Exchange Rate Dynamics and the Bilateral Trade Balance: The Case of U.S. Agriculture

Jungho Baek, Won W. Koo, and Kranti Mulik

This study examines the dynamic effects of changes in exchange rates on bilateral trade of agricultural products between the United States and its 15 major trading partners. Special attention is paid to investigate whether or not the J-curve hypothesis holds for U.S. agricultural trade. For this purpose, an autoregressive distributed lag (ARDL) approach to cointegration is applied to quarterly time-series data from 1989 and 2007. Results show that the exchange rate plays a crucial role in determining the short- and long-run behavior of U.S. agricultural trade. However, we find little evidence of the J-curve phenomenon for U.S. agricultural products with the United States' major trading partners.

Key Words: agricultural trade, autoregressive distributed lag approach to cointegration, bilateral trade, J-curve effect

It is conventional wisdom in economics that the J-curve theory is used to analyze the dynamic effect of exchange rate changes on trade balance. Assuming that the Marshall-Lerner (ML) condition—that the sum of domestic and foreign price elasticities of demand (in absolute value) is greater than one—holds, it is possible that, following a depreciation, an initial decline in the trade balance occurs before showing an improvement. The response of the trade balance over time resembles a tilted J shape. The J-curve phenomenon is attributed to a lagged adjustment of quantities to changes in relative prices (Magee 1973, Junz and Rhomberg 1973). For example, if there is a depreciation of the domestic currency, then the increased competitiveness in the domestic prices leads to exporting more and importing less, thereby improving the trade balance, which is known as the volume effect. At the same time, the depreciation increases the import unit value and results in a deterioration of the trade balance, which is referred to as the value (price) effect. The value effect prevails in the short run, whereas the volume effect dominates in the long run, which causes the time path of the trade balance depicted by the J-curve phenomenon.

In the literature on international economics, a plethora of studies have been conducted to examine the J-curve hypothesis in the United States and other countries. The evidence that has emerged from the literature is rather mixed. Some studies have found evidence of the J-curve phenomenon (e.g., Bahmani-Oskooee 1985, Moffett 1989), while others have found no evidence of it (e.g., Rose and Yellen 1989, Rose 1991). These studies generally can be classified into two groups. With aggregate data, the first group uses a two-country (i.e., between a country and the rest of the world) model to analyze the J-curve phenomenon (e.g., Flemingham 1988, Guptar-Kapoor and Ramakrishnan 1999). Noting that the empirical findings from the first group may suffer from aggregation bias of data, the second group employs bilateral trade data between a country and its major trading partners to test the J-curve hypothesis (e.g., Marwah and Klein 1996, Bahmani-Oskooee and Brooks 1999). More recently, a new body of literature has been emerging that explores the J-curve effect using disaggregated industry data such as individual agriculture, non-agriculture, and manufacturing (Bahmani-Oskooee

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and Ardalani 2006, Ardalani and Bahmani-Oskooee 2007, Bahmani-Oskooee and Wang 2007, Bahmani-Oskooee and Bolhasani 2008).

In the agricultural trade literature, on the other hand, most studies have typically concentrated on the effect of changes in the exchange rate on agricultural export volume and/or prices (Gardner 1981, Bessler and Babula 1987, Bradshaw and Orden 1990, Orden 1999). Limited efforts have been made to investigate the impact of exchange rate fluctuations on the agricultural trade balance. To the best of our knowledge, Carter and Pick (1989) and Doroodian, Jung, and Boyd (1999) are the only two studies that have been done to test the J-curve hypothesis for the U.S. agricultural trade balance. Carter and Pick (1989) employ the polynomial distributed lag model to test the J-curve hypothesis for the U.S. agricultural trade balance. They find empirical evidence that the first segment of the J-curve exists; that is, the depreciation of the U.S. dollar leads to a deterioration of the U.S. agricultural trade balance. Doroodian, Jung, and Boyd (1999) examine the J-curve effect for U.S. agricultural and manufactured goods using the Shiller lag model. Their results support the J-curve effect for agricultural goods, but not for manufacturing goods.

However, the previous two studies have examined the effects of an exchange rate depreciation on the agricultural trade balance between the United States and the rest of the world, instead of a bilateral model. For example, Doroodian, Jung, and Boyd (1999) construct the weighted average of foreign income and exchange variables for their analysis. Obviously, the data compilations of those variables suppress the actual movements taking place at the bilateral levels (Bahmani-Oskooee and Ratha 2004). In addition, the positive effect of the exchange rate on a country's trade balance against one trading partner could be offset by negative effects against another trading partner (Bahmani-Oskooee and Goswami 2003, Bahmani-Oskooee and Ratha 2006). As such, the findings obtained from the previous studies could suffer from aggregation bias. Moreover, the earlier studies have concentrated on the short-run exchange rate effect on the trade balance. Since the short-run effects of exchange rate changes could be different from the long-run effects, it is important to include the long-run dynamics in a model (Bahmani-Oskooee and Ratha 2004).

Furthermore, given the fluctuation of trade surplus in U.S. agriculture and the decrease in the value of the U.S. dollar during the period 2002–2007, it is very interesting to explore the effect of exchange rate changes on the trade balance. During the 2002–2007 period, for example, the value of the U.S. dollar decreased by approximately 30 percent and 6 percent against the Canadian dollar and the Japanese yen, respectively. In addition, the U.S. dollar declined by approximately 31 percent against the euro for the same six years. Following the decrease in the value of the U.S. dollar, on the other hand, the U.S. agricultural trade surplus initially deteriorated before showing an improvement during the same period. For example, the U.S. trade surplus dropped substantially, from \$11.2 billion in 2002 to \$3.9 billion in 2005, but then went up to \$5.6 billion in 2006 and \$18 billion in 2007 (Figure 1).

In this study, therefore, we attempt to extend the scope of previous work by assessing the effect of exchange rate changes on U.S. trade within the context of disaggregating industry data (i.e., agriculture) of bilateral trade. Special attention has been given to assess the characteristics of the short-run dynamics (J-curve phenomenon) and empirically determine whether or not U.S. trade in agriculture benefits from a decline in the value of the U.S. dollar. For this purpose, unlike Carter and Pick (1989) and Doroodian, Jung, and Boyd (1999), we use bilateral trade data between the United States and its 15 major trading partners, which consist of approximately 63 percent of U.S. agricultural trade (Table 1). An autoregressive distributed lag (ARDL) approach to cointegration developed by Pesaran, Shin, and Smith (2001) is employed with these data. The ARDL modeling approach has numerous advantages in comparison to standard cointegration methods such as Engle and Granger (1987) and Johansen (1995). First, the ARDL can be applied irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. The model is thus relieved of the burden of establishing the order of integration among variables and of pre-testing for unit roots (Pesaran, Shin, and Smith 2001). In addition, since a dynamic error correction model (ECM) can be derived from the ARDL via a simple linear transformation, the ARDL model integrates the short-run dynamics with the long-run equilibrium without losing long-run information. Finally, the

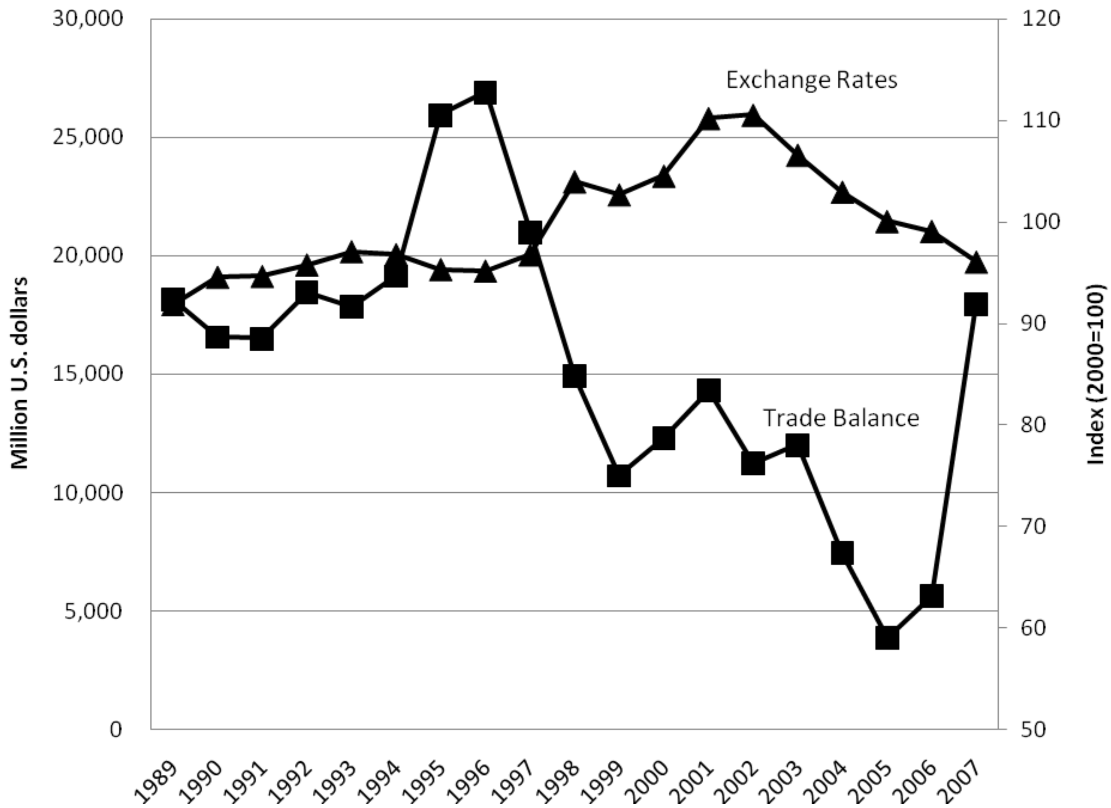

Figure 1. U.S. Agricultural Trade Balance and Real Trade-Weighted Exchange Rates
Source: Economic Research Service, U.S. Department of Agriculture (http://www.ers.usda.gov/Data/ExchangeRates/).

ARDL is more robust and performs better for small sample sizes than other cointegration techniques (Pesaran and Shin 1999). We hope that this study improves our understanding of dynamic effects of exchange rate changes on U.S. agricultural trade and contributes to the literature of international agricultural trade.

The remainder of this paper is organized as follows. The next section briefly introduces the theoretical framework for the J-curve effect. Then we describe the empirical model related to the ARDL estimation. Following that, two sections discuss the data set used in the analysis as well as the empirical results. Finally, we make some concluding remarks.

Some Theoretical Considerations on the J-Curve Effect

The theories dealing with the exchange-rate–trade-balance relationship can be classified into three categories: elasticity, absorption, and monetary approaches. In the elasticity approach, the effects of exchange rate changes on the trade balance are determined by the demand and supply elasticities of exports and imports. According to the absorption approach, the trade balance is determined by real national income and its absorption. Any improvement in the trade balance thus requires an increase of domestic income over total domestic expenditures. In the monetary approach, the trade balance is essentially a monetary phenomenon. Money and asset markets determine the trade balance (i.e., capital account) through changes in supply and demand of the stock of money. In this study, we follow the elasticity approach discussed below. However, we directly deal with the trade balance in the model, instead of analyzing the demand and supply elasticities separately.

Magee (1973) first analyzed the effects of exchange rate changes on the trade balance in the

Table 1. U.S. Agricultural Trade with Its Major Trading Partners, 2003–2007 Average

	Exports ($ mil.)	Imports ($ mil.)	Total (%)
Canada	11,138	12,537	23,675 (18.4)
Mexico	9,881	8,291	18,171 (14.1)
Japan	8,399	427	9,126 (7.1)
China	6,164	1,991	8,155 (6.3)
Netherlands	1,288	2,007	3,295 (2.6)
Italy	553	2,565	3,118 (2.4)
Korea	2,797	209	3,007 (2.3)
Australia	493	2,430	2,923 (2.3)
Indonesia	1,105	1,708	2,813 (2.2)
Taiwan	2,392	190	2,582 (2.0)
France	453	1,892	2,345 (1.8)
Brazil	318	2,012	2,330 (1.8)
Ireland	269	2,050	2,320 (1.8)
Germany	1,122	1,015	2,137 (1.7)
Colombia	775	1,330	2,105 (1.6)
Thailand	725	1,188	1,913 (1.5)
Spain	824	1,000	1,824 (1.4)
New Zealand	164	1,606	1,770 (1.4)
Chile	219	1,538	1,757 (1.4)
UK	1,187	565	1,753 (1.4)
Sub-Total	50,266	46,551	97,118 (76.0)
Total	68,978	59,580	128,559 (100.0)

Source: Foreign Agricultural Trade of the United States (FATUS), U.S. Department of Agriculture (http://www.ers.usda.gov/Data/FATUS/).

Note: Numbers in parentheses are percentage shares of total exports and imports.

framework of the elasticity approach. He identified three different periods after a devaluation in which the adjustment of the trade balance is affected by different factors: the currency-contract period, the pass-through period, and the quantity-adjustment period.

The currency-contract period is defined as the brief period immediately after a devaluation in which the export and import contracts are specified before the change. For example, consider the case in which domestic export contracts are denominated in domestic currency and domestic import contracts denominated in foreign currency. In this case, a devaluation of domestic currency increases the exchange rate expressed as domestic currency against foreign currency and immediately deteriorates the trade balance in the currency-contract period before any price and volume changes. The pass-through period is defined as the period after a devaluation in which prices can change but quantities of exports and imports remain unchanged. This is also known as the value (price) effect. This effect depends on the scale of demand and supply elasticities of exports and imports. For example, consider the situation in which both domestic and foreign demand for imports are inelastic in the short run. As a consequence of a devaluation, the import price measured in domestic currency increases but the demand stays the same, thereby resulting in an increase of value of imports (i.e., full pass-through). On the other hand, the export price in foreign currency decreases by the same proportion of the exchange rate variation (full pass-through) and the export price in domestic currency remains unchanged. To combine both the currency-contract and the pass-through effects, therefore, the trade balance in domestic currency is expected to decrease following a J-curve pattern before any trade volume changes.

The quantity-adjustment period is defined as the period in which quantities start to adjust in response to changes in prices. This is also known as the volume effect. Under this circumstance, as both export and import elasticities increase, domestic volume of exports (imports) increases (decreases) in response to the price drop (increase) in foreign (domestic) currency. As a result, the trade balance eventually improves as long as the Marshall-Lerner condition—that the sum of domestic and foreign price elasticities of demand (in absolute value) exceeds one—is satisfied.

It should be emphasized that in the currency-contract and pass-through periods, there is no logical necessity for a country's trade balance to show the initial portion of the J-curve—the deterioration of trade balance in domestic currency (Magee 1973). The necessary conditions for the initial deficit in trade balance are that (i) domestic export contracts are denominated in domestic currency and import contracts denominated in foreign currency in the currency-contract period, and (ii) domestic and foreign price elasticities of demand are inelastic and yield full pass-through in the pass-through period. As such, a devaluation of domestic currency may lead the trade balance to improve initially or remain constant according to circumstances. For example, when export contracts are denominated in foreign currency and import contracts denominated in domestic currency, the value of exports in domestic currency increases by the same percentage of the devaluation but the value of imports in domestic currency remains unchanged, thereby improving trade balance in the currency-contract period. Or, if both domestic and foreign supplies of exports are inelastic in the short run, export prices in foreign currency increase by the same proportion of the exchange rate variation but import prices in domestic currency remain unchanged (i.e., no pass-through), thereby resulting in an increase of trade balance in the pass-through period.

Development of an Empirical ARDL Model

In examining the J-curve phenomenon, economists have generally relied on a trade balance model developed by Rose and Yellen (1989). The reduced-form equation for the trade balance is specified as follows:

$$(1) \qquad TB = TB(Y, Y^*, ER),$$

where TB is the trade balance, $Y(Y^*)$ is the real income of the home (foreign) country, and ER is the real exchange rate, which is defined as $ER = E \times (P/P^*)$, where E is the bilateral nominal exchange rate of the foreign currency per unit of the domestic currency, P is the domestic consumer price index (CPI), and P^* is the foreign CPI.

To illustrate the ARDL modeling approach, we then express equation (1) in a log linear form as follows:

$$(2) \qquad \begin{aligned} \ln TB_{it} &= \alpha + \beta_1 \ln Y_{US,t} + \beta_2 \ln Y_{i,t} \\ &\quad + \beta_3 \ln ER_{i,t} + \varepsilon_t, \end{aligned}$$

where TB_{it} is the real U.S. trade balance defined as the ratio of the nominal value of U.S. imports from country i to the nominal value of U.S. exports to country i (expressed as trade deficit), i = Canada, Mexico, Japan, the Netherlands, Italy, Korea, Australia, Indonesia, France, Ireland, Germany, Thailand, Spain, New Zealand, and the United Kingdom; $Y_{US,t}$ is the real U.S. income; $Y_{i,t}$ is the real income of trading partner i; $ER_{i,t}$ is the bilateral real exchange rate between the currency of trading partner i and the United States; and ε_t is the error term. With regard to the signs of the coefficients in equation (2), it is expected that $\beta_1 > 0$ and $\beta_2 < 0$, since a rise in U.S. (trading partner) income increases U.S. imports (exports), thereby deteriorating (improving) the trade balance. However, if an increase in U.S. (trading partner) income is a result of a rise in production of import-substitute commodities, U.S. imports (exports) may decline as U.S. (trading partner) income increases. In this case, it is expected that $\beta_1 < 0$ and/or $\beta_2 > 0$ (Magee 1973, Bahmani-Oskooee 1985, Bahmani-Oskooee and Ratha 2004). As to the effect of the exchange rate, it is expected that $\beta_3 > 0$, since the depreciation of the U.S. dollar increases exports and decreases imports, thereby improving the trade balance.

Equation (2) represents the long-run equilibrium relationship among the variables of the trade balance model. As noted in the introduction, however, the main aim of this paper is to analyze both the short- and long-run impacts of changes in exchange rates on U.S. trade in agricultural goods. In addition, to test the J-curve phenomenon, we need to incorporate the short-run dynamics into equation (2). For this purpose, following Pesaran, Shin, and Smith (2001), we reformulate equation (2) as the ARDL framework. This involves estimating the error correction version of the ARDL model for variables under estimation as follows:

(3)

$$\begin{aligned}\Delta \ln TB_{it} &= \alpha + \sum_{k=1}^{p} \varepsilon_k \Delta \ln TB_{i,t-k} + \sum_{k=1}^{p} \phi_k \Delta \ln Y_{US,t-k} \\ &+ \sum_{k=1}^{p} \varphi_k \Delta \ln Y_{i,t-k} + \sum_{k=1}^{p} \gamma_k \Delta \ln ER_{i,t-k} \\ &+ \lambda_1 \ln TB_{i,t-1} + \lambda_2 \ln Y_{US,t-1} \\ &+ \lambda_3 \ln Y_{i,t-1} + \lambda_4 \ln ER_{i,t-1} + u_t,\end{aligned}$$

where Δ is the difference operator, p is lag order, and u_t is assumed serially uncorrelated. Equation (3) is called the error correction version related to the ARDL since the linear combination of lagged variables (terms with "λ"'s) replaces the lagged error-correction term (EC_{t-1}) in a standard error-correction model (ECM). Hence, while λs represent the long-run (cointegration) relationship, the coefficients following the summation signs (Σ) indicate the short-run relationships between changes in bilateral trade balance and exchange rates (i.e., J-curve effect) and changes in bilateral trade balance and income. The traditional cointegration tests such as Engle and Granger (1987) and Johansen (1995) concentrate on cases in which the underlying variables are of equal order of integration [i.e., integrated of order one, or $I(1)$]. This inevitably involves a certain degree of pre-testing and introduces a further degree of uncertainty into the analysis of level relationships (Pesaran, Shin, and Smith 2001, p. 289). To overcome the shortcomings, Pesaran, Shin, and Smith (2001) develop an alternative approach to testing for the existence of cointegration (levels) relationships that is applicable irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. Unlike conventional cointegration tests, therefore, the ARDL model avoids problems associated with non-stationary time-series data (i.e., spurious regression).

It should be pointed out that since the ARDL is based on a single-equation approach, it may not be able to correct the potential endogeneity of the independent variables and thus may yield inefficient estimates of the short- and long-run relationships (Pesaran and Shin 1999). In this study, however, since the size of the agricultural sector is small relative to the entire economy in the United States and its trading partners, the exchange rate and income are expected to behave exogenously in the agricultural sector. As a result, this economic relationship justifies the use of a single-equation procedure to estimate equation (3).

Data and Testing Procedure

Data

To examine the impact of exchange rate changes on the U.S. agricultural trade balance, we collect quarterly data for the first quarter of 1989 through the fourth quarter of 2007 (1989:1–2007:4) (76 observations). The total values of exports and imports for agricultural products between the United States and its major trading partners are obtained from the Foreign Agricultural Trade of the United States (FATUS) database of the U.S. Department of Agriculture (USDA). Based on the average 2003–2007 trade share of each trading partner, we first identify the 20 largest trading partners of the United States (Table 1). However, China, Taiwan, Brazil, Colombia, and Chile are not included in this study, due to the unavailability of data (i.e., GDP index and GDP deflator).[1] As such, the following 15 countries are chosen for the empirical analysis: Canada, Mexico, Japan, the Netherlands, Italy, Korea, Australia, Indonesia, France, Ireland, Germany, Thailand, Spain, New Zealand, and the United Kingdom.[2] The GDP deflator is used to derive real values of exports and imports (2000=100) and is obtained from the International Financial Statistics (IFS) database published by

[1] Recent economic data for Taiwan are no longer provided by the International Financial Statistics (IFS) database.

[2] Because of limited availability of GDP data in the IFS database, the data for Indonesia and Ireland contain 44 observations for 1997:1–2007:4, and Thailand includes 60 observations for 1993:1–2007:4, respectively.

the International Monetary Fund (IMF). The U.S. trade balance is then defined as the ratio of real value of U.S. imports to real value of U.S. exports with the 15 trading partners (expressed as the trade deficit). One of the major reasons for using the ratio is that it is not sensitive to the units of measurement and can be interpreted as the real trade balance. In addition, the ratio can narrow the range of the variable to make it less susceptible to outlying or extreme observations (Wooldridge 2000). Finally, the ratio can be transformed into a logarithmic form without worrying about the possible negative values.

The real gross domestic product (GDP) index (2000=100) is used as a proxy for the real income of the United States and its trading partners and is also taken from the International Financial Statistics (IFS) database published by the International Monetary Fund (IMF). The real bilateral exchange rates between the U.S. dollar and the currencies of the United States' 15 trading partners are obtained from the Economic Research Service (ERS) of the U.S. Department of Agriculture.[3] Since the exchange rate is expressed as the number of trading partner's currency per unit of the U.S. dollar, a decline in the exchange rate indicates a real depreciation of the U.S. dollar. All variables are in natural logarithms.

Testing Procedure for the ARDL

The ARDL modeling procedure starts with determining the lag length (p) in equation (3). As Pesaran, Shin, and Smith (2001) note, it is crucial to balance between choosing p sufficiently large to mitigate the residual serial correlation problems and sufficiently small so that equation (3) is not unduly over-parameterized, particularly in view of the limited time-series data which are available (Pesaran, Shin, and Smith 2001, p. 308). Hence, we use the Akaike Information Criterion (AIC) and Lagrange multiplier (LM) statistics for testing the hypothesis of no serial correlation against order 4, respectively (Table 2).[4]

With the selected lag orders, we then test the existence of a level relationship (cointegration) among variables. For this purpose, the null hypothesis of no level relationship, namely ($\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 0$) in equation (3) is tested, irrespective of whether the regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated. This can be done using an F-test with two sets of asymptotic critical values tabulated by Pesaran, Shin, and Smith (2001) in which all the regressors are assumed to be purely $I(0)$ or purely $I(1)$. This is called a "bounds testing" procedure since the two sets of critical values provide critical value bounds for all possibilities of the regressors into purely $I(0)$, purely $I(1)$, or mutually cointegrated (Pesaran, Shin, and Smith 2001, p. 290). If the computed F-statistic lies outside the upper level of the critical bounds, the null can be rejected, indicating that the variables are cointegrated. If the F-statistic falls below the lower level of the critical bounds, on the other hand, the null cannot be rejected, supporting lack of cointegration. With $p = 5$ for the U.S.-Canada agricultural trade, for example, the F-statistic is 4.11, which lies outside the upper level of the 10 percent critical bounds (Table 2).[5] As a result, the null hypothesis that there exists no cointegrated trade balance equation can be rejected, irrespective of whether the regressors are purely $I(0)$, purely $I(1)$, or mutually cointegrated.[6] However, the test statistics for the agricultural trade with Korea, Indonesia, Germany, Spain, and New Zealand are in the range between 2.87 and 3.50, which falls within the 10 percent bound. If the F-statistic lies between the two bounds, the inference is inconclusive. In these cases, following Kremers, Ericson, and Dolado (1992) and Banerjee, Dolado, and Mestre (1998), the error-correction terms in the ARDL model are used to determine the existence of cointegrated trade balance equations. Hence, if a negative and significant lagged error-correction term is obtained, the variables are said to be cointegrated.

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[3] Real exchange rates are derived by multiplying nominal exchange rates by the ratio of the U.S. to local currency consumer price index (CPI).

[4] To ensure comparability of results for different choices of lag length, all estimators use the same sample period, 1990:4–2007:4 (T=71), with the first eight observations reserved for the construction of lagged variables.

[5] With three regressors ($k = 3$), the 10 percent critical value bound is (2.72, 3.77), which is obtained from Table CI in Pesaran, Shin, and Smith (2001).

[6] To investigate whether a deterministic trend is required, we also estimate equation (3) with a linear time trend. However, the findings are more conclusive when the F-test is applied to equation (3) without a linear trend.

Table 2. Results of F-Test for Cointegration among Variables

	AIC Lags	$\chi^2_{SC}(4)$	F-statistic	Decision
Canada	5	0.78	4.11	cointegration
Mexico	6	7.16	7.80	cointegration
Japan	3	4.44	8.31	cointegration
Netherlands	8	2.32	5.63	cointegration
Italy	8	4.96	6.11	cointegration
Korea	3	2.72	3.50	inconclusive
Australia	6	6.91	5.88	cointegration
Indonesia	4	6.24	3.34	inconclusive
France	5	4.69	4.53	cointegration
Ireland	2	6.47	7.48	cointegration
Germany	4	3.60	3.06	inconclusive
Thailand	4	1.85	4.09	cointegration
Spain	5	6.82	2.87	inconclusive
New Zealand	6	2.38	3.10	inconclusive
UK	3	2.76	3.79	cointegration

Note: A lag order is chosen based on Akaike Information Criterion (AIC). $\chi^2_{SC}(4)$ are Lagrange Multiplier (LM) statistics for testing no serial correlation against order 4. F-statistic for 10 percent critical value bounds is (2.72, 3.77), which is taken from Table CI in Pesaran, Shin, and Smith (2001).

Empirical Results

Results of Short-Run Analysis: Does the J-Curve Exist?

After determining the lag order and the existence of the level relationship, the selected ARDL model outlined by equation (3) is used to estimate the short- and long-run coefficients. The results of short-run coefficient estimates show the short-run dynamic effects of the depreciation on the trade balance or the J-curve effect (Table 3).[7] The sign of the coefficient of the exchange rate determines the existence of the J-curve effect. That is, an initially negative sign followed by a positive one on the lag coefficients would be consistent with the J-curve phenomenon. The results show that, only for the U.S. agricultural trade with France, the signs of the coefficients of the current and three-period lagged exchange rate are negative, followed by a positive sign, indicating the tendency of a J-curve pattern. However, most coefficients are not statistically significant even at the 10 percent level. The results thus suggest that there is no J-curve effect for the U.S. trade with France. Additionally, in all other cases, we observe no specific pattern. Overall, therefore, the findings indicate that the J-curve does not hold for the U.S. agricultural trade with its 15 largest trading partners. This is also consistent with the general findings of previous studies that the short-run adjustment process of the trade balance to an exchange rate depreciation does not follow any specific pattern (Bahmani-Oskooee and Ratha 2004). Notice that, for all cases except Ireland, the exchange rate carries at least one significant coefficient at the 10 percent level, indicating that exchange rate is a major factor in U.S. trade to each of its trading partners' markets in the short run.

[7] In the models of agricultural trade with Canada, for example, the estimated orders of an ARDL (p, p_1, p_2, p_3) model in the four variables $(TB_{i,t}, Y_{US,t}, Y_{i,t}, ER_{i,t})$ are selected by a general-to-specific search, spanned by lag length p = 0,1,2,3,4,5 and p_i = 0,1,2,3,4,5, i = 1,2,3,4,5, using the AIC criterion [see Pesaran and Shin (1999) for details].

Table 3. Coefficient Estimates of Exchange Rate and Error-Correction Terms of the Bilateral Trade Balance Model

Country	Lag Order of Exchange Rate								
	0	1	2	3	4	5	6	7	EC_{t-1}
Canada	1.02** (3.52)	0.49 (1.49)	0.47 (1.61)	0.43 (1.38)					-0.80** (-5.33)
Mexico	0.88** (3.78)								-0.75** (-6.63)
Japan	-0.79** (-2.25)	0.11 (0.77)	-0.75** (-2.15)						-0.67** (-5.93)
Netherlands	0.10 (0.51)	-2.21** (-3.41)	-1.41** (-2.54)	-1.34** (-2.27)	-0.75 (-1.46)	-1.05** (-2.20)	-1.67** (-3.66)	-0.91* (-1.90)	-0.80** (-3.42)
Italy	-0.85** (-2.31)	0.58 (1.39)	0.58 (1.38)	-0.93** (-2.68)					-0.65** (-2.28)
Korea	0.38* (1.78)								-0.55** (-4.81)
Australia	1.35** (4.16)								-0.51** (-3.26)
Indonesia	-0.04 (-0.09)	-0.51* (-1.75)	-0.14 (-0.43)	-0.82** (-2.67)					-0.76** (-5.88)
France	-0.23 (-0.67)	-0.91** (-2.56)	-0.24 (-0.68)	-0.86** (-2.42)	0.43 (1.28)				-0.70** (-3.98)
Ireland	0.56 (0.72)								-0.53** (-3.47)
Germany	0.34** (2.09)								-0.59** (-2.46)
Thailand	-0.79** (-2.43)								-0.91** (-4.02)
Spain	1.05** (2.83)								-0.82** (-3.25)
New Zealand	0.56* (1.73)								-0.30* (-1.69)
UK	0.59** (2.15)								-0.74** (-3.86)

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics. EC_{t-1} refers to the error-correction term.

Because our analysis is based on the elasticity approach, one possible explanation for no evidence of the J-curve for agricultural goods is that the necessary condition for the J-curve in the currency-contract period may not hold for U.S. agricultural trade; that is, U.S. export contracts should be denominated in dollars and U.S. import contracts denominated in foreign currency. However, such an explanation may not be conclusive in view of the fact that the currency-contract analysis deals with a very brief period immediately following a devaluation and because the currency in which prices are quoted presumably would be changed to avoid an exchange rate loss (Magee 1973). Even though the agricultural industry is characterized by contracts that do not change subsequent to a real depreciation, currency and future markets tend to mitigate the effects of exchange rate variability on agricultural trade. Moreover, it is not likely to find qualitative or survey evidence on the currency denomination of U.S. (agricultural) trade. Thus, the most likely explanation for the finding is that U.S. agricultural trade may not meet the necessary condition for the pass-through effect; that is, U.S. and foreign price elasticities of demand are inelastic. In fact, in the short run, supply of U.S. agricultural exports is generally inelastic, while demand is relatively elastic due to the availability of other major exporters such as Australia and the European Union (Table 1). Under this circumstance, as a consequence of a depreciation, the dollar price of U.S. exports increases but the dollar price of U.S. imports remains unchanged (i.e., no pass-through). As a result, U.S. agricultural trade does not show the initial deterioration of the trade balance.

It should be pointed out that the coefficients of the error-correction terms are negative and statistically significant at least at the 10 percent level for all cases, which further provides evidence of the existence of the long-run relationship among variables (Kremers, Ericson, and Dolado 1992, Banerjee, Dolado, and Mestre 1998). The findings thus justify the ARDL modeling of U.S. agricultural trade with Korea, Indonesia, Germany, Thailand, and Spain, in which the results of the *F*-statistics are inconclusive (Table 2).

Results of Long-Run Analysis

The results of the long-run coefficient estimates of the trade balance model show that all the coefficients of the exchange rate are statistically significant at least at the 10 percent level (Table 4). More specifically, in all cases, the U.S. trade balance has a positive long-run relationship with the real bilateral exchange rate, implying that the depreciation of the U.S. dollar indeed improves the trade balance in the long run. Additionally, the coefficients of the real U.S. (foreign) income are statistically significant at the 10 percent level for all cases except Korea, Ireland, Germany, and New Zealand (Japan, Germany, Thailand, New Zealand, and the United Kingdom). For example, the U.S. trade balance with Mexico, France, Italy, and the Netherlands has a positive long-run relationship with real domestic income and a negative relationship with real foreign income. This indicates that a rise in real U.S. (foreign) income increases domestic (foreign) demand for foreign imports (domestic exports), thereby deteriorating (increasing) the trade balance. On the other hand, the U.S. trade balance with Canada, Australia, and Indonesia has a negative long-run relationship with domestic income and a positive relationship with foreign income. This suggests that an increase of real domestic (foreign) income decreases domestic (foreign) demand for foreign imports (domestic exports), thereby improving (deteriorating) the trade balance. The most likely explanation for the finding is that, since imports are defined as the difference between domestic consumption and production, an increase in domestic income could increase the domestic production of import-substitute commodities faster than a rise in domestic consumption, which thus leads to the reduction of domestic imports (Magee 1973, Bahmani-Oskooee 1985, Bahmani-Oskooee and Ratha 2004).

We emphasize here that the trade balance model dealing with imports and exports as a single variable is not able to directly identify which variable is affecting exports or imports and by how much.[8] For completeness, therefore, we also estimate agricultural exports and imports separately in order to measure the effects of exchange rate changes on the trade balance accurately (see Appendix for more details). The results show that U.S. agricultural exports have a significant relationship with the bilateral exchange rate and foreign income in both the short and long run (Tables A1 and A2 in the Appendix).[9] On the other hand, the domestic income is found to be a

[8] The authors would like to thank an anonymous referee for raising this issue.

[9] Notice that U.S. agricultural exports to the European Union countries such as the Netherlands and Italy consistently are found to have a negative long-run relationship with the real foreign income, indicating that an increase in real income of those countries causes a decline in U.S. agricultural exports (Table A3 in the Appendix). The most likely explanation for this case may be that from the perspective of a consumer in these countries, U.S. agricultural products could be inferior (a low propensity to spend additional income on food); thus, foreign import demand for U.S. agricultural goods tends to decrease as foreign income rises (Gehlhar and Dohlman 2007).

Table 4. Estimated Long-Run Coefficients of the Bilateral Trade Balance Model

Country	Exchange Rate	U.S. Income	Foreign Income	Constant
Canada	0.49 (4.52)**	-2.78 (-2.71)**	1.75 (2.76)**	3.95 (1.31)
Mexico	1.17 (3.46)**	1.94 (2.47)**	-2.61(1.81)*	-6.08 (-3.96)**
Japan	0.51 (2.01)**	1.17 (2.87)**	-0.50 (-0.14)	-8.69 (-2.21)**
Netherlands	2.93 (2.56)**	8.13 (1.72)*	-8.89 (-1.66)*	1.89 (0.13)
Italy	0.85 (2.48)**	6.77 (3.54)**	-7.08 (-1.90)*	2.77 (0.33)
Korea	0.69 (1.76)*	-4.23 (-1.22)	-3.62 (-2.98)**	23.32 (1.56)
Australia	2.68 (3.77)**	-7.41 (-2.06)**	5.61 (2.19)**	8.91 (1.47)
Indonesia	2.26 (2.39)**	-11.74 (-2.82)**	9.70 (4.06)**	-11.85 (-0.65)
France	0.70 (2.70)**	4.54 (3.11)**	-3.45 (-2.53)**	-3.84 (-0.99)
Ireland	4.53 (4.44)**	-6.61 (-1.54)	9.06 (2.04)**	34.34 (1.17)
Germany	0.57 (1.67)*	1.59 (1.33)	-0.64 (-0.27)	-4.60 (-0.78)
Thailand	0.88 (2.15)**	1.91 (1.84)*	-0.83 (1.17)	-1.41 (-0.95)
Spain	1.27 (3.19)**	-8.51 (-2.47)**	9.66 (3.22)**	-5.77 (-2.70)**
New Zealand	1.89 (1.78)*	1.55 (0.15)	20.42 (1.54)	-90.84 (-1.11)
UK	0.79 (2.67)**	7.37 (2.91)**	-2.01 (-0.87)	-23.37 (-2.51)**

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics.

significant factor influencing U.S. agricultural imports in both the short and long run, while the exchange rate is found to be a major factor affecting U.S. agricultural imports only in the short run (Tables A2 and A3 in the Appendix). As a result, our findings can be summarized as follows: (i) in the long run, U.S. agricultural exports depend mainly on both the bilateral exchange rate and foreign income, whereas U.S. agricultural imports are driven largely by U.S. income growth and less by exchange rate changes,[10] and (ii) in the short run, U.S. agricultural exports and imports are responsive to both the bilateral exchange rate and income in the United States and its trading partners.

Concluding Remarks

This study examines the short-run (i.e., J-curve) and long-run effects of exchange rate changes on agricultural trade between the United States and its 15 major trading partners in the framework of the ARDL approach. Results show that there is no evidence of the J-curve effect for U.S. agricultural trade with its major trading partners. The finding further suggests that the fluctuation of agricultural trade surplus during the period 2002–2007 cannot be explained by the J-curve effect. We also find that, although the short-run responses of the trade balance in agricultural goods to the U.S. dollar depreciation do not follow any consistent pattern, the long-run effects support the positive long-run relationship between the exchange rate and the trade balance. The finding thus explains why the econometric model should integrate the short-run dynamics with the long-run equilibrium.

[10] This finding is in line with the general conclusions of Gehlhar and Dohlman (2007) that show that income levels and the rate of economic growth are key determinants of foreign demand for U.S. agricultural exports, but that they have little effect on U.S. agricultural imports.

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Appendix. Bilateral Export (In-payments) and Import (Out-payments) Models

To assess the effect of exchange rates on agricultural exports and imports between the United States and its 15 major trading partners, we adopt the bilateral export and import models developed by Bahmani-Oskooee and Goswami (2004) and Bahmani-Oskooee and Ardalani (2006). In this model, the values of exports (in-payments) and imports (out-payments) between the United States and its trading partners are specified as follows:

$$\text{(A1)} \quad \ln VX_{it} = \beta_0 + \beta_1 \ln Y_{it}^* + \beta_2 \ln ER_{it} + \varepsilon_t$$

$$\text{(A2)}\quad \ln VM_{it} = \gamma_0 + \gamma_1 \ln Y_t + \gamma_2 \ln ER_{it} + \mu_t,$$

where VX_{it} (VM_{it}) is the value of U.S. exports to (imports from) its trading partner i, i = Canada, Mexico, Japan, the Netherlands, Italy, Korea, Australia, Indonesia, France, Ireland, Germany, Thailand, Spain, New Zealand, and the United Kingdom; Y_{it}^* is the real income of trading partner i; Y_t is the real U.S. income; and ER_{it} is the bilateral real exchange rate between the currency of trading partner i and the United States. We then reformulate equations (A1) and (A2) in an error-correction modeling format as follows:

$$\begin{aligned}\text{(A3)}\quad \Delta \ln VX_{it} &= \alpha_0 + \sum_{k=1}^{p} \varepsilon_k \Delta \ln VX_{i,t-k} \\ &+ \sum_{k=1}^{p} \phi_k \Delta \ln Y_{i,t-k}^* + \sum_{k=1}^{p} \varphi_k \Delta \ln ER_{t-k} \\ &+ \lambda_1 \ln VX_{i,t-1} + \lambda_2 \ln Y_{i,t-1}^* \\ &+ \lambda_3 \ln ER_{t-1} + \varepsilon_t\end{aligned}$$

$$\begin{aligned}\text{(A4)}\quad \Delta \ln VM_{it} &= \alpha_1 + \sum_{k=1}^{p} \eta_k \Delta \ln VM_{i,t-k} \\ &+ \sum_{k=1}^{p} \delta_k \Delta \ln Y_{t-k} + \sum_{k=1}^{p} \kappa_k \Delta \ln ER_{t-k} \\ &+ \pi_1 \ln VM_{i,t-1} + \pi_2 \ln Y_{t-1} \\ &+ \pi_3 \ln ER_{t-1} + \mu_t,\end{aligned}$$

where Δ is the difference operator and p is lag order. We use equations (A3) and (A4) to estimate the short- and long-run relationships between the values of U.S. exports and imports and their determinants.

Table A1. Coefficient Estimates of Exchange Rate and Error-Correction Terms of the Export (Inpayments) Model

	Lag Order of Exchange Rate				
Country	0	1	2	3	EC_{t-1}
Canada	-0.15** (-3.24)				-0.23** (-2.07)
Mexico	-0.72** (-2.77)				-0.20** (-2.71)
Japan	-0.19** (-2.44)				-0.25** (-3.09)
Netherlands	-0.25** (-2.48)				-0.36** (-2.70)
Italy	0.41 (0.99)				-0.43** (-2.25)
Korea	-0.72** (-2.72)	0.55* (1.97)			-0.25** (-3.16)
Australia	-0.62** (-2.96)				-0.72** (-3.06)
Indonesia	-0.59** (-2.40)				-0.60** (-4.88)
France	-0.12 (-0.35)	0.73** (2.11)			-0.89** (-4.02)
Ireland	-1.30 (-1.51)	0.96 (1.18)	0.93 (1.08)	2.57** (3.05)	-0.81** (-5.19)
Germany	-0.61** (-3.58)				-0.43** (-2.17)
Thailand	-0.10 (-0.53)				-0.50** (-2.14)
Spain	-1.49** (-2.08)				-0.49** (-2.01)
New Zealand	0.73 (1.31)				-0.75** (-6.32)
UK	-0.04 (-0.28)				-0.23* (-1.92)

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics. EC_{t-1} refers to error-correction term.

Table A2. Estimated Long-Run Coefficients of the Export and Import Models

Export Model			
Country	Exchange Rate	Foreign Income	Constant
Canada	-0.64 (-2.03)**	1.37 (10.77)**	1.57 (2.87)**
Mexico	-1.53 (-1.81)*	1.59 (2.73)**	18.31 (4.82)**
Japan	-0.73 (-2.05)**	2.20 (2.06)**	1.05 (0.29)
Netherlands	-0.69 (-2.14)**	-1.79 (-3.56)**	13.86 (6.00)**
Italy	-0.42 (-3.94)**	-0.94 (-10.24)**	8.86 (1.73)*
Korea	-0.94 (-2.31)**	0.07 (0.28)	12.35 (4.94)**
Australia	-0.86 (-3.73)**	0.50 (3.01)**	2.38 (3.39)**
Indonesia	-0.98 (-2.06)**	-1.24 (-1.58)	19.71 (2.62)**
France	-0.65 (-3.49)**	-1.11 (-4.19)**	9.77 (8.02)**
Ireland	-0.74 (-2.27)**	-0.57 (-1.93)*	7.08 (5.05)**
Germany	-1.42 (-2.18)**	0.14 (0.15)	4.70 (1.05)
Thailand	-0.21 (-1.26)	0.95 (4.47)**	1.46 (1.51)
Spain	-1.08 (-1.94)*	-1.28 (-3.37)**	10.88 (6.15)**
New Zealand	-0.40 (-2.09)**	1.44 (8.66)**	-3.09 (-4.13)**
UK	-0.15 (-0.28)	-0.22 (-0.62)	6.57 (4.22)**
Import Model			
Country	Exchange Rate	U.S. Income	Constant
Canada	1.43 (1.33)	2.33 (4.93)**	-3.80 (-1.68)*
Mexico	0.39 (1.13)	1.49 (6.01)**	13.44 (8.28)**
Japan	-0.18 (-0.46)	1.99 (3.56)**	-3.60 (-2.23)**
Netherlands	0.18 (1.44)	1.21 (15.08)**	0.33 (0.90)
Italy	-0.80 (-1.50)	1.83 (19.53)**	-2.91 (-6.41)**
Korea	0.62 (1.42)	2.17 (9.04)**	-1.63 (-1.08)
Australia	0.37 (0.48)	1.95 (3.43)**	-3.02 (-1.15)
Indonesia	-0.45 (-1.13)	-15.04 (-4.38)**	76.33 (4.37)**
France	-0.04 (-0.20)	1.62 (10.29)**	-1.67 (-2.29)**
Ireland	3.23 (4.22)**	4.88 (3.23)**	-18.15 (-2.60)**
Germany	1.15 (2.88)**	1.65 (4.29)**	-2.32 (-1.31)
Thailand	-1.18 (-4.96)**	1.65 (4.63)**	2.21 (1.68)*
Spain	-0.12 (-0.59)	1.17 (5.99)**	-0.32 (-0.36)
New Zealand	-0.07 (-0.21)	1.61 (3.95)**	-1.63 (-0.90)
UK	0.19 (0.63)	7.74 (2.61)**	-28.03 (-2.23)**

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics.

Table A3. Coefficient Estimates of Exchange Rate and Error-Correction Terms of the Import (Outpayments) Model

Country	Lag Order of Exchange Rate					
	0	1	2	3	4	EC_{t-1}
Canada	0.41* (1.80)					-0.27** (-2.78)
Mexico	0.37 (1.13)					-0.92** (-7.35)
Japan	-0.09 (-0.52)	-0.43** (-2.42)	0.02 (0.12)	-0.37** (-2.18)		-0.22* (-1.88)
Netherlands	-0.11 (-0.10)	-0.18* (-1.71)	-0.05 (-0.42)	-0.23** (-2.27)		-0.49** (-3.41)
Italy	-0.43** (-3.75)	-0.04 (-0.23)	0.06 (0.42)	-0.23* (-1.89)	0.21* (1.73)	-0.57** (-2.95)
Korea	-0.29** (-2.35)					-0.46** (-4.82)
Australia	0.08 (0.51)					-0.23** (-2.00)
Indonesia	-0.59** (-2.39)					-0.60** (-4.87)
France	-0.33* (-1.88)					-0.47** (-2.78)
Ireland	1.86** (3.17)					-0.58** (-4.40)
Germany	0.31** (5.04)					-0.27** (-2.76)
Thailand	-0.52** (-3.28)					-0.44** (-4.02)
Spain	-0.41 (-1.54)	-0.78** (-2.78)				-0.45** (-2.57)
New Zealand	-0.02 (-0.21)					-0.26** (-2.34)
UK	-0.51 (-1.41)					-0.32* (-1.67)

Note: ** and * denote significance at the 5 percent and 10 percent levels, respectively. Parentheses are *t*-statistics. EC_{t-1} refers to error-correction term.