries suggest that differences in the extent to which U.S. importers depend on raw coffee of specific origins and associated qualities embedded in Arabia and/or Robusta beans play a major role in determining bilateral trade flows over time.

The presence of asymmetries and nonlinearities in pass-through patterns creates uncertainties on the side of commodity exporters and importers, which in turn affect the stability of bilateral trade relations. Coffee (commodity) exporters' choice of exchange-rate regime therefore can play a direct role in stabilizing export flows with implications for export-dependent producers and their manufacturing customers in destination markets. Moreover, knowledge of the responsiveness of U.S. coffee trade to exogenous shocks also provides valuable insights into the relative competitive position of coffee exporters in the U.S. marketplace. Thus, our analysis holds implications for both trade policy and stakeholders in commodity exports and imports concerned with the impacts that macro-economic uncertainties (e.g., exchange-rate regimes) hold for the stability of commodity trade.

While our particular focus lay on relaxing symmetry and linearity in U.S. coffee import trade to fluctuations in bilateral exchange rates and U.S. per capita income, we followed a stepwise approach of introducing thresholds, making this study the first attempt to comprehensively model asymmetry in agricultural trade. In doing so, our analysis did not attempt to capture or model a complete set of economic factors determining U.S. trade in raw coffee, possibly including spatial dimensions in production and climatic factors. Such questions would be better addressed as part of a gravity model framework, which is beyond the scope of this paper.

Finally, although it has been often noted that bounds testing is more suitable than other cointegration tests in small samples (e.g., Pesaran, Shin, and Smith, 2001; Narayan and Smyth, 2005), it still performs best asymptotically. Hence, we encourage future research to return to this issue in order to complement or challenge our findings based on longer and more robust data series. Moreover, future research on the topic of pass-through in commodity and especially coffee trade should consider additional explanatory variables and market participants (e.g., Vietnam, which was excluded in our sample due to data limitations). These additional data may be critical to gaining deeper insights into the complex interdependencies of sourcing decisions faced by importers of quality-differentiated inputs and the overall increasing importance of global trade in coffee as a high-value commodity.

[Received December 2016; final revision received September 2017.]

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Online Supplement A

Table A1. Descriptive Statistics of U.S. Raw Coffee Import Trade, 1989–2013 (N=100)

	Mean	Median	Maximum	Minimum	Std. Dev.
Arabica imports from					
Brazil (BR)	97,402	61,680	467,000	12,794	89,953
Colombia (CO)	97,925	84,196	302,000	18,012	57,467
Costa Rica (CR)	25,140	24,956	64,957	1,406	15,558
El Salvador (SV)	12,232	10,531	42,443	994	7,797
Guatemala (GT)	52,704	50,023	227,000	3,164	32,927
Honduras (HN)	13,518	9,219	72,723	0	13,553
Indonesia (ID)	22,002	14,588	91,300	1,886	19,368
Mexico (MX)	47,799	34,867	214,000	238	42,287
Peru (PE)	18,899	12,836	132,000	758	19,593
Robusta imports from					
Brazil	49,893	41,343	150,000	14,291	30,380
Colombia	46,273	38,201	114,000	11,244	26,580
Costa Rica	7,415	5,121	27,948	765	5,624
El Salvador	10,563	6,568	56,478	362	9,797
Guatemala	26,134	20,955	108,000	4,320	17,064
Honduras	6,694	4,640	25,869	483	5,603
Indonesia	14,296	13,097	44,355	1,848	8,651
Mexico	37,092	10,704	497,000	1,818	61,786
Peru	9,589	7,680	42,601	377	7,412
US per capita GDP	43,346.88	44,615.47	50,288.55	35,044.54	5,084.71
USD real exchange rate					
Brazil (BRL)	2.21	1.95	4.77	1.04	0.77
Colombia (COP)	2,545.52	2,533.43	3,524.09	1,762.86	511.12
Costa Rica (CRC)	619.02	637.99	708.74	461.47	61.88
El Salvador (SVC)	9.96	9.52	14.60	8.54	1.49
Guatemala (CTG)	10.35	10.51	15.33	7.26	1.97
Honduras (HNL)	23.92	23.88	32.9	15.12	4.40
Indonesia (IDR)	11,030.65	9,651.89	25,460.9	7,934.22	3,434.83
Mexico (MXN)	12.86	12.37	18.79	10.72	1.47
Peru (PEN)	3.12	3.04	5.18	2.21	0.41

Notes: Arabica and Robusta import data are in USD (1,000s), deflated and seasonally adjusted (United States Department of Agriculture, Foreign Agricultural Servive, Global Agricultural Trade System, 2016). U.S. per capita GDP data are constant prices, seasonally adjusted (Federal Reserve Bank of St. Louis, 2016). USD real exchange rates are local currency per USD (**?**).

Online Supplement B. Estimation Results for Arabica Imports

Table B1. Arabica Model 1 with Double Symmetry, Equation (3)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	−14.63***	2.13	−15.44***	−5.87	10.30	13.59	−12.32***	−11.68*	−28.62***
	(4.70)	(2.63)	(5.32)	(10.51)	(6.53)	(15.96)	(4.57)	(6.78)	(8.33)
imp_{t-1}	−0.27***	−0.38***	−0.38***	−0.35***	−0.40***	−0.58***	−0.20***	−0.21***	−0.54***
	(0.08)	(0.10)	(0.12)	(0.11)	(0.12)	(0.11)	(0.06)	(0.06)	(0.12)
rer_{t-1}	−0.40***	−0.66***	−0.26	0.24	−0.99**	−1.75*	−0.21***	1.44**	−1.72**
	(0.15)	(0.18)	(0.28)	(0.72)	(0.43)	(0.98)	(0.08)	(0.73)	(0.73)
gdp_{t-1}	1.85***	0.92***	2.21***	1.04	−0.10	0.11	1.65***	1.08*	3.68***
	(0.56)	(0.25)	(0.69)	(0.86)	(0.46)	(1.29)	(0.53)	(0.58)	(0.98)
Δrer_t	−0.25	−0.44	−1.56	1.21	−0.72	−9.82***	−0.16	0.66	−2.47**
	(0.25)	(0.37)	(1.67)	(1.24)	(1.24)	(2.07)	(0.26)	(0.56)	(1.08)
Δgdp_t	0.67	1.79	2.40	3.04	3.30	−1.83	3.58	2.24	0.12
	(3.26)	(1.97)	(4.29)	(5.76)	(3.82)	(9.12)	(2.36)	(4.14)	(5.98)
Δimp_{t-1}						−0.08		−0.16***	
						(0.07)		(0.05)	
Δrer_{t-1}						11.41***		−0.40	
						(1.41)		(0.70)	
Δgdp_{t-1}						2.40		−0.14	
						(10.13)		(4.66)	
Δimp_{t-2}						−0.30***		−0.08*	
						(0.06)		(0.04)	
Δrer_{t-2}						−28.63***		0.79	
						(3.54)		(0.94)	
Δgdp_{t-2}						16.49		11.23	
						(11.23)		(8.03)	
Δimp_{t-3}						−0.11***			
						(0.04)			
Δrer_{t-3}						7.73***			
						(2.83)			
Δgdp_{t-3}						16.14			.
						(10.83)			
Adj. R^2	0.07	0.14	0.15	0.14	0.24	0.87	0.07	0.15	0.25
LM(4)	0.01	0.10	0.21	0.99	0.12	0.23	0.96	0.64	0.44
RESET	0.72	0.96	0.45	0.14	0.39	0.01	0.19	0.34	0.01
J-B.	0.49	0.12	0.10	0.00	0.02	0.14	0.06	0.00	0.03
CUSUM	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*N*	*Y*
Bounds	4.00	6.01**	4.08	3.99	3.48	10.52***	5.41**	3.86	6.85***

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table B2. Arabica Model 2 with Asymmetry in Exchange-Rate Effects, Equation (10)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	−5.49	−0.04	−7.44	15.31	4.44	47.86***	−22.68***	−0.84	−23.67*
	(8.82)	(5.19)	(14.43)	(10.26)	(7.98)	(17.14)	(7.58)	(13.78)	(14.09)
imp_{t-1}	−0.31***	−0.40***	−0.42***	−0.48***	−0.47***	−0.89***	−0.22***	−0.21***	−0.58***
	(0.08)	(0.10)	(0.13)	(0.12)	(0.09)	(0.09)	(0.05)	(0.06)	(0.12)
rer^{-}_{t-1}	−0.50***	−0.69***	−0.08	−0.57	−1.23***	−3.61***	−0.31***	0.69	−2.29***
	(0.17)	(0.18)	(0.38)	(0.69)	(0.38)	(0.87)	(0.10)	(0.64)	(0.67)
rer^{+}_{t-1}	−0.33**	−0.63***	0.86	1.74	−1.81**	−0.87	−0.41***	0.85	−1.79**
	(0.16)	(0.21)	(1.57)	(1.10)	(0.88)	(0.58)	(0.10)	(0.60)	(0.77)
gdp_{t-1}	1.03	0.67	1.33	−0.82	0.38	−3.28**	2.48***	0.41	2.93**
	(0.89)	(0.50)	(1.36)	(0.97)	(0.79)	(1.62)	(0.75)	(1.31)	(1.41)
Δrer^{-}_{t}	−0.32	−0.75	−2.67	−2.47	1.21	−20.21***	−1.12***	−2.45	−2.91
	(0.56)	(0.76)	(2.28)	(4.87)	(2.14)	(5.70)	(0.21)	(2.53)	(1.98)
Δrer^{+}_{t}	−0.15	−0.08	3.29	3.32**	−3.49	−6.06***	0.18	1.14*	−1.96
	(0.39)	(0.63)	(5.93)	(1.61)	(2.45)	(1.14)	(0.18)	(0.64)	(1.54)
Δgdp_{t}	0.21	2.21	2.37	2.64	3.04	−1.04	3.77	3.73	−1.66
	(3.43)	(2.05)	(4.96)	(5.98)	(4.14)	(10.06)	(2.40)	(4.02)	(6.01)
Δimp_{t-1}						0.06		−0.12***	
						(0.07)		(0.04)	
Δrer^{-}_{t-1}						−0.07		−2.77	
						(6.67)		(2.01)	
Δrer^{+}_{t-1}						10.47***		0.25	
						(0.87)		(0.83)	
Δgdp_{t-1}						0.97		2.24	
						(10.72)		(4.72)	
Δimp_{t-2}						−0.03			
						(0.04)			
Δrer^{-}_{t-2}						0.24			
						(6.32)			
Δrer^{+}_{t-2}						−27.31***			
						(1.96)			
Δgdp_{t-2}						7.33			.
						(9.38)			
Adj. R^2	0.09	0.16	0.17	0.19	0.30	0.88	0.12	0.11	0.26
LM(4)	0.01	0.18	0.08	0.79	0.05	0.07	0.77	0.17	0.78
RESET	0.91	0.75	0.52	0.07	0.00	0.58	0.17	0.87	0.02
J-B.	0.66	0.41	0.02	0.00	0.13	0.04	0.16	0.00	0.01
CUSUM	*Y*	*Y*	*N*	*Y*	*Y*	*Y*	*Y*	*N*	*Y*
Bounds	3.85	4.51*	3.35	4.47*	6.83***	33.42***	7.32***	3.15	5.74**

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table B3. Arabica Model 3 with Asymmetry in Income Effects, Equation (11)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	5.64***	12.74***	6.51	5.23**	11.84***	10.03**	4.58***	−0.16	11.11***
	(1.57)	(2.98)	(4.64)	(2.59)	(3.00)	(4.84)	(1.26)	(1.38)	(2.61)
imp_{t-1}	−0.32***	−0.42***	−0.40***	−0.42***	−0.48***	−0.75***	−0.20***	−0.22***	−0.58***
	(0.09)	(0.10)	(0.13)	(0.12)	(0.11)	(0.16)	(0.07)	(0.07)	(0.13)
rer_{t-1}	−0.39***	−0.68***	−0.04	0.50	−1.41***	0.31	−0.18**	1.47**	−2.26**
	(0.16)	(0.18)	(0.62)	(0.76)	(0.56)	(1.65)	(0.08)	(0.63)	(0.94)
gdp^-_{t-1}	−0.41	0.38	0.87	−2.93	0.74	−13.16***	0.60	0.48	4.04
	(1.28)	(1.12)	(2.92)	(2.36)	(1.58)	(6.63)	(0.95)	(1.90)	(4.04)
gdp^+_{t-1}	1.40**	0.83**	1.88**	0.16	−0.11	−1.12	1.27***	1.00	4.15***
	(0.62)	(0.33)	(0.93)	(0.83)	(0.46)	(1.23)	(0.42)	(0.84)	(1.61)
Δrer_t	−0.21	−0.53	−1.22	1.41	−0.70	−6.59***	−0.14	0.51	−2.31**
	(0.27)	(0.37)	(1.81)	(1.32)	(1.25)	(2.17)	(0.26)	(0.56)	(1.18)
Δgdp^-_t	1.11	−5.16	5.23	−1.79	1.38	26.11	−4.06	2.36	−8.66
	(6.01)	(3.55)	(6.90)	(11.68)	(7.21)	(19.56)	(3.97)	(7.69)	(9.41)
Δgdp^+_t	−0.85	8.39*	−0.06	7.94	6.58	−29.81	10.33**	3.31	6.13
	(6.55)	(4.42)	(9.03)	(9.14)	(7.20)	(19.01)	(4.40)	(7.70)	(14.11)
Δimp_{t-1}						0.00		−0.13***	
						(0.09)		(0.05)	
Δrer_{t-1}						10.80***		−0.50	
						(1.40)		(0.69)	
Δgdp^-_{t-1}						−17.39		−6.41	
						(21.70)		(7.85)	
Δgdp^+_{t-1}						16.96		6.03	
						(21.21)		(10.34)	
Δimp_{t-2}						−0.24***			
						(0.07)			
Δrer_{t-2}						−28.51***			
						(3.08)			
Δgdp^-_{t-2}						43.44*			
						(22.85)			
Δgdp^+_{t-2}						13.30			
						(20.55)			
Δimp_{t-3}						−0.10***			
						(0.04)			
Δrer_{t-3}						6.59**			
						(3.29)			
Δgdp^-_{t-3}						5.07			
						(22.45)			
Δgdp^+_{t-3}						32.01**			
						(15.73)			
Adj. R^2	0.09	0.19	0.15	0.16	0.28	0.88	0.09	0.09	0.25
LM(4)	0.00	0.21	0.15	0.44	0.07	0.76	0.64	0.29	0.81
RESET	0.59	0.74	0.52	0.02	0.00	0.01	0.61	0.57	0.02
J-B.	0.57	0.36	0.01	0.01	0.07	0.06	0.33	0.00	0.03
CUSUM	*Y*	*Y*	*Y*	*Y*	*N*	*Y*	*Y*	*N*	*Y*
Bounds	3.48	5.15**	12.32***	3.26	5.02**	8.93***	3.89	3.06	5.24**

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table B4. Arabica Model 4 with Double Asymmetry, Equation (8)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	5.37***	9.00***	6.89***	7.02***	8.15***	13.19***	5.22***	4.93***	6.49***
	(1.59)	(1.97)	(2.09)	(1.99)	(1.42)	(1.27)	(1.50)	(1.28)	(1.72)
imp_{t-1}	−0.32***	−0.52***	−0.46***	−0.49***	−0.45***	−0.86***	−0.36***	−0.31***	−0.58***
	(0.09)	(0.11)	(0.14)	(0.13)	(0.08)	(0.09)	(0.10)	(0.08)	(0.12)
rer^{-}_{t-1}	−0.45**	−0.85***	−1.35	−0.31	−0.51	4.76*	−0.04	−0.63	−2.96***
	(0.21)	(0.34)	(0.90)	(0.77)	(0.98)	(2.60)	(0.15)	(1.07)	(1.15)
rer^{+}_{t-1}	−0.35*	−0.81**	3.22	1.64	−2.28*	0.61	−0.46***	1.44*	−2.12**
	(0.20)	(0.34)	(2.57)	(1.11)	(1.24)	(0.68)	(0.11)	(0.77)	(0.89)
gdp^{-}_{t-1}	0.31	0.46	8.83*	−2.55	−1.97	−19.81***	0.63	4.52	6.75
	(1.64)	(2.77)	(5.07)	(2.45)	(3.30)	(5.50)	(0.99)	(3.63)	(4.59)
gdp^{+}_{t-1}	1.11	0.85	0.17	−0.99	1.38	3.31	4.10***	−3.74	3.59**
	(1.13)	(0.98)	(1.70)	(0.99)	(1.78)	(2.34)	(1.32)	(3.17)	(1.61)
Δrer^{-}_{t}	−0.26	−0.90	−3.81*	−1.11	1.77	−13.92**	−0.88***	−2.50	−4.08
	(0.58)	(0.80)	(2.16)	(5.07)	(2.41)	(6.54)	(0.22)	(2.58)	(2.87)
Δrer^{+}_{t}	−0.17	−0.51	6.47	2.92*	−4.61	−5.71***	0.05	1.62**	−1.81
	(0.40)	(0.67)	(6.05)	(1.77)	(3.35)	(1.10)	(0.20)	(0.81)	(1.54)
Δgdp^{-}_{t}	1.66	−4.46	11.79*	2.05	−2.41	17.02	−1.71	3.93	−4.43
	(6.35)	(5.10)	(6.83)	(11.26)	(8.15)	(18.76)	(3.77)	(7.07)	(10.37)
Δgdp^{+}_{t}	−1.25	8.93*	−6.93	3.09	7.41	−19.69	7.82**	4.65	1.69
	(6.81)	(4.50)	(9.58)	(9.50)	(7.60)	(20.05)	(3.91)	(8.36)	(14.63)
Δimp_{t-1}		0.05				0.02		−0.08*	
		(0.12)				(0.08)		(0.05)	
Δrer^{-}_{t-1}		−0.11				−3.48		−2.08	
		(0.85)				(8.55)		(2.27)	
Δrer^{+}_{t-1}		0.90				9.07***		0.02	
		(0.77)				(0.93)		(0.78)	
Δgdp^{-}_{t-1}		0.96				−22.60		−8.61	
		(5.39)				(19.78)		(8.32)	
Δgdp^{+}_{t-1}		2.36				24.07		8.42	
		(3.77)				(20.11)		(11.18)	
Δimp_{t-2}						−0.05			
						(0.04)			
Δrer^{-}_{t-2}						−3.52			
						(6.94)			
Δrer^{+}_{t-2}						−28.97***			
						(1.70)			
Δgdp^{-}_{t-2}						21.06			
						(14.79)			
Δgdp^{+}_{t-2}						10.18			
						(19.16)			
Adj. R^2	0.15	0.17	0.18	0.18	0.29	0.90	0.17	0.11	0.25
LM(4)	0.00	0.03	0.12	0.90	0.02	0.81	0.83	0.37	0.78
RESET	0.87	0.52	0.81	0.04	0.00	0.70	0.89	0.94	0.04
J-B.	0.64	0.33	0.00	0.00	0.09	0.11	0.04	0.00	0.01
CUSUM	*Y*	*Y*	*N*	*Y*	*N*	*N*	*Y*	*N*	*Y*
Bounds	2.88	4.58*	2.82	4.47*	6.36***	30.11***	6.34**	3.89	4.56*

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Online Supplement C. Estimation Results for Robusta Imports

Table C1. Robusta Model 1 with Double Symmetry, Equation (3)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	9.89**	17.05**	21.03***	32.88*	21.81***	31.46***	4.82	41.46***	19.38***
	(4.28)	(7.19)	(6.40)	(17.05)	(7.09)	(9.16)	(3.80)	(7.88)	(7.63)
imp_{t-1}	−0.30***	−0.29***	−0.37***	−0.42***	−0.32***	−0.58***	−0.37***	−0.31***	−0.51***
	(0.07)	(0.08)	(0.09)	(0.13)	(0.07)	(0.10)	(0.09)	(0.06)	(0.11)
rer_{t-1}	−0.23**	−0.16	0.47	−0.37	−0.06	−0.67	−0.28	−0.65	0.54
	(0.11)	(0.17)	(0.49)	(0.75)	(0.36)	(0.42)	(0.23)	(0.50)	(0.64)
gdp_{t-1}	−0.41	−1.01**	−1.71***	−2.38*	−1.52***	−1.91***	0.35	−3.25***	−1.13*
	(0.34)	(0.49)	(0.47)	(1.33)	(0.56)	(0.73)	(0.44)	(0.64)	(0.66)
Δrer_t	0.47	−0.01	−0.06	−0.50	−1.28	−0.33	0.00	−0.37	−0.73
	(0.33)	(0.60)	(2.14)	(1.22)	(1.36)	(0.80)	(0.50)	(1.30)	(0.65)
Δgdp_t	−1.72	3.28	0.42	−6.69	0.03	−9.65	6.20	−1.99	8.23
	(4.16)	(2.58)	(5.03)	(6.81)	(3.58)	(10.29)	(6.32)	(5.70)	(6.54)
Adj. R^2	0.16	0.12	0.14	0.17	0.14	0.26	0.17	0.12	0.24
LM(4)	0.18	0.00	0.01	0.91	0.52	0.23	0.41	0.82	0.31
RESET	0.50	0.81	0.50	0.00	0.82	0.38	0.12	0.18	0.03
J-B.	0.20	0.01	0.70	0.00	0.04	0.55	0.11	0.00	0.00
CUSUM	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*
Bounds	7.54***	5.24**	6.30**	3.65	7.16***	12.21***	5.94**	9.75***	8.29***

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table C2. Robusta Model 2 with Asymmetry in Exchange-Rate Effects, Equation (10)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	12.02	−6.14	−0.51	29.87	0.58	18.59	−34.14	18.00	−2.54
	(10.96)	(7.99)	(12.67)	(21.45)	(8.87)	(21.75)	(21.94)	(13.83)	(10.54)
imp_{t-1}	−0.31***	−0.44***	−0.43***	−0.42***	−0.45***	−0.59***	−0.45***	−0.40***	−0.58***
	(0.07)	(0.10)	(0.09)	(0.14)	(0.09)	(0.09)	(0.11)	(0.08)	(0.11)
rer^{-}_{t-1}	−0.23**	−0.31*	−0.10	−0.19	−0.12	−0.35	−0.27	−1.10**	0.82
	(0.11)	(0.18)	(0.53)	(0.85)	(0.32)	(0.89)	(0.32)	(0.55)	(0.73)
rer^{+}_{t-1}	−0.20	−1.05***	−2.54*	−0.83	−2.50***	−0.65	−0.81*	−1.79***	−0.27
	(0.15)	(0.37)	(1.41)	(1.35)	(0.77)	(0.46)	(0.43)	(0.50)	(0.74)
gdp_{t-1}	−0.62	1.34*	0.74	−2.16	0.76	−0.85	3.94*	−1.05	1.25
	(1.01)	(0.80)	(1.20)	(1.93)	(0.85)	(2.03)	(2.20)	(1.29)	(1.01)
Δrer^{-}_{t}	0.70	0.18	−1.24	5.07	−1.12	−7.66	−1.45	−4.43	0.77
	(0.56)	(1.25)	(2.99)	(6.94)	(2.29)	(5.97)	(0.89)	(3.04)	(1.67)
Δrer^{+}_{t}	0.45	−0.42	−0.56	−2.30	−3.57	0.91	0.21	0.25	−1.64*
	(0.50)	(0.69)	(5.40)	(2.64)	(2.46)	(0.86)	(0.53)	(1.40)	(0.99)
Δgdp_{t}	−2.02	1.54	3.29	−7.59	1.53	−4.11	5.88	0.46	9.65
	(4.37)	(2.79)	(4.85)	(7.47)	(4.02)	(10.78)	(5.97)	(6.15)	(6.77)
Δimp_{t-1}								0.01	
								(0.08)	
Δrer^{-}_{t-1}								−3.49*	
								(1.88)	
Δrer^{+}_{t-1}								1.59	
								(1.14)	
Δgdp_{t-1}								3.73	.
								(5.02)	
Adj. R^2	0.15	0.19	0.16	0.15	0.22	0.26	0.19	0.2	0.25
LM(4)	0.09	0.00	0.04	0.78	0.57	0.72	0.20	0.22	0.44
RESET	0.46	0.51	0.24	0.00	0.05	0.46	0.09	0.23	0.01
J-B.	0.17	0.00	0.70	0.00	0.37	0.40	0.00	0.05	0
CUSUM	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*N*	*Y*
Bounds	5.62**	5.02**	5.73**	3.86	7.24***	11.23***	5.41**	8.79***	6.98***

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table C3. Robusta Model 3 with Asymmetry in Income Effects, Equation (11)

	BR	CO	CR	SV	GT	HN	ID	MX	PE
C	5.62***	15.80***	16.19***	8.21***	12.26***	13.24***	8.08**	8.74***	12.73***
	(1.25)	(3.80)	(4.81)	(3.29)	(2.69)	(2.63)	(4.10)	(1.94)	(2.50
imp_{t-1}	−0.31***	−0.50***	−0.49***	−0.42***	−0.48***	−0.60***	−0.34***	−0.34***	−0.68***
	(0.07)	(0.11)	(0.10)	(0.14)	(0.09)	(0.10)	(0.10)	(0.07)	(0.12)
rer_{t-1}	−0.22*	−0.84***	−1.25*	−0.48	−1.43***	−1.08*	−0.27	−0.99**	−1.66**
	(0.13)	(0.27)	(0.67)	(0.78)	(0.54)	(0.59)	(0.33)	(0.50)	(0.80)
gdp^{-}_{t-1}	−0.66	6.68)***	7.50***	−0.49	5.26**	4.02	0.44	−1.05	16.28***
	(1.84)	(2.10)	(2.84)	(3.43)	(2.23)	(3.65)	(3.60)	(1.81)	(4.95)
gdp^{+}_{t-1}	−0.47	0.31	0.05	−1.92	−1.38**	−0.55	0.33	−2.66***	3.82***
	(0.74)	(0.51)	(0.73)	(1.58)	(0.60)	(1.23)	(1.33)	(0.83)	(1.41)
Δrer_t	0.54	−0.21	−0.88	−0.63	−1.60	−0.87	−0.02	−0.59	−2.22***
	(0.34)	(0.53)	(1.97)	(1.43)	(1.16)	(1.00)	(0.49)	(1.48)	(0.67)
Δgdp^{-}_t	0.86	1.66	−5.92	−7.46	8.08	−6.96	16.23	−1.64	−15.86
	(8.27)	(4.78)	(9.12)	(12.29)	(6.09)	(19.07)	(13.12)	(13.40)	(9.88)
Δgdp^{+}_t	−4.19	3.76	10.78	−5.46	−4.27	−8.99	−2.58	−1.62	32.49**
	(6.73)	(5.27)	(8.59)	(12.00)	(7.49)	(18.89)	(9.84)	(9.53)	(14.74)
Δimp_{t-1}								0.00	
								(0.09)	
Δrer_{t-1}								0.52	
								(0.60)	
Δgdp^{-}_{t-1}								−9.85	
								(9.41)	
Δgdp^{+}_{t-1}								8.65	
								(9.81)	
Adj. R^2	0.15	0.24	0.20	0.15	0.23	0.26	0.14	0.14	0.34
LM(4)	0.12	0.00	0.10	0.94	0.44	0.33	0.15	0.61	0.71
RESET	0.54	0.45	0.23	0.00	0.19	0.40	0.03	0.09	0.03
J-B.	0.12	0.01	0.79	0.00	0.45	0.43	0.05	0.17	0.00
CUSUM	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*
Bounds	5.70**	5.75**	6.75***	4.11	7.23***	10.34***	4.08	6.18***	8.06***

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Table C4. Robusta Model 4 with Double Asymmetry, Equation (8)

	BR	**CO**	**CR**	**SV**	**GT**	**HN**	**ID**	**MX**	**PE**
C	5.38***	9.72***	8.24***	7.20***	8.73***	11.28***	7.32***	6.99***	10.35***
	(1.19)	(1.98)	(1.64)	(2.33)	(1.74)	(1.62)	(1.68)	(1.55)	(2.00)
imp_{t-1}	−0.30***	−0.53***	−0.50***	−0.42***	−0.47***	−0.70***	−0.46***	−0.41***	−0.70***
	(0.07)	(0.11)	(0.10)	(0.14)	(0.09)	(0.10)	(0.11)	(0.08)	(0.12)
rer^{-}_{t-1}	−0.26	−1.46***	−1.66**	−0.40	−0.89	−6.41**	0.23	−0.34	−1.49*
	(0.21)	(0.42)	(0.81)	(1.01)	(1.07)	(2.71)	(0.51)	(1.08)	(0.91)
rer^{+}_{t-1}	−0.18	−0.26	−0.36	−0.72	−2.03	−1.68***	−0.76*	−1.89***	−1.71**
	(0.16)	(0.38)	(1.91)	(1.46)	(1.48)	(0.65)	(0.42)	(0.52)	(0.78)
gdp^{-}_{t-1}	−0.12	12.57***	10.01**	−0.79	3.49	12.19**	0.54	−2.93	15.95***
	(2.34)	(3.75)	(4.43)	(3.43)	(3.99)	(5.28)	(3.79)	(2.63)	(5.45)
gdp^{+}_{t-1}	−0.69	−1.81*	−0.50	−2.02	−0.35	−5.62*	4.91**	0.96	4.23***
	(1.08)	(1.10)	(1.42)	(1.94)	(1.99)	(2.90)	(2.05)	(2.54)	(1.39)
Δrer^{-}_{t}	0.75	−0.18	−1.89	3.95	−1.71	−13.16**	−1.06	−4.21	−2.91*
	(0.58)	(1.12)	(2.81)	(6.61)	(2.12)	(5.56)	(1.02)	(3.08)	(1.52)
Δrer^{+}_{t}	0.48	−0.28	1.82	−1.94	−2.34	1.51	−0.17	0.05	−1.31
	(0.52)	(0.69)	(5.15)	(2.78)	(3.62)	(1.14)	(0.58)	(1.41)	(0.92)
Δgdp^{-}_{t}	1.49	4.37	−3.61	−7.86	7.06	7.62	16.40	−0.46	−17.03*
	(8.68)	(5.43)	(9.34)	(12.07)	(6.85)	(20.05)	(12.34)	(14.13)	(10.32)
Δgdp^{+}_{t}	−4.97	5.24	8.52	−7.18	−2.88	−16.89	−7.56	2.94	36.15**
	(6.92)	(5.58)	(9.89)	(13.70)	(7.95)	(18.55)	(9.05)	(9.58)	(15.30)
Δimp_{t-1}								0.00	
								(0.09)	
Δrer^{-}_{t-1}								−4.25**	
								(2.05)	
Δrer^{+}_{t-1}								1.36	
								(1.04)	
Δgdp^{-}_{t-1}								−4.93	
								(10.11)	
Δgdp^{+}_{t-1}								13.66	.
								(10.25)	
Adj. R^2	0.13	0.25	0.19	0.13	0.21	0.30	0.21	0.19	0.33
LM(4)	0.10	0.00	0.11	0.78	0.35	0.88	0.54	0.37	0.74
RESET	0.42	0.41	0.30	0.00	0.06	0.50	0.22	0.15	0.02
J-B.	0.17	0.07	0.79	0.00	0.52	0.76	0.00	0.10	0.00
CUSUM	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*
Bounds	4.40*	5.15**	5.59**	3.40	5.71**	10.61***	5.78**	7.00***	6.58***

Notes: Single, double, and triple asterisks (*, **, ***) indicate significance at the 10%, 5%, and 1% level. Numbers in parentheses are robust standard errors. LM = Breusch-Godfrey serial correlation test, p-value. RESET = Ramsey RESET test, p-value. J.-B. = Jarque-Bera test, p-value; CUSUM = stability test, Y marks stable results and N marks results leaving 5% band. Bounds = F-stat. referring to bounds testing (for critical values see Pesaran, Shin, and Smith, 2001, Table CI(iii), k = 2, 10% = 4.14, 5% = 4.85, 1% = 6.36).

Online Supplement D:

Table D1. Results of (A)Symmetry Testing

	Arabica Imports									Robusta Imports								
	BR	**CO**	**CR**	**SV**	**GT**	**HN**	**ID**	**MX**	**PE**	**BR**	**CO**	**CR**	**SV**	**GT**	**HN**	**ID**	**MX**	**PE**
Model 2																		
$er^+_{t-1} = er^-_{t-1}$	0.11	**0.68**	0.46	**0.03**	**0.39**	**0.00**	**0.32**	0.54	**0.32**	**0.77**	**0.00**	**0.03**	0.69	**0.00**	**0.71**	**0.08**	**0.03**	**0.00**
$\Delta er^+_t = \Delta er^-_t$	0.83	**0.56**	0.42	**0.31**	**0.24**	**0.02**	**0.00**	0.20	**0.74**	**0.78**	**0.72**	**0.92**	0.39	**0.53**	**0.18**	**0.17**	**0.20**	**0.32**
$\Delta er^+_{t-1} = \Delta er^-_{t-1}$						**0.12**		0.22									**0.05**	
$\Delta er^+_{t-2} = \Delta er^-_{t-2}$						**0.00**												
Model 3																		
$gdp^+_{t-1} = gdp^-_{t-1}$	0.06	**0.61**	0.65	0.13	**0.56**	**0.06**	0.48	0.67	**0.97**	**0.87**	**0.00**	**0.00**	0.53	**0.00**	0.09	0.96	**0.22**	**0.00**
$\Delta gdp^+_t = \Delta gdp^-_t$	0.85	**0.06**	0.70	0.58	**0.67**	**0.12**	0.04	0.94	**0.48**	**0.68**	**0.81**	**0.27**	0.92	**0.31**	0.95	0.33	**0.99**	**0.03**
$\Delta gdp^+_{t-1} = \Delta gdp^-_{t-1}$						**0.36**		0.44									**0.28**	
$\Delta gdp^+_{t-2} = \Delta gdp^-_{t-2}$						**0.42**												
$\Delta gdp^+_{t-3} = \Delta gdp^-_{t-3}$						**0.38**												
Model 4																		
$er^+_{t-1} = er^-_{t-1}$	0.66	**0.94**	0.16	**0.08**	**0.38**	**0.08**	**0.04**	0.15	**0.28**	**0.74**	**0.05**	**0.57**	0.87	**0.63**	**0.04**	**0.01**	**0.14**	**0.57**
$\Delta er^+_t = \Delta er^-_t$	0.90	**0.75**	0.16	**0.50**	**0.21**	**0.22**	**0.01**	0.16	**0.53**	**0.77**	**0.95**	**0.58**	0.49	**0.90**	**0.01**	**0.52**	**0.24**	**0.42**
$\Delta er^+_{t-1} = \Delta er^-_{t-1}$		**0.43**				**0.14**		0.43									**0.04**	
$\Delta er^+_{t-2} = \Delta er^-_{t-2}$						**0.00**												
$gdp^+_{t-1} = gdp^-_{t-1}$	0.72	**0.91**	0.16	**0.47**	**0.49**	**0.00**	**0.06**	0.19	**0.41**	**0.82**	**0.00**	**0.04**	0.66	**0.50**	**0.01**	**0.14**	**0.38**	**0.01**
$\Delta gdp^+_t = \Delta gdp^-_t$	0.80	**0.10**	0.17	**0.95**	**0.46**	**0.29**	**0.14**	0.96	**0.79**	**0.62**	**0.92**	**0.47**	0.97	**0.46**	0.45	**0.18**	**0.84**	**0.03**
$\Delta gdp^+_{t-1} = \Delta gdp^-_{t-1}$		**0.86**				**0.17**		0.32									**0.29**	
$\Delta gdp^+_{t-2} = \Delta gdp^-_{t-2}$						**0.72**												

Notes: Coefficients in bold refer to models for which the H0 of no cointegration was rejected by means of the bounds-testing approach.
p-values reported.