EXCHANGE RATES, EXCHANGE RISK AND US CODFISH IMPORTS

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ABSTRACT

The effects of exchange rates and risk on U.S. codfish imports from Canada, China, Norway and Iceland are examined in the context of the Armington framework. The exchange rate volatility is measured by the generalized autoregressive conditional heteroskedasticity (GARCH) method. The nonstationarities of time-series data are explicitly taken into account by employing the Johansen test, the fully modified ordinary least squares (FMOLS) method and the autoregressive distributed lag (ARDL) model. Significant long-run effects of exchange rate volatilities and competing suppliers' currency uncertainties are supported by the data and cointegration tests. However short-run effects of volatilities cannot be verified by the relevant error correction models (ECM). The estimation results also show that the long-run impact of exchange rate on the import demand is generally larger than that of the relative price. The empirical results in general indicate the important impacts of exchange rates and risk on the U.S. codfish imports.

Keywords: codfish, import demand, exchange rates, volatility

INTRODUCTION

The important influence of exchange rate on trade has been recognized by economists and decision-makers for long time. The U.S. attributes the huge trade deficit with China to the non-market determined exchange rate of Chinese currency. Trade policy targeting the exchange rate is based on the assumption that the exchange rate directly affects trade flows. If exchange rate pass-through is complete, a one-percent appreciation of Chinese currency would lead to a reduction of imports from China by the same degree as a one-percent rise in the destination price. One important factor causing an incomplete exchange rate pass-through is exchange rate uncertainty (Bacchetta and Wincoop, 2005) [6].[a]

Extensive literature explores the impacts of exchange rates and volatility on international trade. The consensus is that changes in exchange rate can affect terms of trade and comparative advantages between countries as long as they affect the relative prices between traded goods and nontraded goods. Evidence is found in the literature in support of the long-run co-movement of the real exchange rate and commodity prices. Campa and Goldberg (2005) [9] provide evidence of partial exchange rate pass-through in the short-run, but over the long-run, the producer currency pricing is more prevalent for many imported goods. Gervais and Khraief (2007) [19] estimate the exchange rate pass-through elasticity for Canadian exports to the United States in the range of −0.2 to −0.7. These findings indicate that exchange rate may have a monetary effect independent of the price effect.

A substantial amount of research focuses specially on the effect of exchange rate risk on international trade flows within a bilateral world framework. Ariccia (1999) [15] confirms the inverse relationship between exchange rate uncertainty and international trade with data on Eastern European countries using different proxies for uncertainty. The same result is uncovered by Doroodian (1999) [16] using data on three developing countries, by Chowhury (1993) [12] in the case of each of G-7 countries, and by Bahmani-Oskooee and Ltaifa (1992) [8] in both developed countries and developing countries. The study of Sauera and Bohara (2001) [32] provides evidence in support of the negative effect of exchange rate volatility on most Low Developed Country (LDC) exports, but not for exports from Asian LDCs or industrialized countries. Except for different responses to volatility between developed countries and developing countries, responses of different trade sectors are also different. In a study of Taiwan, Wang and Barrett (2007) [34] find that real exchange rate risk has an insignificant effect in most sectors although agricultural trade volumes appear highly responsive to real exchange rate volatility. The study of Cho, Sheldon and McCorriston (2002) [11] shows that, in ten developed countries, the negative impact of uncertainty on agricultural trade is more significant compared to other sectors. Kandilov (2008) [25] extends the study of Cho, Sheldon and McCorriston [11] by taking into account agricultural export subsidies and incorporating a broad sample. The author concludes that, controlling for subsidy results, the effect of exchange rate risk declines by 50% compared to the results found in the original paper and the effect is generally larger for developing countries.

One possible reason explaining the above mixed empirical results is related to the aggregation problem. Aggregate trade flows can obscure the specific effect of exchange rate volatility on a particular commodity. There is increasing evidence supporting that exchange rate uncertainty has different effects on different markets, and the impact of exchange risk should be evaluated in the context of disaggregate data (McKenzie, 1999) [27]. Few papers focus on the effect of exchange rate uncertainty on a particular commodity. The exceptions are several papers about agricultural commodities. The agricultural industry is more sensitive to exchange rate risk because agriculture is typically traded with flexible pricing and agricultural goods are less storable than manufactured products. Anderson and Garcia (1989) [5] explain the effect of exchange rate risk on bilateral soybean trade flows and they conclude that the effect varies across countries for the reasons of different forward markets, alternative suppliers and so on. Langley, Giugale, Meyers and Hallahan (2000) [26] reveal the negative effect of exchange rate uncertainty on Thailand's poultry exports rather than aggregate exports.

Another reason for the indeterminate results in the literature may be attributed to lag effects of exchange rate risk. Most of the literature does not distinguish the short-run and long-run effects of exchange rate risk. The short-run exchange rate uncertainty can easily be offset by importers (or exporters) through financial markets. In contrast, the long-run exchange rate risk cannot be easily offset due to liquidity constraints and transaction costs.

Further, most of the study on the effect of exchange rate volatility is in the framework of the bilateral world. In the real multi-country world, bilateral trade is always affected by the factors of third-country, i.e. competing importers or exporters. Alston, Carter, Gray and Sumner (1997) [3] add the third-country factors into the analysis of the market responses when the grain trade barriers between the U.S. and Canada are reduced. Andino, Mulik and Koo (2005) [4] provide evidence that soybean imports of Asian countries from the U.S. are affected by depreciation (appreciation) of competitor's currencies relative to the importers' currencies. With *a prior* belief that multinational exchange rates are well-integrated, third-country effects in terms of exchange rates and volatility deserve more attention in the analysis of trade flows. Movements in one exchange rate can be offset by other factors, such as movements in other exchange rates relative to the currency of the same importer (or exporter). In the study of the U.S. bilateral export flows to its six largest trading partners, Cushman (1986) [14] finds individually and jointly significant third-country risk effects in most trading flows. Jin, Cho and Koo (2004) [24] point out that U.S. currency value and volatility, along with prices and volatilities of substitutes have significant effects on U.S. market shares in the Asian wheat market. However, the study of Dell'Ariccia [15] reflects a negative effect of third-country exchange rate volatility on bilateral trade probably due to multicollinearity.

Among the above literature, few studies are based on a complete demand system, which may also cause the inclusive empirical results. When focusing on a specific commodity, importers are assumed to allocate budgets among different suppliers. Therefore, the import demand system based on *a priori* utility function can reveal fully the decision process of importers.

Against this background, the objective of the present paper is to determine whether exchange rate and its risk affect a particular imported commodity. The import demand system is derived from a utility framework. The regression techniques in terms of characteristics of the time-series data generating processes are employed to distinguish the short-run and long-run impacts of exchange rate volatility. When modeling, third-county factors are controlled to improve the regression, and the impacts of price and exchange rate are distinguished in the demand equation. The U.S. codfish import market is taken as an empirical example. This market is a good case study because of variations of the market in the past twenty years and various characteristics of suppliers, to which we return below.

In what follows, we first present a brief overview of the U.S. codfish imports market. The theoretical framework is discussed in the next section, followed by empirical models of the U.S. codfish import demand. The latter includes stationary analysis of variables and discussion of estimation. The final section consists of summary and conclusions.

U.S. CODFISH IMPORT MARKET

Codfish, or cod, is a species of deep sea fish common to the North Pacific Ocean and North Atlantic Ocean. Except for the domestic production in the waters off Alaska, the U.S. also imports codfish from different countries. Canada, Norway, China, Iceland, Russia, Denmark and Fayene are the leading countries exporting codfish to the U.S. during the past twenty years. From 1990 to 2009, the U.S. codfish import volumes have experienced dramatic changes. The U.S. imported 283 thousand metric tons of codfish from these seven countries in 1990; however the volume sharply declined to 120 thousand metric tons in 2008. But the variation of imports has become weak since 1994. Figure 1 shows the trend of codfish imports in the U.S. market during the investigation period.

The sharp decline in the first several years of the sample period mainly affects the imports from Canada. One reason is a moratorium on the cod fishery declared by the Canadian government in 1992. From 1994, China displaced Canada and dominated the U.S. codfish import market. In 1990, the import volume of Canadian codfish accounted for 88.43% of the total codfish imports while Chinese codfish only 0.46%. In 2008, the positions for these two countries are inversed in the market (4.43% vs. 84.00%).

Since target of the present empirical study is to evaluate impacts of exchange rates and risk on import demands, sample selection of countries is based on the trade volume, regular adjusting exchange rate and variation of exchange rate. The Russian currency values fluctuate dramatically during the data period. The market shares of Fayene and Denmark are relatively small. Therefore, Russia, Fayene and Denmark are excluded from the sample. On average, the annual imports of codfish from Canada, China, Norway and Iceland account for about 86% of the total codfish imports.

Import from Canada is the Pacific codfish, which is also caught in the Gulf of Alaska state waters. Norway and Iceland export to the US the Atlantic codfish caught in the northern Atlantic Ocean. Codfish import from China is mainly re-exported. Volumes of codfish imports from the four countries during the sample period are also presented in figure 1. Codfish import from China has substantially increased during this period, while imports from Canada, Norway and Iceland show decreasing trends.

The U.S. codfish import market is a good empirical test case for our study. This market is relatively concentrated and the market shares of the main suppliers varied dramatically over the sample period, which makes the construction of an empirical demand system feasible. Agricultural products are more sensitive to exchange rate volatility. Further, the sample includes the low developed country (China) and the developed country (Canada) that may have different responses to volatility, as explored in the literature.

THEORETICAL FOUNDATION

After considering above-mentioned factors, the Armington framework (Armington 1969) [1], based on a two-step budgeting process assumption, is applied to derive the demand system. In the first stage, total expenditure is allocated over broad groups of goods. In the second stage, the expenditure of a particular group such as codfish import is divided among various suppliers. As Alston, Carter and Pick (1990) [2] states, ease of use and flexibility are two main reasons why the Armington specification is popular in the disaggregate model. Moreover, the relative price in the Armington model reflects the comparative advantage which is the main force directing trade flows. In this model, it is easy to isolate the effect of exchange rate from the demand system and to untangle its risk from other variables. Consistency between the exchange rate pass-through elasticity and the 'Armington' elasticity of substitution is explored by Warr (2005) [35], indicating that the Armington model can be applied to clarify the determinants of a specific commodity from different exporters. Solomon and Kinnucan (1993) [33] employ the extended Armington model to estimate the effect of government-subsidized export promotion on the import demand for the U.S. cotton in the Pacific Rim. In their study, the exchange rate is isolated and treated as a shift factor. Duffy, Wohlgenant and Richardson (1990) [17] use the Armington approach to develop estimates of the elasticity of import demand for the U.S. cotton.

Based on the Armington theory, the demand function for a given good in a particular group is

$$\frac{P_i M_i}{P^* M^*} = b_i^{\sigma} \cdot \left(\frac{P_i}{P^*}\right)^{(1-\sigma)} \qquad \text{(Eq. 1)}$$

where P_i and M_i are respective price and import from country i; P^* represents a price index for the group in question ($\ln P^* = \Sigma w_i \ln P_i$); M^* is a quantity index ($\ln M^* = \Sigma w_i \ln M_i$); the symbol $P_i M_i / P^* M^*$ indicates market share of country i in the market; and σ represents the elasticity of substitution between any pairs of products in this market.

Taking the logarithm of both sides of equation (1) yields

$$\ln\left(\frac{P_i M_i}{P^* M^*}\right) = \alpha_0 + \alpha_1 \ln\left(\frac{P_i}{P^*}\right) \qquad \text{(Eq. 2)}$$

where $\alpha_0 = \ln b_i^{\sigma}$, and $\alpha_1 = 1-\sigma$.

According to Alston, Carter and Pick [2], assumptions embedded in the Armington model may be overly restrictive; and it would be desirable and appropriate to use a less severely restrictive set of assumptions about demand relationships. In this paper, we consider homotheticity a testable hypothesis. Through simple algebra transformation of equation (2), we obtain the following import demand equation

$$\ln M_i = \alpha_0 + \ln M^* + (\alpha_1 - 1) \ln\left(\frac{P_i}{P^*}\right). \qquad \text{(Eq. 3)}$$

Finally, the corresponding empirical model is in the form

$$\ln M_i = \beta_0 + \beta_1 \ln M^* + \beta_2 \ln\left(\frac{P_i}{P^*}\right) \qquad \text{(Eq. 4)}$$

where β_0, β_1 and β_2 are parameters to estimate.

As discussed in the previous sector, exchange rate may have a monetary effect which is independent of price effect. For importers, the responses to a change of price may differ from the responses to a change of exchange rate, i.e. money illusion. In the Armington model, the effect of all cross prices can be ignored. But as argued by Chamber and Just (1979) [10], the exchange rate perhaps affects all cross prices, and thus the aggregate impact of exchange rate may be very significant, especially for a small group of commodities. Therefore, after considering price in equation (4) importer currency (US dollar), the demand equation becomes

$$\ln M_i = \beta_0 + \beta_1 \ln M^* + \beta_2 \ln \left(\frac{P_i^{US}}{P^*}\right) \qquad \text{(Eq. 5)}$$

where P_i^{US} is the US dollar price (destination price) of imported cod. The relationship between the exporter currency price and the destination price is expressed as

$$P_i^{US} = P_i^F \cdot E^{us/F} \qquad \text{(Eq. 6)}$$

where P_i^F is the foreign currency price and $E^{us/F}$ is the US dollar price of foreign currency. Taking the logarithm of both sides of equation (6) yields the equation

$$\ln P_i^{US} = \ln P_i^F + \ln E^{us/F}. \qquad \text{(Eq. 7)}$$

By substituting (7) into equation (5), we get the extended empirical model

$$\ln M_i = \beta_0 + \beta_1 \ln M^* + \beta_2 \ln \left(\frac{P_i^F}{P^*}\right) + \beta_3 \ln E_i \qquad \text{(Eq. 8)}$$

where the constraint $\beta_2 = \beta_3$ is used to test against no money illusion.

The import demand function specifies retrospectively the utility function which the importing firm maximizes. Within the determinants of total imports and unit prices, the form of the utility function is defined as a mean-variance utility function (Hopper 1978, Cushman 1986) [22, 14], which means that the firm's utility is a linear function of the mean profit and the variance of profit. Herein, the negative expenditure of import is treated as "profit". The bigger the "profit" (the smaller absolute value of expenditure), the greater the utility obtained by the firm. And a lower variance of "profit" improves the firm's utility.

Without loss of generality, the firm is supposed to import from two suppliers: A and B. The utility function and the profit function are as follows

$$U = E(\Pi) - \frac{\gamma}{2} V(\Pi) \qquad \text{(Eq. 9)}$$

$$\Pi = -P_1 M_1 - P_2 M_2 \qquad \text{(Eq. 10)}$$

where Π is total profit and $V(\Pi)$ is volatility of profit. M_1 and P_1 are the amount and unit price of codfish imported from country A; and M_2 and P_2 from country B. The constant γ measures the degree of risk aversion.

Substituting Π into (9) yields the following objective function

$$U = -E(P_1)M_1 - E(P_2)M_2 - \frac{\gamma}{2}\left[M_1^2 V(P_1) + M_2^2 V(P_2) + 2M_1 M_2 \mathrm{Cov}(P_1, P_2)\right] \qquad \text{(Eq. 11)}$$

Maximizing the firm's utility yields the first order conditions

$$-E(P_1) - \gamma M_1 V(P_1) - \gamma M_1 \mathrm{Cov}(P_1, P_2) = 0 \qquad \text{(Eq. 12)}$$

$$-E(P_2) - \gamma M_2 V(P_2) - \gamma M_2 \mathrm{Cov}(P_1, P_2) = 0 \qquad \text{(Eq. 13)}$$

Solving (12) and (13) simultaneously, we obtain the import demand functions of codfish as follows

$$M_1 = \frac{\mathrm{Cov}(P_1,P_2)E(P_2) - E(P_1)V(P_2)}{\gamma D} \qquad \text{(Eq. 14)}$$

$$M_2 = \frac{\mathrm{Cov}(P_1,P_2)E(P_1) - E(P_1)V(P_1)}{\gamma D} \qquad \text{(Eq. 15)}$$

From equations (14) and (15), the effects of own-price risk and competitor's price risk can be expressed as

$$\frac{\partial M_1}{\partial V(P_1)} = \frac{-M_1 V(P_2)}{\gamma D} \prec 0 \qquad \text{(Eq. 16)}$$

$$\frac{\partial M_1}{\partial V(P_2)} = \frac{M_2 \mathrm{Cov}(P_1, P_2)}{D} \lesseqgtr 0 \qquad \text{(Eq. 17)}$$

where $D = V(P_1) * V(P_2) - \mathrm{Cov}(P_1, P_2)^2 \succ 0$, unless the coefficient of correlation between P_1 and P_2 equals ± 1 (Cushman) [14]. Therefore the expected sign of own-price risk, which is restricted to the exchange rate uncertainty, is negative. But the sign of equation (17) is uncertain depending on the sign of $\mathrm{Cov}(P_1, P_2)$.

EMPIRICAL MODEL AND DATA

The empirical model of U.S. codfish import demand is specified by relaxing the homotheticity assumption and incorporating the exchange rate and its volatility in the specification. For an import demand equation, except for the bilateral exchange rate volatility, other three suppliers' currency volatilities are also incorporated to capture the third-country impacts. The resulting specification of the U.S. codfish demand function is as follows

$$\ln M_{i,t} = \beta_0 + \beta_1 \ln M_t^* + \beta_2 \ln \left(\frac{P_{i,}^F}{P^*}\right)_t + \beta_3 \ln E_{i,t} + \beta_4 \ln V(E_i)$$
$$+ \beta 5 \ln V(Ej)t + \beta 6 \ln V(Ek)t + \beta 7 \ln V(El)t + ei,t \qquad \text{(Eq. 18)}$$

where i denotes supplier, i.e. Canada, Norway, China or Iceland (simply, can, nor, chn and ice, respectively); t stands for the time quarter; and

$M_{i,t}$ ≡ volume of codfish import from supplier i
M_t^* ≡ index of total imports
$(\frac{P_i^F}{P^*})_t$ ≡ relative price of cod imported from supplier i
$P^F_{i,t}$ is measured in the supplier's currency
P_t^* represents the price index
$E_{i,t}$ ≡ nominal exchange rate (USD / supplier's currency)
$V(E_i)_t$ ≡ volatility of bilateral real exchange rates
$V(E_j)_t$, $V(E_k)_t$ and $V(E_l)_t$ ≡ competing suppliers' currency volatilities
$e_{i,t}$ ≡ error term

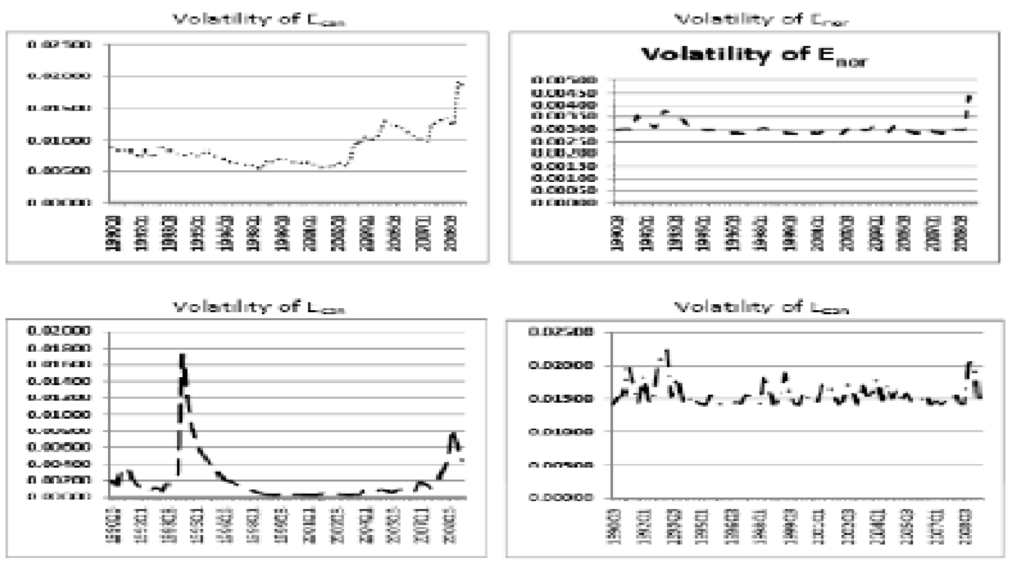

Figure 1: Exchange rate volatilities (GARCH method)

The coefficient of $M_{i,t}^*$ is expected to have a positive sign since the volume of codfish import from a particular supplier depends on the total amount of codfish imports from all sources. An increase in the relative price or appreciation of a specific supplier's currency is assumed to depress import demand. The coefficient of bilateral exchange rate volatility is expected to be negative; however the coefficient of competing supplier's currency volatility remains obscure.

The data on the U.S. codfish imports are from the Norwegian Seafood Export Council (NSEC). Prices are freight on board (FOB) measured at the wholesale level. The number of observations is 78, from 1990 Q1 to 2009 Q2. The data of real exchange rates are collected from the USDA-ERS. The nominal exchange rates are from the International Monetary Fund (IMF) series *International Financial Statistics*.

Different measures of exchange rate volatility are used in the literature. Generally, these measures can be divided into two types. One is the unconditional volatility represented by the moving standard deviation; another is the conditional volatility represented by the generalized autoregressive conditional heteroskedasticity method (GARCH). The unconditional volatility reflects the realized (*ex post*) uncertainty and ignores the relevant information on the random generalized exchange rates (Kandilov 2008) [25]. On the other hand, the GARCH method is based on the past information and can fully reveal the characteristics of financial data. In addition, the GARCH model can be tested whether the movement in the conditional variance of a variable overtime is statistically significant or not.

Therefore, the conditional volatility rather than traditional volatility is employed in this paper. In the GARCH method, the stochastic error is obtained from an autoregressive moving-average ARMA(1,1) process, and the lags in the GARCH are specified to be one. The process follows that

$$\text{ARMA(1,1):} \quad E_{i,t} = \pi_0 + \pi_1 E_{i,t-1} + \pi_2 v_{i,t-1} + v_{i,t} \qquad \text{(Eq. 19)}$$

$$\text{GARCH(1,1):} \quad V(Ei)t = \delta 0 + \delta 1\, v_{i,t-1}^2 + \delta 2\, V(Ei)^2_{t-1} + \mu_{i,t} \qquad \text{(Eq. 20)}$$

where, $E_{i,t}$ is bilateral exchange rate; $V(E_i)_t$ represents the conditional volatility of supplier's currency values; and $\nu_{i,t}$ and $\mu_{i,t}$ are error terms of ARMA(1,1) and GARCH (1,1) processes, respectively.

Figure 1 illustrates the trends of volatilities for the U.S. dollar exchange rate of each supplier. The exchange rate volatility of Canada has a high degree of variation during the investigating period. In contrast, except for several outliers, the exchange rate volatility of China has a considerably low degree of fluctuation, indicating the consequences of government policies and macroeconomic factors.

REGRESSION RESUTLS

In time series analysis, the presence of a unit-root process may lead to serious errors in inference. Thus, in the multivariate time series context, the first step is always to test stationarity of each variable (Enders 1995, p. 266) [18]. If the variables are integrated in the same order of one, the next step is to test a linear combination of the variables resulted in a stationary process. Given the existence of a cointergrating relationship among variables, the last step is to evaluate the error adjustment mechanism.

As a first step, the stationarity properties of variables are evaluated by implementing a typical Dickey Fuller (DF) test or Augmented Dickey Fuller (ADF) test. The ADF test with drift and trend and ADF test with drift only reflect different characteristics of a data series. The order of augment in the ADF test is determined by the Akaike Information Criterion (AIC). If the null hypothesis cannot be rejected by one of the tests, we can conclude that the data series is integrated of order one, I(1).

At the 5 percent level, the null hypothesis of a stationary process cannot be rejected for all variables, indicating the variables included in the models are integrated of order one. After confirming that all variables are integrated processes of the same order one, we test for the presence of a common stochastic trend for the variables in each demand equation. If there is no linear combination of these variables resulted in a stationary process, the regression of equation (18) is spurious. The Johansen procedure is implemented to test for cointegration in the import demand expression. Since the method is based on the vector autoregressive regression (VAR) approach, all variables incorporated in the model are assumed to be endogenous. Hence, temporarily, the competing exporters' currency volatilities are excluded from the import demand equation. For the trace tests, only the null hypothesis r□0 can be uniformly rejected at the 10% level, reflecting that alternative hypothesis $r \geq 1$ is preferred. The eigenvalue tests produce the same results as the trace tests. Only the null hypothesis r□0 is uniformly rejected in favor of the alternative hypothesis r□1. Consequently, these test results indicate that, for each demand equation, the variable import (Mi) is cointegrated with the total import, relative price, exchange rate and volatility. Therefore equation (12) reflects a long-run or equilibrium relationship between the dependent variable and explanatory variables. Hence the estimated parameters can be interpreted as long-run elasticities.

Ignoring the third-country impacts from the demand equation may lead to misspecification. By control for the competing suppliers' currency volatility on the demand system, the long run model represented by equation (18) can be regressed by the Ordinary Least Squares (OLS) approach. But, in the demand system, price is always expected to be endogenous. Thus estimates of the long-run relationship using the OLS approach may be biased. Moreover, for the error term from the regression, the null hypothesis of no autocorrelation is rejected by the DW test in all cases.b One alternative to estimate the long-run relationship is the fully modified OLS (FMOLS) method proposed by Phillips and Hansen (1990) [30]. In the FMOLS approach, firstly, the innovations from the cointegrating equation and stochastic regressors are used to derive the long-run covariance matrices. From these matrices, the data are modified and an estimated bias correction term is created. In this step, endogeneity can be controlled through the bias correction term. By including the bias correction term, a semi-parametric procedure is applied to produce the FMOLS estimator for the cointegrating equation and the regressors equations jointly. In the last step, the serial correlation between the two innovations is controlled. Table 1 reports the FMOLS estimates of the cointegrating equations

Table 1: Estimates of Long Run Model Using the FMOLS Approach

	Canada	Norway	China	Iceland
lnM*	1.270***	-0.700***	-0.687***	0.171*
	(3.494)	(-3.447)	(-2.782)	(1.648)
$\ln(P_i/P^*)$	-4.976***	-2.775***	2.897***	-3.806**
	(-5.144)	(-5.102)	(5.358)	(-10.002)
$\ln E_i$	-8.485***	-21.782***	9.871	-4.809***
	(-3.219)	(-3.529)	(1.482)	(-8.118)
$\ln V(E_{can})$	-0.972*	-1.607	2.514***	0.016
	(-1.682)	(-4.524)	(5.428)	(0.060)
$\ln V(E_{nor})$	3.480***	-1.899*	-2.824*	-0.665
	(2.773)	(-1.798)	(-1.851)	(-1.074)
$\ln V(E_{chn})$	-0.182**	0.164*	-0.605***	-0.157***
	(-2.009)	(1.713)	(-6.786)	(-3.543)
$\ln V(E_{ice})$	0.334	1.061*	-0.00007	0.010
	(0.420)	(1.722)	(-0.00008)	(0.032)
R-square	0.785	0.824	0.723	0.838
Number of Obs	75	75	75	75

Note: t-ratios in parentheses; ***P≤0.01; **P≤0.05; **P≤0.1.

In all cases but one (Iceland), the bilateral exchange rate volatility has a significant and negative impact on import, indicating that a high volatility depresses codfish import. The Norwegian codfish import with respect to bilateral exchange rate volatility is more elastic (-1.899), whereas the responses of Canadian codfish and Chinese codfish imports are less elastic (-0.972 and -0.605). For the three import demand functions, effects of almost all competing suppliers' currency volatilities are captured with either a positive or negative sign. This confirms the hypothesis that the effect of third-county exchange risk is ambiguous. Icelandic currency volatility only has an impact on the Norwegian codfish import. Moreover, the bilateral exchange rate volatility of Iceland has no effect on codfish import from this country.

The estimated coefficients of relative price and exchange rate are statistically significant with expected negative signs in all import demand equations except for the Chinese codfish import equation. Further, in each equation, the magnitude of relative price coefficient is smaller than that of exchange rate coefficient, reflecting the strong and different impact of exchange rate on imports. For example, in the Canadian codfish demand equation, a one-percent increase of relative price leads to a reduction of Canadian codfish imports by 4.976%, whereas the same change of exchange rate causes 8.485% reductions. The estimation results overall indicate that the exchange rate and risk play more different roles in the U.S. import demand for codfish compared to codfish price itself.

The error term in the cointegration model reflects the deviation from equilibrium in the short-run. By incorporating the residue from the cointegration model, the error-correction model (ECM) is applied to investigate the error correction adjustment and the short-run volatility elasticity. This specification gives rise to the estimation equation

$$\Delta \ln M_{i,t} = \gamma_0 + \gamma_1 \Delta \ln M_t^* + \gamma_2 \Delta \ln \left(\frac{P_i^F}{P^*}\right)_t + \gamma_3 \Delta \ln E_{i,t} + \gamma_4 \Delta \ln V(E_i)_t + \gamma 5\, \Delta \ln V(E_j)_t + \gamma 6\, \Delta \ln V(E_k)_t + \gamma 7\, \Delta \ln V(E_l)_t + \gamma 8\, e_{i,t-1}^c + u_{i,t} \quad \text{(Eq. 21)}$$

where $e_{i,t-1}^c$ is the one period lagged value of the error term from the cointegrated regression; Δ as usual denotes the first difference operator.

Table 2: Estimates of ECM (lagged residual from FMOLS)

	Canada	Norway	China	Iceland
Residual $_{t-1}$	-0.259***	-0.541***	-0.262***	-0.724***
	(-3.005)	(-5.258)	(-2.794)	(-5.491)
Δ lnM*	0.658***	-0.011	0.57**	0.265
	(3.452)	(-0.050)	(2.146)	(1.676)
Δ ln(P_i/P*)	-2.625***	-2.022***	-1.162***	-2.177***
	(-4.373)	(-3.516)	(-2.846)	(-2.753)
Δ lnE_i	-5.824***	-25.973***	12.271	-2.614**
	(-3.196)	(-3.750)	(1.289)	(-2.425)
Δ lnV(E_{can})	0.330	-0.636	0.893	0.020
	(0.735)	(-1.184)	(1.520)	(0.051)
Δ lnV(E_{nor})	-0.123	0.169	-0.137	-0.009
	(-1.331)	(1.473)	(-1.104)	(-0.110)
Δ lnV(E_{chn})	1.222	-0.08	-1.296	-0.173
	(1.525)	(-0.084)	(-1.306)	(-0.249)
Δ lnV(E_{ice})	-0.205	0.418	0.675*	-0.073
	(-0.726)	(1.18)	(1.841)	(-0.298)
R-square	0.442	0.478	0.508	0.502
DW Test	2.248	1.956	2.168	2.083
Number of Obs	74	74	74	74

Note: t-ratios in parentheses; ***P≤0.01; **P≤0.05; **P≤0.1.

The estimated results of the ECM are reported in table 2. Except for the significant coefficient of Icelandic exchange rate volatility in the Chinese codfish import demand equation, none of volatility variables has a transitory effect on bilateral codfish trade flows, due to the insignificant coefficient. Changes of exchange volatilities affect the import demand after a particular time point. But, the parameter of error correction term is statistically significant with an expected negative sign in each model. This finding indicates that the validity of an equilibrium relationship among the variables in each cointegrating demand equation. The magnitudes of error correction terms vary from a high of 0.724 (absolute value) for import of Icelandic codfish to a low of 0.259 (absolute value) for import of Canadian codfish. This hints different adjustment mechanisms among the import demand equations. Accordingly, the carryover periods are in the range from 1.381 to 3.861 quarters with the shortest period in the Icelandic codfish case and the longest period in the Canadian codfish case.

Considering the strong caveat that the DF or ADF stationarity test has a low power, the variables may have different stationarity rather than I(1). In response, the autoregressive distributed lag (ARDL) cointegration procedure (Pesaran and Shin, 1999 Pesaran, Shin and Smith 2001) [28, 29] is applied to estimate the cointegrating relationship, for it is irrespective of whether the underlying regressors are I(1) or I(0). This method can be divided into three steps. In the first step, a conditional error correction ARDL equation is constructed and the joint insignificance of the lagged-level variables (null hypothesis) is tested against the existence of a conditional level relationship. If the null hypothesis is rejected, the next step is to estimate the long-run level relationship by including the lagged-level variables in the regression. Later the ECM is regressed where the order of lag for each variable is determined by the lag length in the long-run model. Consistency among the cointegration test, the long-run estimates and the error correction adjustment mechanism is guaranteed by the choice of order of lags for each variable.

Since the FMOLS regression of Icelandic codfish import is far from appealing, this equation is dropped from the demand system. Consequently, Icelandic currency volatility is excluded in other three import demand equations. To begin with, the long-run relationship between codfish import and a set of regressors is examined on the basis of the error correction representation of the ARDL model as follows

$$\begin{aligned}\Delta \ln M_{i,t} = {} & \alpha_0 + \sum_{q=1}^{n} b_q \Delta \ln M^*_{t-q} + \sum_{q=0}^{n} c_q \Delta \ln \left(\frac{P_i^F}{P^*}\right)_{t-q} + \sum_{q=0}^{n} d_q \Delta \ln E_{i,t-q} + \sum_{q=0}^{n} e_q \Delta \ln V(E_i)_{t-q} \\ & + \sum_{q=0}^{n} f_q \Delta \ln V(E_j)_{t-q} + \sum_{q=0}^{n} g_q \Delta \ln V(E_k)_{t-q} + \sum_{q=0}^{n} f_q \Delta \ln V(E_l)_{t-q} + \beta_1 \ln M^*_{t-1} \\ & + \beta_2 \ln \left(\frac{P_{i,}^F}{P^*}\right)_{t-1} + \beta_3 \ln E_{i,t-1} + \beta_4 \ln V(E_i)_{-1}) + \beta_5 \ln V(E_j)_{t-1} + \beta_6 \ln V(E_k)_{t-1} + \beta_7 \ln V(E_l)_{t-1} \\ & + u_{i,t} \qquad \text{(Eq. 22)}\end{aligned}$$

The ARDL cointegration test is performed by testing the joint significance of level variables in equation (22). The null hypothesis (H_0: $\beta_1 = \beta_2 = \beta_3 = \beta_4 = \beta_5 = \beta_6 = \beta_7 = 0$) means nonexistence of a long-run relationship among variables in the import demand equation. Pesaran and Smith [28] point out that the distribution of the calculated F-statistic is nonstandard. They produce the critical values including an upper bound and a lower bound in terms of all possible classifications of variables stationarities. Test results are sensitive to the order of lags in equation (22) (Bahmani-Oskooee and Bohl, 2000) [7]. When the order one of lag is imposed, F-statistic values are all less than the low bound critical value, indicating the null hypothesis cannot be rejected in each case. But, presuming the order two of lags, the computed F-statistic values are all bigger than the upper bound critical value at the 5% level, reflecting that there is a long-run equilibrium relationship between the import codfish and a set of regressors.[c, d]

All lagged level variables should not be retained in the regression for the long-run parameters. The long run model in terms of a more parsimonious specification is advised by Persaran, Shin and Smith [29]. In this study, seven variables are included in the long run model, thus there are in total 2^7 regressions since the maximum of lags for each variable is two. In the broad literature, the AIC or Schwarz's Bayesian Criterion (SBC) method is commonly used to determine an optimal lag length. But Pesaran and Shin (p386) [28] state, "….., in the context of the ARDL model inference on the long run parameters, δ and θ is quite simple and requires *a priori* knowledge or estimation of the orders of the extend ARDL(p, m) model." After considering the sample size in this study, we depend on *a prior* knowledge to set lag length in order to save the degree of freedom and avoid any over-parameterized specification. A crude benchmark of the adjustment period is from 1.381 to 3.861 quarters based on the FMOLS-ECM approach. The *a priori* knowledge is that the trade contracts in many sectors (especially agricultural exports) include agreement for delivery in less than 90 days (Wang and Barret) [34]. Moreover, the bilateral exchange rate volatility is central to this paper. Consequently, all variables in each import demand equation take order one of lag with the exception that the bilateral exchange volatility imposes order two of lags.

Table 3: The Estimated Long-Run Volatility Elasticicties

Model: ARDL(1,1,1,2,1,1)

Regressor	Canada	Norway	Iceland
$\ln VE_{can}$	-3.234*	5.000*	-0.706**
	(-1.735)	(1.814)	(-2.120)
$\ln VE_{nor}$	-2.176**	-5.304**	
	(-1.911)	(-2.237)	
$\ln VE_{chn}$			-1.006**
			(-2.011)

Note: t-ratios in parentheses; ***P≤0.01; **P≤0.05; **P≤0.1.
The propagation method is used to calculate the standard error.
The insignificant elasticities are not produced.

Estimation results of the long-run relationship for each model using the ARDL approach are reported in table 3. Like the findings from the FMOLS approach, the bilateral exchange rate volatility has a significant and negative effect on the codfish import demand in each demand equation, whereas the magnitudes are much greater than the counterparts from the FMOLS estimates, -3.234 vs. -0.972 (Canada), -5.304 vs. -1.900 (Norway) and -1.006 vs. -0.605 (China).[e] But, both of the two sets of parameters are generally larger than estimated effects of volatility on the bilateral aggregate agricultural commodity trade between the U.S. and other countries (Cho, Sheldon and McCorriston; and Kandilov) [11, 25]. On the other hand, they are generally closer to the estimates based on a specific agricultural commodity in the recent research (Jin, Cho and Koo; Anderson and Garcia) [24, 6]. This verifies the hypothesis that aggregate data can mitigate effects of exchange volatility on trade flows.

Table 4: Estimates of ECM (lagged residual from ARDL(1,1,1,2,1,1))

Regressor	Canada	Norway	China
$Residual_{t-1}$	-0.498***	-0.624***	-0.383***
	(-3.42)	(-4.41)	(-2.67)
Δ $lnM*_{t}$	0.600***	-0.181	1.077***
	(3.04)	(-0.86)	(4.33)
Δ $lnM*_{t-1}$	-0.334*	-0.182	0.394*
	(-1.75)	(-0.87)	(1.67)
Δ $ln(P_i/P*)_t$	-1.961***	-2.020***	-1.706***
	(2.90)	(-3.18)	(-3.77)
Δ $lnP_i/P*)_{t-1}$	-0.016	1.149**	-0.204
	(-0.03)	(2.12)	(-0.43)
Δ lnE_t	-4.789***	-19.141***	4.250
	(-2.44)	(-3.09)	(0.42)
Δ lnE_{t-1}	1.148	7.079	2.588
	(0.48)	(1.10)	(0.26)
Δ $lnV(E_{can})_t$	0.410	-0.933*	0.675
	(0.84)	(-1.91)	(1.10)
Δ $lnV(E_{can})_{t-1}$	-0.684	0.136	-0.417
	(-1.41)	(0.27)	(-0.69)
Δ $lnV(E_{can})_{t-2}$	0.498		
	(1.07)		
Δ $lnV(E_{nor})_t$	0.771	1.534	-0.561
	(0.85)	(1.64)	(-0.51)
Δ $lnV(E_{nor})_{t-1}$	0.043	-1.595*	-0.480
	(0.05)	(-1.78)	(-0.43)
Δ $lnV(E_{nor})_{t-2}$		-0.266	
		(-0.19)	
Δ $lnV(E_{chn})_t$	-0.150	0.158	-0.045
	(-1.48)	(1.51)	(-0.32)
Δ $lnV(E_{chn})_{t-1}$	-0.010	-0.105	0.131
	(-0.11)	(-1.00)	(0.98)
Δ $lnV(E_{chn})_{t-2}$			-0.013
			(-0.10)
R-square	0.549	0.557	0.572
DW Test	2.068	1.967	1.898
Number of Obs	71	71	71

Note: t-ratios in parentheses; ***P≤0.01; **P≤0.05; **P≤0.1.

The ARDL approach does not improve the estimates of third-country volatility in the long run models. For the significant variables, they have the same signs as the counterparts from the FMOLS approach in Canadian codfish and Norwegian codfish cases. In contrast, the competing suppliers' currency risks have no effects on codfish import from Chinese. This seems reasonable in that import of Chinese codfish keep rising during the sample period.

We further estimate the error correction model by incorporating the lagged error correction term from the ARDL long-run regression in the specification. The result (table 4) indicates that the efforts of bilateral exchange rate volatilities and competing suppliers' currency uncertainties cannot be captured in the Canadian and Chinese codfish import demand equations. But the significant effects of the bilateral exchange rate risk and one competing supplier's currency uncertainty are found in the Norwegian codfish demand equation, reflecting a slight improvement of the regression results compared to the FMOLS-ECM approach. In order to save the degree of freedom, we limit the number of lags imposed in the ARDL model, which may be insufficient to capture the dynamic effect of exchange rate volatility. But, in each case, the error correction term is statistically significant showing that the system converges to the long-term equilibrium. The carryover periods, ranging from 1.600 to 2.604, are shorter than the counterparts from the FMOLS-ECM approach, though the long-run estimates of volatilities are larger in the ARDL approach.

CONCLUDING COMMENTS

The effect of exchange rate volatility on trade flows has been well documented in the literature. However, empirical work in the context of a complete demand system has been fragmented. Under the framework of Armington formulation, the paper is concerned with impacts of exchange rate and risk on the import demand for codfish in the U.S. market. It can be argued that exchange rate volatility has a significant effect on the allocation of resources by the market participants. Based on the structure of the time-series data, the cointegrating relationship between codfish import and bilateral exchange rate volatility, total import, relative price and exchange rate is confirmed in each import demand equation of the demand system. But the relevant error correction models fail to capture the short-run effect of exchange rate volatility. These estimation results confirm the hypothesis that exchange rate volatility would affect trade through its effect on capacity and the lags could be fairly long (Cote, 1994) [13]. The effect of volatility in the long-run is accordance with the individual effect of exchange rate on import volumes. In almost all cases, by control for exchange rate volatility, the import demand is more sensitive to exchange rate compared to relative price in the short-run and long-run.

Four additional implications from the empirical findings deserve special attention. First, the exchange rate volatility has a different role in different market. In order to avoid aggregate bias, it is better to focus on individual commodities when estimating the effect of exchange rate volatility. Second, ignoring the stationary behaviour of the variables in the speciation causes the inconsistent results about exchange rate volatility. Another advantage of the time-series frame is that the long-run and short-run responses of the market can be separated on the basis of the data cointegrated processes. Third, exchange rate and its risk are important powers affecting the market structure. The relative stable exchange rate of exporting country may be one of the important factors driving importers to allocate resources to this supplier for a given commodity. Fourth, evaluating the economic policy about exchange rate should be based on the disaggregate data on commodities or sectors. Sources of exchange rate volatility are associated with macroeconomic policies. Including exchange rate volatility in the specification and focusing on disaggregate data may change the choice of economy policy options. For example, many trade policies are associated with the impact of exchange rate on trade flows. Although exchange rate movements can theoretically pass completely through into the import price in the long-run, effect of exchange rate volatility possibly makes persistent deviations from the long-run equilibrium occur.

REFERENCES

[1]. Armington, P.S. 1969 "A Theory of Demand for Products Distinguished by Place of Production." International Monetary Fund Staff Paper. 26: 159-78.

[2]. Alston, J. M., C.A. Carter and D. Pick. 1990 "Whither Armington Models?" *American Journal of Agricultural Economics*. 72(2): 455-67.

[3]. Alston, J. M., C.A. Carter, D. Pick and D.A. Summer. 1997. "Third-Country Effects and Second-Best Grain Trade Policies: Export Subsidies and Bilateral Liberalization." *American Journal of Agricultural Economics*. 79: 1300-10.

[4]. Andino, J., K. Mulik and W.W. Koo. 2005. "The Impact of Brazil and Argentina's Currency Devaluation on U. S. Soybean Trade." Agribusiness & Applied Economics Report No. 574, Center for Agricultural Policy and Trade Studies, North Dakota State University.
[5]. Anderson, M. and P. Garcia. 1989. "Exchange Rate Uncertainty and the Demand for U.S. Soybeans." *American Journal of Agricultural Economics*. 71(3): 721-29.
[6]. Bacchetta, P. and E.V. Wincoop, 2005. "A Theory of the Currency Denomination of International Trade." Journal of International Economics. 67: 295-319.
[7]. Bahmani-Oskooee, M. and M.T. Bohl. 2000. "German Monetary Unification and the Stability of the German M3 Money Demand Function," *Economics Letters*. 66: 203-208.
[8]. Bahmani-Oskooee, N. and N. Ltaifa. 1992. "Effects of Exchange Rate Risk on Exports: Cross country Analysis." *World Development.* 20: 1173-81.
[9]. Campa, J.M. and L.S. Goldberg. 2005. "Exchange Rate Pass-Through into Import Prices." *The Review of Economics and Statistics*. 87(4): 679-690.
[10]. Chambers, R.G. and R.E. Just. 1979. "A Critique of Exchange Rate Treatment in Agricultural Trade Models." *American Journal of Agricultural Economics*. 61: 249-57.
[11]. Cho, G., M. Sheldon and S. Mccorriston. 2002. "Exchange Rate Uncertainty and Agricultural Trade." *American Journal of Agricultural Economics*. 84(4): 932-42.
[12]. Chowdhury, A.R. 1993 "Does Exchange Rate Volatility Depress Trade Flows: Evidence From Error-Correction Models." *Review of Economics and Statistics*. 75(4): 45-63
[13]. Cote, A. 1994 "Exchange Rate Volatility and Trade: Survey". Bank of Canada Working Paper. No. 94-5
[14]. Cushman, D.O. 1986 "Has Exchange Risk Depressed International Trade? The Impact of Third-Country Exchange Risk*." Journal of International Money and Finance*. 5: 361-79.
[15]. Dell'Ariccia, G. 1999. "Exchange rate fluctuations and trade flows: evidence from the European Union." International Monetary Fund Staff Paper. 46: 315-34.
[16]. Doroodian, K. 1999 "Does Exchange Rate Volatility Deter International Trade in Developing Countries?" *Journal of Asian Economics* 10: 465-74.
[17]. Duffy, P., M.K. Wohlgenant, and J.W. Richardson. 1990. "The Elasticity of Export Demand for U. S. Cotton". *American Journal of Agricultural Economics*. 72(2): 485-96.
[18]. Enders, W. 1995. Applied econometric time series. Wiley, New York. 443 p.
[19]. Gervais, J.P. and N. Khraief. 2007. "Is Exchange Rate Pass-Through in Pork Meat Export Prices Constrained by the Supply of Live Hogs?" *American Journal of Agricultural Economics*. 89: 1058–72.
[20]. Grigsby, S.E., and C.A. Arnade. 1986. "The Effect of Exchange Rate Distortions on Grain Export Markets, the Case of Argentina." *American Journal of Agricultural Economics*. 68: 434-40.
[21]. Gujarati, D.N. 2003. "Basic Econometrics." McGraw-Hill Higher Education Press.
[22]. Hopper, P. and S.W. Kohlhagen. 1978 "The Effects of Floating Exchange Rate Uncertainty on the Prices and Volume of International Trade." *Journal of International Economics* 8: 483-511.
[23]. Houck, J.P. 1986 "Elements of Agricultural Trade Policies." Prospect Heights, Illinois: Waveland Press, Inc.
[24]. Jin, H.J., G. Cho and W.W. Koo. 2004. "Third-Country Effects on the Market Shares of U.S. Wheat in Asian Countries." *Journal of Agricultural and Applied Economics*. 36(3): 797-813.
[25]. Kandilov, I.V. 2008. "The Effects of Exchange Rate Volatility on Agricultural Trade." *American Journal of Agricultural Economics*. 90(4): 1028-43.
[26]. Langley, S.V., M. Giugale, W.H. Meyers and C., Hallahan. 2000. "International Financial Volatility and Agricultural Commodity Trade: A Primer." *American Journal of Agricultural Economics*. 82: 695-700.
[27]. McKenzie, M.D. 1999. "The Impact of Exchange Rate Volatility on International Trade Flows." *Journal of Economic Survey*. 13(1): 71-106.
[28]. Pesaran, M.H. and Y. Shin. 1999. An autoregressive distributed lag modelling approach to cointegration analysis. In: Strom, S. (Ed.), Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium, chapter 11, Cambridge University Press, Cambridge, MA.
[29]. Pesaran, M.H., Y. Shin. and R.J. Smith. 2001. "Bound Testing Appraoches to the Analysis of Level Relationships*." Journal of Applied Economics*. 16: 289-326.
[30]. Phillips, P.C.B. and B.E. Hansen 1990. "Statistical Inference in Instrumental Variables Regression with I(1) Process." *Review of Economic Studies*. 57: 99-125.
[31]. Reziti, I, and A. Ozanne. 1999 "Testing Regularity Properties in Static and Dynamic Duality Models: the Case of Greek Agriculture". *European Review of Agricultural Economics*. 26: 461-477.
[32]. Sauer, C. and A.K. Bohara. 2001. "Exchange Rate Volatility and Exports: Regional Differences between Developing and Industrialized Countries." *Review of International Economics*. 9(1): 133-152.

[33]. Solomon, H. and H.W. Kinnucan. 1993. "Effects of Non-Price Export Promotion: Some Evidence for Cotton." *Australian Journal of Agricultural Economics*. 37(1):1-15.
[34]. Wang, K.L. and C.B. Barrett. 2007. "A New Look at the Trade Volume Effects of Real Exchange Rate Risk." *Journal of Agricultural and Resource Economics*. 32(2): 225-255.
[35]. Warr, P. 2008. "The Transmission of import prices to domestic prices: an application to Indonesia". *Applied Economics Letters*. 15:499-503.

ENDNOTES

[a] In this paper, exchange rate uncertainty, risk and volatility have the same meaning.
[b] The OLS estimate results are available upon request.
[c] If the computed F-statistic falls inside of the critical bounds, the test result is inclusive.
[d] In the ARDL model, when the lag length is long, severe multicollinearity can occur. Thus the polynomial lag method is used in the regression to avoid this problem.
[e] This finding is consistent with the sample simulation results in the paper of Pesaran and Shin (1999). They reveal that the bias of FMOLS estimator is much bigger than that of ARDL estimator.