How sensitive Is U.S. agricultural trade to the bilateral exchange rate?: Evidence from bulk and consumer-oriented products

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Abstract

This paper examines the dynamic effects of changes in the bilateral exchange rate on changes in the bilateral trade of bulk and consumer-oriented agricultural products between the U.S. and its 10 major trading partners. We find that, for consumer-oriented products, U.S. exports are highly sensitive to the bilateral exchange rate and foreign income in both the short- and long-run, while U.S. imports are mostly responsive to the U.S. domestic income. For bulk products, on the other hand, U.S. exports and imports are driven largely by the income of the U.S. and its trading partners and less by exchange rate changes in both the short- and long-run.

Key Words: Agricultural exports; agricultural imports; autoregressive distributed lag approach to cointegration; bulk; consumer-oriented; exchange rate

INTRODUCTION

Analyzing the effect of currency devaluation on the trade balance has long been an active field of empirical research in international economics. Given modeling approach and data uses, these studies generally can be classified into three groups. The first group includes early studies that have adopted aggregate trade data in the framework of a two-country model ─ between a country and the rest of the world ─ in examining the exchange rate effect on the trade balance (e.g., Felmingham 1988, Mahdavi and Sohrabian 1993, Guptar-Kapoor and Ramakrishnan 1999). For example, Guptar-Kapoor and Ramakrishnan (1999) use data on trade flows between Japan and the rest of the world to investigate the effects of depreciation of the Japanese yen on the trade balance.

The second group argues that, since a country's trade balance could improve with one trading partner while simultaneously deteriorating with another, the empirical findings of the first group could suffer from aggregation bias of data.[1] This group thus adopts bilateral trade data between a country and its major trading partners to tackle the issue (e.g., Wilson 2001, Arora et al. 2003, Bahmani-Oskooee and Ratha 2004). For example, Bahmani-Oskooee and Ratha (2004) use the bilateral trade balance between the U.S. and thirteen developing countries to examine the dynamic effects of depreciation of the U.S. dollar on the trade balance.

More recently, a new body of literature has been emerging, which argues that the second group may also suffer from aggregation bias of data due to the fact that a country tends to export/import different commodities to/from different trading partners (e.g.,

[1] Additionally, because studies in the first group generally rely on trade data between a country and the rest of the world, instead of at a bilateral level, they need to construct the weighted averages of variables (i.e., exchange rates) used for their analyses. The data compilations could suppress the actual variations of variables taking places at the bilateral level, which may result in aggregation bias as well.

Bahmani-Oskooee and Ardalani 2006, Bahmani-Oskooee and Wang 2007, Bahmani-Oskooee and Bolhasani 2008). Further, this group claims that by dealing with exports and imports as a single variable in their trade balance models, studies both in the first and second groups are not able to directly detect what variables (i.e., exchange rates) are affecting exports and/or imports and by how much, thereby providing misleading results.[2] Accordingly, the third group uses industry/commodity level data (e.g., agriculture, non-agriculture, manufacturing, etc) on bilateral basis and at the same time analyzes exports and imports separately in order to measure the effects of exchange rate fluctuations on the trade balance accurately.[3] For example, Bahmani-Oskooee and Bolhasani (2008) use exports and imports data from 152 commodities to estimate the impact of real depreciation on the bilateral trade balance between Canada and its major trading partners.

Until recently, on the other hand, agricultural economists have typically relied on aggregate trade data in examining the effects of exchange rate changes on the agricultural trade balance (e.g., Carter and Pick 1989, Doroodian et al. 1999, Baek and Koo 2007).[4] For example, Carter and Pick (1989) and Doroodian et al. (1999) analyze the effects of exchange rate change on the agricultural trade balance between the U.S. and the rest of the world. The former finds that market factors other than exchange rate fluctuations are the primary determinants of the U.S. agricultural trade. The latter suggests that depreciation of the dollar has a significant effect on the U.S. agricultural trade balance. Accordingly, relatively limited efforts have been made to investigate the impact of

[2] Studies in the first two groups generally define the trade balance as (1) the excess of value of exports over that of imports or (2) the ratio of value of exports to value of imports.

[3] For example, Bahmani-Oskooee and Ardalani (2006) show that U.S. exports are more responsive to changes in the value of the U.S. dollar, while U.S. imports are driven largely by U.S. income growth and less by exchange rate changes. Hence, they conclude that analyzing imports and exports separately is indeed desirable to draw more robust findings.

[4] In other words, all studies in the agricultural trade literature constitute the first group as classified earlier.

exchange rate fluctuation on the agricultural trade balance at the bilateral commodity level. In other words, no study has dealt with disaggregate bilateral trade data in examining a direct link between the agricultural trade balance and exchange rates.[5]

In this paper, therefore, we have attempted to expand the literature on agricultural trade by assessing the effects of exchange rate fluctuations on U.S. agricultural exports and imports within the context of disaggregating agricultural product data (e.g., bulk and consumer-oriented products) of bilateral trade — between the U.S. and its major 10 trading partners. Special attention has been paid to assess the characteristics of the short- and long-run dynamics and empirically identify whether or not U.S. exports and imports in agricultural products could benefit from dollar depreciation. In fact, given the continuing decline in the value of the U.S. dollar since 2002, it is very timely and important to explore the issue.[6] For this purpose, we use an autoregressive distributed lag (ARDL) approach to cointegration or an ARDL bound testing approach (referred to here as ARDL model) developed by Pesaran et al. (2001). Because an error-correction model (ECM) can be derived from the ARDL model through a simple linear transformation, this model is widely used to estimate the short- and long-run parameters of the model simultaneously.

The remainder of this paper is organized as follows. The next section briefly describes overview of U.S. agricultural trade over the last two decades. The following section introduces the empirical model associated with the ARDL estimation as well as

[5] In fact, Bahmani-Oskooee and Ardalani (2006) examine the effect of real depreciation of the U.S. dollar on exports and imports of 66 U.S. industries, including agricultural commodities. The focus of the analysis, however, is also on trade flow between the U.S. and the rest of the world, instead of between the U.S. and its major trading partners. Accordingly, their results could also suffer from aggregation bias as they admitted.

[6] During the 2002-2007 period, for example, the value of the U.S. dollar decreased by approximately 30%, 6% and 31% against the Canadian dollar, the Japanese yen and the euro, respectively.

the dataset used for the analysis. The last two sections discuss the empirical results and make some concluding remarks.

OVERVIEW OF U.S. AGRICULTURAL TRADE

Agricultural Trade Balance

In the past two decades, U.S. agricultural exports have often experienced volatile swings, while U.S. agricultural imports have been relatively steady, even becoming increasingly strong in recent years, thereby resulting in fluctuations in the agricultural trade balance (Figure 1). During the early 1990s, for example, U.S. agricultural exports had increased substantially, from $39.5 billion in 1990 to $60.4 billion in 1996. During the same period, on the other hand, U.S. agricultural imports were fairly stable, ranging from $22.9 billion in 1990 to $33.5 billion in 1996. As a result, the agricultural trade surplus had reached a record high of $26.9 billion in 1996. This surplus had begun to decline as a result of the slow growth of U.S. agricultural exports relative to imports since 1996. U.S. agricultural imports, for example, have increased from $36.1 billion in 1997 to $59.3 billion in 2005. Meanwhile, U.S. agricultural exports have fluctuated from a low of $48.4 billion in 1999 to a high of $63.2 billion in 2005. Accordingly, the agricultural trade surplus dipped below $5 billion in 2005. In 2007, however, the trade surplus has begun its long-awaited improvement as U.S. exports rise to a record high of $89.9 billion and U.S. import growth, while still strong, is at its slowest pace in 5 years; hence, the trade surplus rebounded to $18 billion in 2007 and then jumped to $34.8 billion in 2008.

Types of Agricultural Products Traded

The Foreign Agricultural Service's (FAS) BICO data classifies agricultural exports and imports into bulk, intermediate, and consumer-oriented products. Bulk products include grains such as wheat and rice, cotton, tobacco and other bulk commodities. Intermediate products include products such as wheat flour, soybean meal, soybean oil, vegetable oils, live animals, animal fats and other intermediate commodities. Consumer-oriented products include snack foods, red meats, dairy products, processed food and vegetables, nursery products and other processed or ready-to-eat products.

On the export side, bulk products accounted for a larger percentage of U.S. agricultural trade in the mid-1990s (Figure 2). Since 1996, however, the export share for bulk commodities has declined, while the export share of consumer-oriented products has increased. During the period of 2003-07, therefore, exports of these two products have been close to equal. During the same period, China, Japan and Mexico have been the major markets for U.S. exports of bulk products, followed by Taiwan, Korea, Egypt, Canada, Turkey, Indonesia and Colombia (Table 1). Additionally, Canada, Mexico and Japan are the top three importers of consumer-oriented products from the U.S. over the last five years, followed by Korea, Russia, UK and China.

On the import side, on the other hand, consumer-oriented products have accounted for the largest share of agricultural trade over the last two decades (Figure 3). During the period of 2003-07, for example, consumer-oriented products consist of 67.4%, while bulk and intermediate products account for 13.4% and 19.2%, respectively. During the same period, Canada and Mexico have been the two largest exporters of consumer-oriented products to the U.S., followed by Australia, Italy, the Netherlands, France and New Zealand (Table 1). Additionally, Indonesia has been the top exporter of bulk products to the U.S. over the past five years, followed by Canada, Brazil, Thailand and Colombia.

THE MODELS

To construct the model, we define the trade balance as the difference between the value of exports (inpayments) and the value of imports (outpayments) as follows:

$$TB = XQ \times P_x - MQ \times P_m \tag{1}$$

where TB is the trade balance; XQ (MQ) is the volume of exports (imports); and P_x (P_m) is the domestic price of exports (imports). In examining the exchange rate effects on trade flows, most previous studies have typically related the volume of exports and imports to a measure of relative prices and income, and have estimated demand elasticities of (aggregate) imports and exports to determine whether the Marshall-Lerner condition holds.[7] One major limitation of this approach, however, is the assumption of perfectly elastic supply of imports and exports. Further, when analyzing bilateral trade flows, the traditional approach is no longer suitable because of the unavailability of import and export prices of different commodities at a bilateral level.[8]

To overcome these shortcomings, Bahmani-Oskooee and Goswami (2004) and Bahmani-Oskooee and Ardalani (2006) have attempted to directly link bilateral export ($XQ \times P_x$) and import values ($MQ \times P_m$) and changes in exchange rates and (domestic

[7] The Marshall-Lerner condition states that as long as the sum of domestic and foreign price elasticities of demand (in absolute value) exceeds unity, currency devaluation will increase a country's inpayments and decrease outpayments, thereby improving the trade balance.

[8] In some instances, of course, bilateral export and import price indices at commodity levels may be available. One major shortcoming of such indices, however, is the assumption that exporters (importers) charge (pay) the same prices at domestic and foreign countries (all exporting countries) for each exported (imported) commodity. As Cushman (1987 and 1990) noted, this assumption raises the issue of specification error due mainly to an inadequate specification of data elements.

and foreign) income.[9] This direct method allows easily determining whether currency devaluation has favorable effects on a country's export and import values. For each product group j, therefore, we formulate the bilateral inpayments and outpayments models between the U.S. and trading partner i in a log linear form as follows:

$$\ln(VX_{ijt}) = a_0 + a_1 \ln(Y_{it}^*) + a_2 \ln(ER_{it}) + \varepsilon_t \tag{2}$$

$$\ln(VM_{ijt}) = b_0 + b_1 \ln(Y_t) + b_2 \ln(ER_{it}) + \mu_t \tag{3}$$

where VX_{ijt} (VM_{ijt}) is the value of product j's U.S. exports to (imports from) its trading partner i; 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 U.S. In equations (2) and (3), to the extent that an increase in U.S. (trading partner) income results in an increase in U.S. agricultural imports (exports), it is expected that $a_1 > 0$ and $b_1 > 0$. Additionally, it is expected that $a_2 < 0$ and $b_2 > 0$, since the depreciation of the U.S. dollar causes a decrease (increase) in imports (exports) of agricultural goods through a rise (decline) in import (export) prices.[10]

We employ the ARDL approach to cointegration developed by Pesaran et al. (2001) to examine the dynamic relationship between changes in exchange rates and changes in U.S. agricultural trade. The main advantage of this testing and estimation approach is that it can be applied irrespective of whether the regressors are $I(0)$ or $I(1)$, and this avoids the well-known pre-testing problems associated with standard cointegration techniques (e.g., Engle and Granger 1987, Johansen 1995) that requires the

[9] Their empirical models have initially relied on a theoretical framework developed by Haynes et al. (1986) and Cushman (1987 and 1990).
[10] These expected signs are based on the definition of ER, which is defined here as number of units of foreign currencies per U.S. dollar.

classification of the variables into stationary or non-stationary. In addition, the ARDL model takes sufficient numbers of lags to capture the data generating process in a dynamic framework of a general-to-specific modeling; it is thus more robust and performs better than other cointegration tests, even with small or finite sample size (Pesaran and Shin 1999).

Following Pesaran et al. (2001), the error-correction representations of the ARDL specification model for equations (2) and (3) are given by

$$\Delta \ln VX_{it} = \alpha_0 + \sum_{k=0}^{p} \varepsilon_k \Delta \ln VX_{i,t-k} + \sum_{k=0}^{p} \phi_k \Delta \ln Y^*_{i,t-k} + \sum_{k=0}^{p} \varphi_k \Delta \ln ER_{t-k} \tag{4}$$

$$+ \delta_1 \ln VX_{i,t-1} + \delta_2 \ln Y^*_{i,t-1} + \delta_3 \ln ER_{t-1} + \varepsilon_t$$

$$\Delta \ln VM_{it} = \alpha_1 + \sum_{k=0}^{p} \eta_k \Delta \ln VM_{i,t-k} + \sum_{k=0}^{p} \delta_k \Delta \ln Y_{t-k} + \sum_{k=0}^{p} \kappa_k \Delta \ln ER_{t-k} \tag{5}$$

$$+ \eta_1 \ln VM_{i,t-1} + \eta_2 \ln Y_{t-1} + \eta_3 \ln ER_{t-1} + \mu_t$$

where Δ is the difference operator; and p is lag order. Equations (4) and (5) are different from a standard VAR model in that a linear combination of lagged-level variables are used as proxy for lagged error terms.

We begin by testing for the presence of the long-run relationship between variables. For this, we use the F-test to determine the joint significance of lagged levels of the variables in equations (4) and (5). Pesaran et al. (2001) provide two sets of asymptotic critical values for the F-test. One set assumes that all the variables are $I(0)$ and another assumes they all are $I(1)$. For this purpose, the null hypotheses of the non-existence of the long-run relationship are $H_0: \delta_1 = \delta_2 = \delta_3 = 0$ against $H_1: \delta_1 \neq \delta_2 \neq \delta_3 \neq 0$ in equation (4), and $H_0: \eta_1 = \eta_2 = \eta_3 = 0$ against

$H_1 : \eta_1 \neq \eta_2 \neq \eta_3 \neq 0$ in equation (5). The null hypotheses of no cointegration among the variables in both equations can be rejected if the computed F-statistic falls outside the upper bound of the critical values. Conversely, if the computed F-statistic falls below the lower bound of the critical values, the null cannot be rejected. Finally, if the F-statistic falls inside the two bounds, the inference is inconclusive and knowledge of the order of the integration of the underlying variables is necessary to make a decision on long-run relationships (Pesaran et al. 2001).

DATA AND EMPIRICAL PROCEDURE

Data

To analyze the dynamic effects of exchange rate changes on agricultural commodity trade between the U.S. and its major trading partners, quarterly export and import data between the first quarter of 1989 and the fourth quarter of 2007 (1989:q1-2007:q4) are collected. For this purpose, based on the BICO reports classified by the Foreign Agricultural Service (FAS) in the U.S. Department of Agriculture (USDA), we divide U.S. agricultural trade into two groups such as bulk and consumer-oriented products.[11] Based on the average 2003-07 trade share of each trading partner for bulk and consumer-oriented products, we then identify the 10 largest trading partners of the U.S. for each product. For example, Canada, Mexico, Japan, Australia, Netherlands, Italy, France, New Zealand, Korea and Germany are chosen for the 10 major trading partners for consumer-

[11] The BICO (**B**ulk, **I**ntermediate, and **C**onsumer-**O**riented) reports are derived from the Harmonized Tariff System (HTS) to the 6-digit level for generalized categories. Additionally, we emphasize here that, because of the relatively small share of agricultural trade, intermediate products (19.8% of exports and 19.2% of imports) are not included in our study. Hence, bulk and consumer-oriented products account for 81.2% of exports and 80.8% of imports, respectively.

oriented products. Likewise, Japan, Mexico, Indonesia, Canada, Korea, Thailand, Turkey, Germany, Netherlands and Spain are selected for bulk products (Table 1).[12]

The total values of exports and imports for bulk and consumer-oriented products between the U.S. and its 10 major trading partners are collected from the USDA's FAS Online. The income of the U.S. and its trading partners is measured as real gross domestic product (GDP) index (2000=100) and is taken from the International Financial Statistics (IFS) published by the International Monetary Fund (IMF). The real exchange rates between the U.S. dollar and the currencies of its 10 major trading partners are collected from the USDA's Economic Research Service (ERS). Since the exchange rate is expressed as the number of trading partner's currency per unit of the U.S. dollar, a decline in exchange rate indicates a real depreciation of the U.S. dollar. The GDP deflator (2000=100) obtained from the IFS is used to derive real values of exports and imports of bulk and consumer-oriented products. Finally, the data are converted to natural logarithms and used throughout.

Empirical Procedure

As noted earlier, the ARDL modeling starts with testing the existence of the long-run relationship between the variables in equations (4) and (5) using the F-values. The outcome of the bounds tests critically depends on the choice of the lag order (p); it is

[12] It should be noted that, based on the availability of data, we select these 10 countries out of 20 top trading partners for bulk and consumer-oriented products (Table 1). For example, because of unavailability of GDP data in the IFS, two Asian countries (China and Philippines) and six South/Central American countries (Colombia, Brazil, Chile, Dominican Republic, Cote d'ivoire and Costa Rica) are excluded from the analysis. Additionally, due to the unavailability of recent economic data in the IFS, Taiwan is also excluded from this study. Finally, because of limited availability of GDP data in the IFS, the data for Thailand contains 60 observations for 1993:q1-2007:q4 and Indonesia includes 44 observations for 1997:q1-2007:q4, respectively. Additionally, since U.S.-Netherlands trade for both bulk and consumer-oriented products nearly disappear in 2006 and 2007, the data for Netherlands covers 68 observations for 1989:q1-2005:q4 for both cases

thus critical that the lag order of the underlying VAR is selected appropriately. To this end, we adopt the Akaike Information Criterion (AIC) and Lagrange multiplier (LM) statistics for testing the hypothesis of no serial correlation to select the optimum number of lags.[13] The results show that, of the 40 cases, the F-statistics are found to lie outside the upper level of the 10% critical bounds for 32 cases (Table 2). With $p = 2$ for the U.S. exports (imports) of consumer-oriented products to (from) Canada, for example, the F-statistic is 5.04 (5.34), which lies outside the upper level of the 10% critical bounds.[14] As such, this result supports the existence of cointegrated export (import) equation when using $p = 2$. With the eight cases, on the other hand, the computed F-statistics fall within the 10% bound. For the U.S. exports of consumer-oriented products to Japan and Korea, for example, the test statistics are 3.54 and 3.37, respectively, which falls inside the 10% bound. In this case, following Kremers et al. (1992) and Banerjee et al. (1998), the error-correction term in the ARDL model can be used to determine the existence of the long-run relationship. Hence, if a negative and significant lagged error-correction term is obtained, the variables are said to be cointegrated.

With the existence of the level relationship identified, we then use the selected ARDL model to estimate the long-run coefficients and error-correction model. More specifically, the long-run model can be estimated from the reduced-form solution of equations (4) and (5), when the first-differenced variables jointly equal zero. The error-correction model is estimated by the ARDL approach. For this purpose, a general-to-

[13] As Pesaran et al. (2001) note: "there is a delicate balance between choosing p sufficiently large to mitigate the residual serial correlation problem and, at the same time, sufficiently small so that the conditional error-correction model in equations (4) and (5) are not unduly over-parameterized, particularly in view of the limited time series data which are available (p. 308)."

[14] The upper bound critical value for F-statistic with unrestricted intercept and restricted trend at the 10% significance level is 4.02 and the lower bound critical value is 3.38. These values are obtained from Pesaran et al. (2001).

specific modeling approach guided by the AIC is used to select the optimal lag structure of the ARDL specification. With the U.S. exports of consumer-oriented products to Canada ($p = 2$), for example, the estimated orders of an $ARDL(p, p_1, p_2)$ model in the three variables (VX_t, Y_t^*, ER_t) are selected by a general-to-specific search, spanned by lag length $p = 0,1,2$ using the AIC criterion, which results in the choice of an $ARDL(2,0,0)$ specification.

EMPIRICAL RESULTS

In this section we divide our findings into the short- and long-run analyses. The long-run results of export and import functions for consumer-oriented and bulk products are summarized in Table 3, while the short-run results are summarized in Table 4.

Results of Long-Run Analysis

The results of the long-run coefficient estimates of export function for consumer-oriented products show that the exchange rate is statistically significant at least at the 10% level for all cases except Korea, and has the expected negative sign (Table 3). This implies that, in the long-run, a decrease in the value of the U.S. dollar (depreciation) causes an increase in U.S. exports of consumer-oriented products through a decline in export prices. Additionally, of the 9 cases in which the real foreign income is statistically significant, 7 cases show a positive long-run relationship between U.S. exports and real foreign income. This suggests that a rise in real income of those countries boosts purchasing power of foreign consumers and leads to growth in foreign demand for U.S. exports of consumer-oriented products. For the remaining 2 cases (Mexico and France), on the other hand, U.S.

exports of consumer-oriented products have a negative long-run relationship with the foreign real income, implying that an increase in real income of those countries causes a decline in U.S. agricultural exports. The most likely explanation for this finding is that, given the definition of imports (consumption minus production), an increase in foreign income could lead to an increase in the foreign production of import-substitute commodities faster than foreign consumption, which could lead to reduced foreign imports (Magee 1973, Bahmani-Oskooee and Ratha 2004).

For bulk products, on the other hand, the results of the long-run coefficient estimates of export function show that the exchange rate is not statistically significant even at the 10% level in most cases (7 out of 10 countries), indicating that the bilateral exchange rate plays little role in determining U.S. exports of bulk commodities in the long-run (Table 3). One possible explanation for this finding is that, since bulk commodities generally consist of essential goods such as wheat, corn, rice and soybeans, foreign demand for U.S. exports of those products may not respond to exchange-rate driven price changes significantly; in other words, the price elasticity of import demand in the major U.S. export markets is inelastic for bulk products. The coefficients of the real foreign income, by contrast, are statistically significant at the 10% level in the majority of cases (8 out of 10 countries). For example, U.S. exports of bulk products have a positive long-run relationship with the real income of Japan, Canada, Thailand and Turkey, and have a negative relationship with the real income of Mexico and three EU countries (Germany, Netherlands and Spain). As seen in U.S. exports of consumer-oriented products, therefore, this finding shows that real foreign income is an important factor in affecting U.S. exports of bulk commodities.

The results of the long-run coefficient estimates of import function for consumer-oriented products show that the exchange rates are not statistically significant even at the 10% level for all cases except Canada, indicating that the bilateral exchange rate has little effect on U.S. imports of consumer-oriented products (Table 3). The finding that more imports of consumer-oriented products are weakly responsive to the exchange rate than exports of those products could be due to the fact that foreign exporters tend to squeeze their profit margins to maintain their share of the U.S. market as the value of the dollar decreases. Another possible explanation for this is that the structural changes in U.S. diet toward high-quality products over the last two decades may have been an important factor in driving U.S. imports higher; under this circumstance, exporters may tend to supply more premium and high-quality products to the U.S. market as exchange rates fluctuate and profit margins shift. Notice that the exchange rate has a significant effect on U.S. imports of consumer-oriented products from Canada, which accounts for approximately 20% of U.S. imports of those products (Table 1). The coefficients of the real U.S. income, by contrast, are statistically significant at the 5% level for all cases, suggesting that, in the long-run, a rise in real U.S. income boosts American purchasing power, thereby increasing U.S. imports of consumer-oriented products. Similarly, for bulk commodities, in the majority of cases, the real U.S. income carries a significant coefficient, while the bilateral exchange rate does not.

Results of Short-Run Analysis

Now, we turn our attention to the short-run dynamics, which is identified by coefficient estimates of first differenced variables in equations (4) and (5). The results of export function show that, for consumer-oriented products, the coefficients of the exchange rate are statistically significant at least at the 10% level for all cases except New Zealand, indicating that, in the short-run, the bilateral exchange rate is an important determinant of U.S. exports of consumer-oriented products (Table 4). For bulk products, by contrast, the exchange rate carries a significant coefficient only in 3 cases, implying that, by and large, the bilateral exchange rate has little effect on U.S. exports of bulk products. Additionally, the coefficients of the real foreign income for both bulk and consumer-oriented products are statistically significant at least at the 10% level for almost all cases (18 out of 20 cases), suggesting that foreign income plays the dominant role in determining U.S. exports of bulk and consumer-oriented products.[15]

The results of import function show that the real U.S. income has a significant short-run effect on U.S. imports of both bulk and consumer-oriented products for all cases except Italy (in consumer-oriented products). However, exchange rate generally carries an insignificant coefficient in both bulk and consumer-oriented products, showing lack of significant relation between the value of the dollar and the value of imports. From these, therefore, the results of short-run analysis seem to be consistent with those of long-run analysis; in the both short- and long-run, income effects hold for both U.S. exports and imports of bulk and consumer-oriented products, while exchange rate effects hold only for U.S. exports of consumer-oriented products.

It should be noted that the coefficients of the error-correction terms are negative and statistically significant at least at the 10% level for all cases (Table 4). A highly significant error-correction term is further proof of the existence of a stable long-run

[15] The real income of Japan and Netherlands is insignificant for consumer-oriented and bulk products, respectively. To save space, the short-run coefficients of the real income of U.S. and its major partners are not reported here, but can be obtained from the authors upon request.

relationship among variables (Kremers et al. 1992, Banerjee et al. 1998). Thus, the findings justify the ARDL modeling of the bilateral export and import models for bulk and consumer-oriented products in which the results of the F -statistics are inconclusive (Table 2).

Finally, we use the cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests to the residuals of error-correction models (equations (4)-(5)) in order to ensure that estimated coefficients are stable over time.[16] For stability of all estimated coefficients, the plot of these two statistics should stay within the 5% significance level. For U.S. exports and imports of consumer-oriented products to/from Canada, for example, since the plot of these two statistics stays within the critical bounds, the estimated coefficients are indeed stable over time (Figure 4). The overall results of stability test suggest that, in the majority of the models, the estimated coefficients are generally stable over the sample period (Table 5).

SUMMARY AND CONCLUSIONS

This study examines the dynamic interaction between changes in the exchange rate and changes in U.S. agricultural trade. Given the continuing decline in the value of the U.S. dollar over the last seven years, it is very timely and important to explore the linkage. Although the literature on the relationship between the exchange rate and U.S. agricultural trade exists, relatively little attention has been paid to the direct effects of exchange rates on U.S. agricultural trade at the bilateral commodity level. Hence, this study has attempted to quantify the effect of exchange rate fluctuations on agricultural

[16] It should be pointed out that these tests are known to have low power and could miss important breaks. However, the diagnostic tests indicate no serious problems with serial correlation, heteroskedasticity, and normality; overall, therefore, the ARDL models presented are well defined and provide sound findings.

trade in the context of disaggregating agricultural product data of bilateral trade – agricultural trade in bulk and consumer-oriented products between the U.S. and its 10 major trading partners. For this purpose, we use the ARDL approach and consider separating the analysis of exports and imports in order to measure the effects of exchange rate changes on the agricultural trade accurately. We find that, for U.S. bilateral trade in consumer-oriented products, exports are highly sensitive to the bilateral exchange rate and foreign income in both the short- and long-run, while imports are mostly responsive to the U.S. domestic income. For bulk products, on the other hand, U.S. exports and imports are driven largely by the income of the U.S. and its trading partners and less by exchange rate changes in both the short- and long-run.

REFERENCES

Arora, S., M. Bahmani-Oskooee, and G.G. Goswami. 2003. Bilateral J-curve between India and her trading partners. Applied Economics 35: 1037-1041.

Baek, J. and W.W. Koo. 2007. Dynamic interrelationship between the U.S. agricultural trade balance and the macroeconomy. Journal of Agricultural and Applied Economics 39(3): 457-470.

Bahmani-Oskooee, M., and M. Bolhasani. 2008. The J-curve: evidence from commodity trade between Canada and the US. Journal of Economics and Finance 32: 207-225.

Bahmani-Oskooee, M., and Y. Wang. 2007. United States-China trade at the commodity level and the yuan-dollar exchange rate. Contemporary Economic Policy 25(3): 341-361.

Bahmani-Oskooee, M., and Z. Ardalani. 2006. Exchange rate sensitivity of U.S. trade flows: evidence from industry data. Southern Economic Journal 72(3): 542-559.

Banerjee, A., J.J. Dolado, and R. Mestre. 1998. Error-correction mechanism tests for cointegration in a single-equation framework. Journal of Time Series Analysis 19: 267-284.

Carter, C.A., and D.H. Pick. 1989. The J-curve effect and the U.S. agricultural trade balance. American Journal of Agricultural Economics 71: 712-720.

Cushman, D.O. 1987. U.S. bilateral trade balances and the dollar. Economics Letter 24: 363-367.

Cushman, D.O. 1990. U.S. bilateral trade equations: forecasts and structural stability. Applied Economics 22: 1093-1102.

Doroodian, K., C. Jung, and R. Boyd. 1999. The J-curve effect and U.S. agricultural and industrial trade. Applied Economics 31: 687-695.

Engle, R.F., and C.W.J. Granger. 1987. Cointegration and error correction representation: estimation and testing. Econometrica 55: 251-276.

Felmingham, B.S. 1988. Where is the Australian J-curve. Bulletin of Economic Research 40: 43-56.

Guptar-Kapoor, A., and U. Ramakrishnan. 1999. Is there a J-curve: a new estimation for Japan. International Economic Journal 13(4): 71-79.

Haynes, S.E., M. M. Hutchison, and R.F. Mikesell. 1986. U.S.-Japanese bilateral trade and the yen-dollar exchange rate: an empirical analysis. Southern Economic Journal 65: 626-632.

Johansen, S. 1995. Likelihood-based inference in cointegrated vector autoregressive models. Oxford: Oxford University Press.

Kremers, J.J.M, N.R. Ericson, and J.J. Dolado. 1992. The Power of cointegration tests. Oxford Bulletin of Economics and Statistics 54(3): 325-348.

Magee, S.P. 1973. Currency contracts, pass-through, and the J-curve. Brookings Papers of Economic Activity 1: 303-323.

Mahdavi, S., and A. Sohrabian. 1993. The exchange value of the dollar and the US trade balance: an empirical investigation based on cointegration and Granger causality tests. Quarterly Review of Economics and Finance 33(4): 343-358.

Pesaran, M.H., and Y. Shin. 1999. An autoregressive distributed lag modeling approach to cointegration analysis. Chapter 11 in Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium, Strom S (ed.). Cambridge University Press: Cambridge, UK.

Pesaran, M.H., Y. Shin, and R.J. Smith. 2001. Bounds testing approaches to the analysis of level relationships. Journal of Applied Econometrics 16: 289-326.

Wilson, P. 2001. Exchange rates and the trade balance for dynamic Asian economies: does the J-curve exist for Singapore, Malaysia and Korea? Open Economies Review 12: 389-413.

Table 1. U.S. exports and imports of bulk and consumer-oriented agricultural products, 2003-07 Average (million $)

Consumer-Oriented				**Bulk**			
Country	Exports	Imports	Total	Country	Exports	Imports	Total
Canada	8,357	8,467	16,824	China	4,433	142	4,575
Mexico	4,134	7,122	11,256	Japan	3,961	14	3,975
Japan	3,566	316	3,882	Mexico	3,476	269	3,746
Australia	308	2,250	2,558	Indonesia	739	1,349	2,089
Netherlands	577	1,650	2,227	Canada	780	938	1,718
Italy	229	1,943	2,172	Taiwan	1,462	16	1,478
China	653	1,288	1,941	Korea	1,126	4	1,130
France	200	1,619	1,819	Colombia	542	558	1,100
New Zealand	97	1,524	1,621	Thailand	393	581	974
Chile	52	1,385	1,438	Egypt	916	19	935
Korea	999	190	1,189	Turkey	743	156	899
Germany	473	658	1,131	Brazil	116	757	872
Spain	325	779	1,104	Germany	478	146	624
UK	709	376	1,086	Guatemala	213	340	553
Brazil	88	928	1,017	Nigeria	466	25	490
Costa Rica	56	830	886	Philippines	376	75	451
Colombia	81	722	803	Dominican	298	138	436
Russia	765	16	781	Cote d'ivoire	11	412	422
Taiwan	584	130	714	Costa Rica	220	171	391
Thailand	143	561	704	Netherlands	376	5	381
Hong Kong	622	58	679	Spain	365	7	372
Sub-Total	23,019	32,813	55,832	Sub-Total	21,489	6,122	27,612
Total	27,868	40,172	68,040	Total	27,422	7,984	35,406

Table 2. Results of F-test for cointegration among variables

		Exports				Imports			
	Country	Lag	$\chi^2(1)$	F - statistics	Decision	Lag	$\chi^2(1)$	F - statistics	Decision
Consumer-Oriented	Canada	2	0.04	5.04	O	2	0.41	5.34	O
	Mexico	1	1.86	4.23	O	3	4.63	6.26	O
	Japan	1	3.42	3.54	Δ	2	0.01	4.92	O
	Australia	3	4.06	7.32	O	3	1.13	4.58	O
	Netherlands	2	0.35	13.12	O	2	0.05	4.92	O
	Italy	2	0.79	4.20	O	8	1.53	3.80	Δ
	France	1	1.09	9.29	O	1	1.39	13.15	O
	New Zealand	6	0.12	5.22	O	4	7.13	3.32	Δ
	Korea	2	0.78	3.37	Δ	2	0.56	3.82	Δ
	Germany	6	0.03	4.50	O	1	2.58	7.93	O
Bulk	Japan	1	1.13	10.94	O	1	3.42	3.21	Δ
	Mexico	2	0.01	8.71	O	1	2.75	14.04	O
	Indonesia	1	1.68	16.53	O	4	0.23	6.71	O
	Canada	1	0.18	4.03	O	1	3.91	3.63	Δ
	Korea	2	0.06	3.77	Δ	1	0.10	8.96	O
	Thailand	3	0.87	7.31	O	1	0.02	5.43	O
	Turkey	2	0.34	12.80	O	3	0.21	4.25	O
	Germany	4	0.15	8.67	O	1	0.01	6.79	O
	Netherlands	2	0.01	4.14	O	1	0.12	7.86	O
	Spain	6	0.28	4.12	O	1	2.94	7.63	O

Note: For exports (imports) the first lag is for $\Delta \ln VX$ ($\Delta \ln VX$), the second is for $\Delta \ln Y^*$ ($\Delta \ln Y$), and the last is for $\Delta \ln ER$. A lag order is selected based on Akaike Information Criterion (AIC). $\chi^2(1)$ are Lagrange Multiplier (LM) statistics for testing no serial correlation against order 1. The upper bound critical value for F-statistic with unrestricted intercept and restricted trend at the 10% significance level is 4.02 and the lower bound critical value is 3.38. These values are obtained from Pesaran et al. (2001). O and Δ represent cointegration and inconclusive, respectively.

Table 3. Estimated long-run coefficients of export and import functions

		Exports			Imports		
	Country	Exchange rate	Foreign income	Constant	Exchange rate	U.S. income	Constant
Consumer-Oriented	Canada	-0.57 (-3.39)**	1.40 (14.09)**	1.05 (2.40)**	0.66 (5.71)**	3.01 (32.87)**	-6.72 (-16.32)**
	Mexico	-2.86 (-4.69)**	-1.43 (-3.41)**	19.69 (7.04)**	-0.24 (-0.64)	-7.69 (-2.96)**	40.22 (3.65)**
	Japan	-0.69 (-1.71)*	-2.06 (-0.96)	19.39 (2.09)**	0.12 (0.33)	1.65 (4.56)**	-3.91 (-3.01)**
	Australia	-2.62 (-1.96)**	2.55 (3.42)**	-6.86 (-2.30)**	1.65 (1.25)	2.97 (2.83)**	-8.39 (-1.77)*
	Netherlands	-0.32 (-1.92)*	1.66 (9.56)**	-2.82 (-3.54)**	0.13 (0.69)	1.20 (6.98)**	-0.02 (-0.02)
	Italy	-1.68 (-3.93)**	2.63 (3.47)**	-6.75 (-1.94)*	0.21 (0.59)	1.51 (12.84)**	-3.59 (-15.24)**
	France	-1.14 (-4.29)**	-0.82 (-2.28)**	7.40 (4.49)**	-0.06 (-0.32)	1.87 (11.99)**	-2.33 (-3.25)**
	New Zealand	-0.53 (-1.65)*	2.37 (8.66)**	-8.01 (-6.52)**	0.10 (0.23)	1.90 (3.62)**	-3.09 (-1.33)
	Korea	0.21 (0.42)	4.66 (3.20)**	-15.11 (-2.02)**	-0.67 (-1.41)	2.23 (4.73)**	-2.16 (-0.72)
	Germany	-1.55 (-3.53)**	1.39 (1.69)*	-2.03 (-0.53)	-0.18 (-1.03)	0.79 (5.51)**	1.29 (1.93)*
Bulk	Japan	-0.73 (-1.77)*	7.10 (3.22)**	-21.24 (-2.25)**	-0.54 (-0.44)	4.01 (3.20)**	-15.05 (-3.49)**
	Mexico	-0.11 (-0.25)	-1.83 (-5.42)**	15.19 (7.10)**	1.64 (2.05)**	-4.53 (-7.58)**	21.05 (5.68)**
	Indonesia	-0.93 (-2.82)**	-1.52 (-2.51)**	20.53 (3.81)**	-0.49 (-0.69)	-19.84 (-4.19)**	97.75 (4.10)**
	Canada	0.61 (1.35)	1.27 (4.31)**	-0.77 (-0.58)	-0.83 (-0.76)	1.30 (1.72)*	-0.74 (-0.22)
	Korea	-1.89 (-2.65)**	0.04 (0.08)	18.55 (4.16)**	-0.64 (-1.31)	0.09 (0.18)	4.00 (1.27)
	Thailand	-0.32 (-1.11)	0.70 (1.09)*	2.89 (1.72)*	-1.58 (-1.35)	1.46 (2.82)**	3.74 (2.20)**
	Turkey	0.18 (0.22)	1.45 (2.52)**	4.88 (1.72)*	1.05 (0.95)	2.20 (16.03)**	15.77 (4.98)**
	Germany	-0.83 (-0.78)	-2.17 (1.48)	15.18 (2.27)**	-0.89 (-1.62)	0.67 (22.40)**	-5.57 (-3.67)**
	Netherlands	0.76 (0.77)	-5.67 (-5.54)**	32.50 (6.75)**	-1.01 (-1.36)	-1.29 (-1.89)*	5.91 (1.89)*
	Spain	-0.56 (-1.33)	-2.33 (-5.88)**	15.01 (8.21)**	-0.67 (-1.27)	0.38 (0.71)	-1.48 (-0.59)

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

Table 4. Estimated short-run coefficients of export and import functions

		Exports			Imports			
	Country	Lag order of exchange rate			Lag order of exchange rate			
		0	1	ec_{t-1}	0	1	2	ec_{t-1}
Consumer-Oriented	Canada	-0.27		-0.48	0.39			-0.59
		(-3.20)**		(-3.13)**	(4.05)**			(-4.09)**
	Mexico	-0.94		-0.33	-0.13			-0.57
		(-4.57)**		(-4.80)**	(-0.65)			(-2.99)**
	Japan	-0.23		-0.33	-0.21	-0.56		-0.35
		(-1.76)*		(-3.81)**	(-0.82)	(-2.23)**		(-3.43)**
	Australia	-0.40		-0.15	0.23			-0.14
		(-3.27)**		(-1.78)*	(1.49)			(-2.21)**
	Netherlands	-0.21		-0.68	-0.46			-0.35
		(-1.92)*		(-5.16)**	(-2.68)**			(-3.11)**
	Italy	-1.27	0.62	-0.33	-0.65	-0.61		-0.17
		(-3.18)**	(1.54)	(-3.43)**	(-3.42)**	(-2.84)**		(-2.52)**
	France	-0.77		-0.68	-0.03			-0.44
		(-3.59)**		(-5.86)**	(-0.32)			(-4.30)**
	New Zealand	-0.21		-0.40	0.02			-0.25
		(-1.59)		(-2.95)**	(0.23)			(-2.11)**
	Korea	-1.68		-0.30	-0.24			-0.35
		(-4.32)**		(-3.72)**	(-1.25)			(-2.96)**
	Germany	-0.63		-0.41	-0.08			-0.43
		(-3.62)**		(-3.21)**	(-1.04)			(-4.77)**
Bulk	Japan	-0.23		-0.31	-0.13			-0.23
		(-1.84)*		(-3.43)**	(-0.44)			(-2.70)**
	Mexico	-0.07		-0.62	0.76			-0.46
		(-0.25)		(-5.07)**	(1.84)*			(-4.16)**
	Indonesia	-0.20		-0.74	-0.27	-0.10	0.41	-0.58
		(-0.65)		(-4.82)**	(-0.85)	(-0.49)	(1.88)*	(-5.63)**
	Canada	0.19		-0.31	-0.19			-0.22
		(1.23)		(-3.41)**	(-0.85)			(-2.77)**
	Korea	0.09	1.01	-0.30	-0.50			-0.78
		(0.16)	(1.63)	(-3.22)**	(-1.32)			(-6.83)**
	Thailand	-0.25		-0.80	-0.66			-0.42
		(-1.08)		(-6.22)**	(-3.15)**			(-4.04)**
	Turkey	-2.06		-0.42	0.50			-0.47
		(-2.73)**		(-4.19)**	(0.91)			(-2.81)**
	Germany	-0.37		-0.45	-0.24			-0.26
		(-1.06)		(-1.76)*	(-1.47)			(-3.01)**
	Netherlands	2.36		-0.49	-0.74			-0.73
		(1.94)*		(-3.73)**	(-1.37)			(-5.77)**
	Spain	-0.52		-0.93	-0.55			-0.82
		(-1.33)		(-3.55)**	(-1.24)			(-6.49)**

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

Table 5. Results of stability test

		Exports		Imports	
	Country	CUSUM	CUSUMSQ	CUSUM	CUSUMSQ
Consumer-Oriented	Canada	Stable	Stable	Stable	Stable
	Mexico	Stable	Stable	Stable	Stable
	Japan	Stable	Stable	Stable	Stable
	Australia	Stable	Unstable	Stable	Unstable
	Netherlands	Stable	Stable	Stable	Stable
	Italy	Stable	Stable	Stable	Stable
	France	Stable	Stable	Stable	Stable
	New Zealand	Stable	Stable	Stable	Stable
	Korea	Stable	Stable	Stable	Stable
	Germany	Stable	Stable	Stable	Stable
Bulk	Japan	Stable	Unstable	Stable	Stable
	Mexico	Stable	Stable	Stable	Unstable
	Indonesia	Stable	Stable	Stable	Stable
	Canada	Stable	Stable	Stable	Stable
	Korea	Stable	Stable	Stable	Stable
	Thailand	Stable	Stable	Stable	Stable
	Turkey	Stable	Stable	Stable	Unstable
	Germany	Stable	Unstable	Stable	Stable
	Netherlands	Stable	Stable	Stable	Unstable
	Spain	Stable	Stable	Stable	Stable

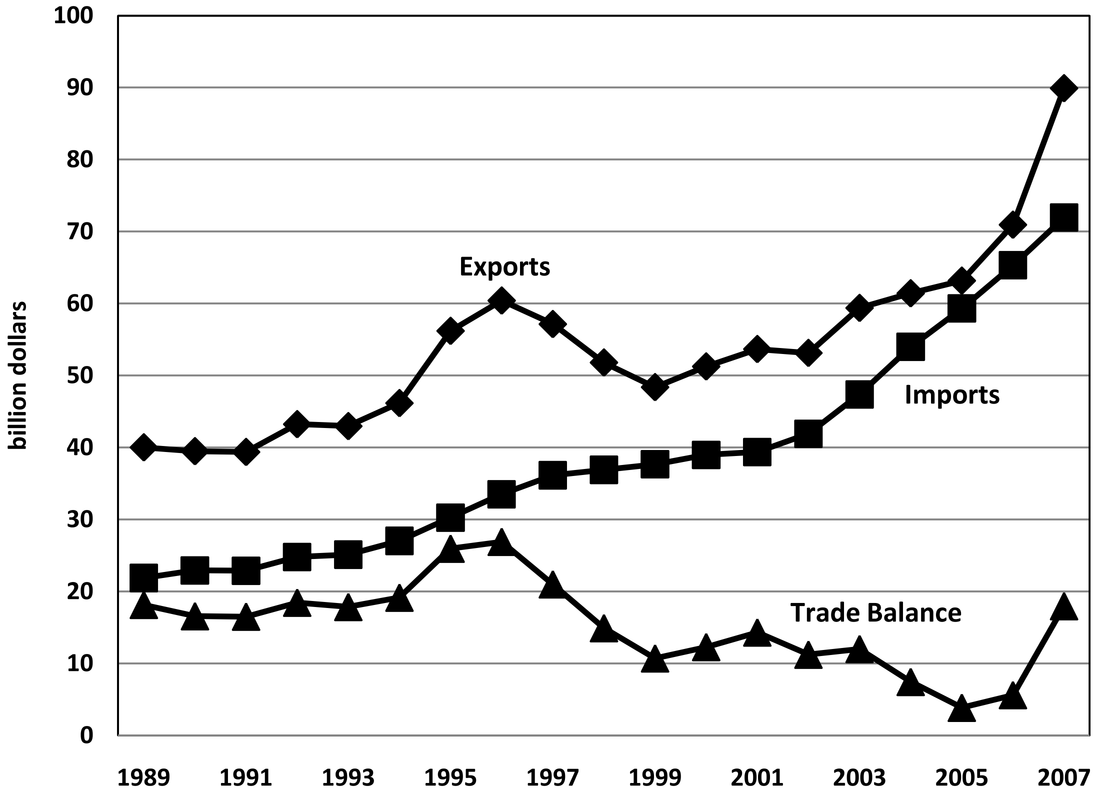


Figure 1. U.S. agricultural trade

Source: Economic Research Service (ERS), USDA

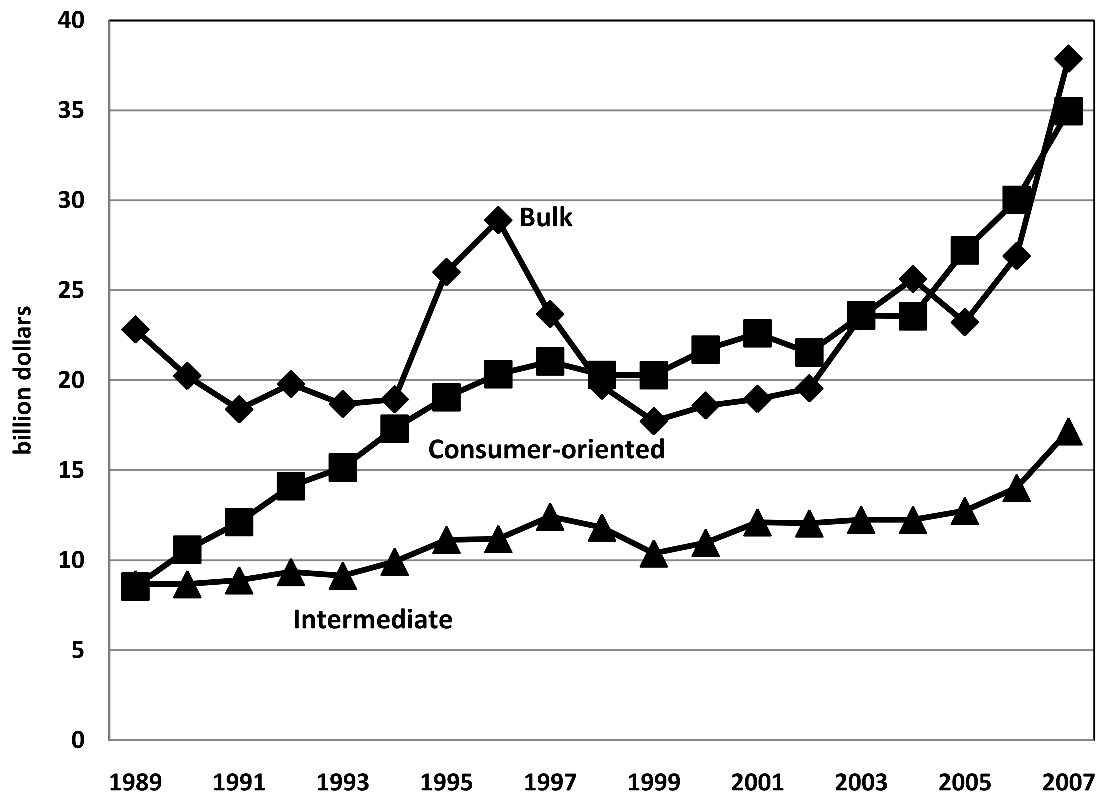


Figure 2. U.S. exports of bulk, intermediate and consumer-oriented products

Source: Economic Research Service (ERS), USDA

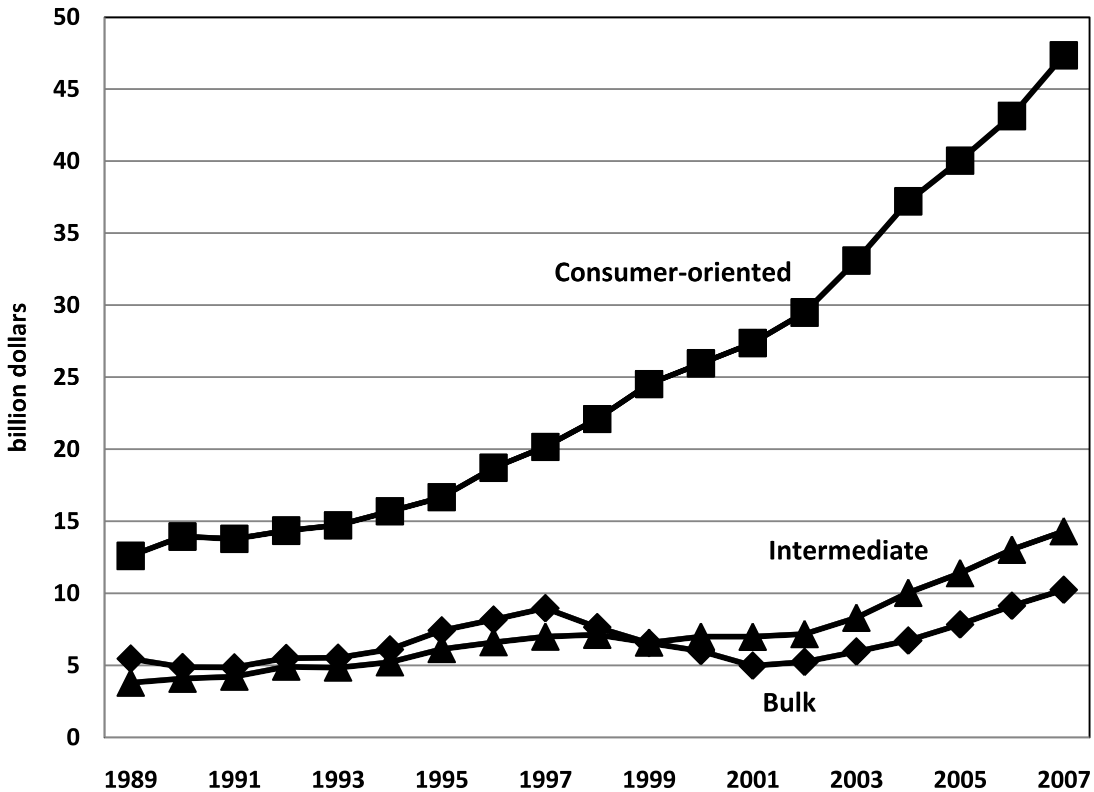
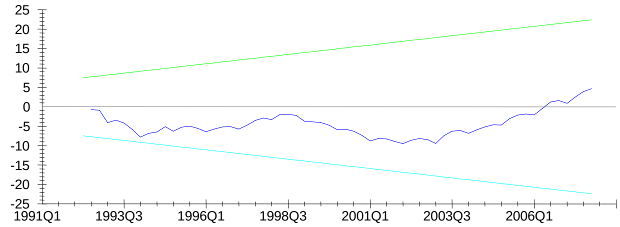


Figure 3. U.S. imports of bulk, intermediate and consumer-oriented products

Source: Economic Research Service (ERS), USDA

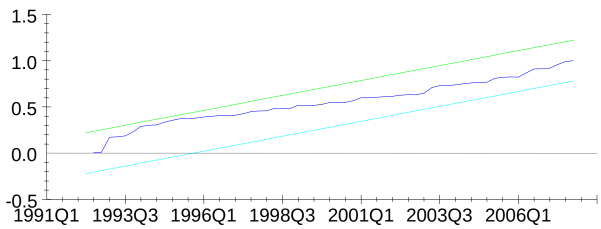
(a) U.S. exports of consumer-oriented products to Canada

Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

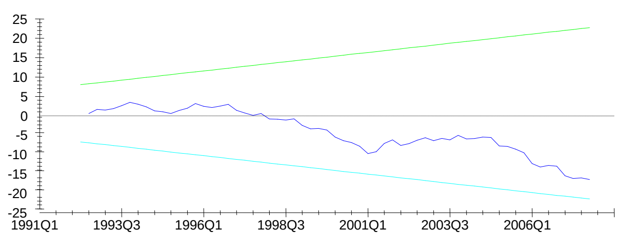
Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

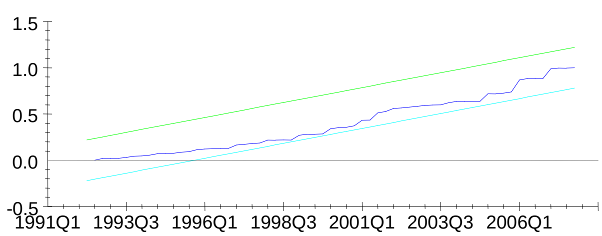
(b) U.S. imports of consumer-oriented products from Canada

Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

Plot of Cumulative Sum of Squares of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

Figure 4. An example of stability test results (U.S. exports and imports of consumer-oriented products to Canada)