Three Essays on Antidumping Duties and Risk Factors Affecting International Seafood Trade

by

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Abstract

This dissertation includes three essays to address influences of the antidumping policy, price risk, and exchange risk in the US seafood import market. In the first chapter, I attempt to evaluate the effect of trade diversion on the effectiveness of antidumping duties. Antidumping duties can have the unintended consequence of increasing imports from countries not named in the dispute. Previous research suggests "trade diversion" significantly undermines the effectiveness of antidumping duties. However, domestic consumers could reduce their losses, as they can substitute out of the dutied good and into both the domestic good and the non-dutied good, which is, as might be expected, less expensive than the domestic good. Testing these propositions using the 2003 antidumping duty imposed by the United States on catfish imports from Vietnam, I find that the trade-diversion effect was significant in the sense that the quasi-rents of US producers enjoyed from the antidumping duty were reduced. The impact on consumer welfare was modest due to the great loss generated from the higher price of the dutied-good. An implication for public policy of the empirical results is that, in terms of the national welfare gain, it is optimal to lower the tariff rate when the amount of trade diversion is large.

The second chapter examines the impact of the export price risk and exchange-risk on the import demand. An extended Rotterdam model reveals that risk factors take effect on marginal utility via "adjusting" prices. This coincides with viewpoints in trade literature that risk-averse importers attach a risk premium as an extra mark-up to cover the cost of exchange risk (Balg and Metcalf, Bergin) and/or price risk (Wolak and Kolstad). The derived model further demonstrates that the trade effects of risk factors depend on own-price elasticities and substitutability between products within the same group. The modified specification makes testing restrictions on the effects of the risk variables plausible, resulting in a reduction in the parameter space. By further decomposing import price risk into export price risk and exchange-rate risk, the empirical model is applied to the US salmon import market. The results support the hypothesis that importers are sensitive to price and exchange-rate risks but reject the proposition that those two risk factors exert an identical effect on import demand.

The third chapter focuses on the difference between the dynamic and static specification for an import demand system inclusive of the price risk factor. I build risk factors into an Almost Ideal Demand System (AIDS) where the price risk may play a role by influencing baseline imports or import responsiveness to price. The risk-augmented AIDS model tests causality between price risk and trade. A multivariate GARCH approach estimates the conditional variances of prices. Nonstationarities in the data and endogeneity of price and price volatility are taken into account with Johansen's approach and a Vector Error Correction Model (VECM). The empirical results uncover a substantial role of risk factors in the US codfish import market. China's import share (CIF) increased 18% from 2004 to 2005, while Canada lost 9%. Holding other factors constant, high fluctuation of Canada's price would reduce its share by 69%, and raise China's share almost two times. China's relatively stable price further diminishes Canada's share by 4% and increases China's share by 9%.

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1.1 Introduction

An antidumping (AD) duty aims to assist domestic producers by raising price in the home market. Blonigen (2003) reports that AD cases worldwide increased from only a few in the 1970s to 2,200 in the 1990s. In the United States, calculated dumping margins between 1980 and 2000 rose from 15% to over 63%, and the probability of an affirmative ruling rose from 45% to over 60% (Blonigen 2006). Durling and Prusa (2006 p. 676) note that the range of products subjected to AD has been expanded, with measures frequently targeting agricultural products. Frequent enhancements of AD measures induce extensive disputes about the validity of this policy, which spurs persistent interest in the international trade literature. The most common consensus is that the targeted tariff is *a priori* inefficacy due to: (*i*) a large import demand elasticity (Kinnucan 2003, Kinnucan and Myrland 2006), (*ii*) the less competitive import-competing industry (Chang and Gayle 2006), and (*iii*) the trade-diversion effect (Prusa 2001, Gallaway et al. 1999). Hencefore, the AD duty will induce the non-targeted foreign suppliers to increase their shipments. The trade-diversion effect on AD duties is the primary focus of this paper.

The standard result in the economics literature is that trade diversion undermines the effectiveness of AD duties. However, as noted by Carter and Gallant-Trant (2010), most of the studies relate to industrial products, and there are good reasons to expect a different outcome for agricultural products. Unlike industrial products, agricultural products tend to be perishable, are produced seasonally, and must meet stringent quality and food safety standards before they can be admitted for importation. These factors, coupled with weak control over production and export decisions due to the biological and atomistic nature of agricultural production, constrain the extent to which non-targeted countries can fill the gap in the domestic market left by a particular antidumping action. Indeed, in their empirical investigation Carter and Gallant (2010, p. 119) find "a relatively small amount of trade diversion for agricultural products, and this effect does not extend beyond the year in which the case was initiated."

Although economists find little justification for AD duties because of the losses to consumers and downstream industries in the importing countries, the policy persists, filings have multiplied, and rulings have become more protectionist. Irwin (2005) reports that import-competing firms in the United States are increasingly turning to AD duties as an "easy" means to gain protection. A likely explanation is that the government distributes the revenue gained from AD duties to domestic producers alleging harm.[1] Therefore, the national welfare can be used to assess effectiveness of AD duties, which should also be influenced by trade diversion. Furthermore, part of what domestic producers gain is matched by the loss to domestic consumers. Trade diversion should benefit US consumers if the price of the non-dutied good is less expensive than the price of the protected good.

The purpose of this research is to determine the effects of trade diversion on the effectiveness of the antidumping duty that was imposed by the United States on catfish imports from Vietnam (US Department of Commerce 2003, 2009). The "Catfish War" generated national media attention, with articles appearing in *The New York Times*, *The Wall Street Journal*, *The Christian Science Monitor*, and *The Economist* discussing the policy and ethical dilemmas posed by the dispute.[2] For the purposes of this paper, catfish is a useful case study in that the trade diversion issue is clearly indicated in the data (table 1). Specifically, prior to the implementation

[1] In the US, the Continued Dumping and Subsidy Offset Act, commonly known as the "Byrd Amendment" (see Chang and Gayle 2006), was enacted by the US Congress in 2000 to provide for the distribution of antidumping duty revenues to petitioners. It was repealed in 2006, with a phase-out period for duty distribution.

[2] The *WSJ* article entitled "Catfish Case Muddies Waters for Bush 'Fast Track'" appeared 13 July 2001. For citations of the remaining articles and a good discussion of the issues involved, see Coleman (2005, pp. 6-8).

of antidumping duties ranged from 37% to 64% in 2003 US producers, in essence, owned the domestic market with a market share of 0.84. One year after the antidumping duty went into effect, China entered the US market, and within four years, saw its market share increase from 0.01 to 0.17. Although by 2008 Vietnam had re-established itself as an important competitor with a market share of 0.18, it seems clear that without the entry of China: *a*) the quasi-rents US producers enjoyed from the antidumping duty would have been higher, and *b*) the ability of the duty to protect US producers' market share would have been greater. Examining these hypotheses, I find that the entry of China did indeed undermine duty effectiveness. But the degree of the trade-diversion effect is primarily limited by the domestic industry's dominant market share. This "home bias" in an Armington (1969) framework implies an attenuated pass-through elasticity and weak cross-price effects for the domestic product.[3]

Previous research on the pass-through elasticity shows that the Armington substitution elasticity (σ) plays an important role in the extent to which a tariff-induced increase in the price of an imported good will raise the price of the protected good (Warr 2008). However, no study to our knowledge investigates the relationship between the market share, the pass-through elasticity, and trade-diversion effects. Thus, as a by-product of our analysis, I fill a gap in the literature relative to a key elasticity used to evaluate trade policy. A key finding is that the pass-through elasticity is inversely related to the market share of domestic producers. This "home bias" means a small market share of imports. I will further illustrate that the smaller market share of imports, the less important trade-diversion effects are apt to be.

We begin with the presentation of the structural model. Analytical expressions for the pass-through elasticity and the trade-diversion effect are then derived. The welfare effects of the

[3] According to Whalley and Xin (2009, p. 309), home bias is commonly defined in the literature as "an Armington preference for domestic over comparable foreign products in a trade model where goods are heterogeneous across countries, a definition I adopt in this study.

catfish antidumping duty are then measured, with special attention to trade-diversion effects. The paper concludes with a summary of the key findings.

1.2 Model

Consider a country that imports two products, Q_2 and Q_3, differentiated by source origin. The imported products compete with a home good, Q_1. The three goods are weakly separable from all other goods in the home market, and consumers allocate income in accordance with the two-stage budgeting hypothesis. The home country imposes an antidumping duty on Q_2 defined as follows:

(1) $$\tilde{P}_2 = P_2 \cdot T$$

where $T \geq 1$ is the duty expressed in "iceberg" or proportionate form, $\widetilde{P}_2$ is price inclusive of the duty, hereafter "demand price", and P_2 is price exclusive of the duty, hereafter "supply price." Prices for the three goods are determined under competitive conditions.

Given these assumptions, an issue is the effect of the duty on the price of the domestic good when the presence of the non-dutied good, Q_3, is taken into account. To determine that, let the demand and supply equations for the three goods be defined as follows:

(2) – (4) $$Q_i^* = \eta_{i1} P_1^* + \eta_{i2} \widetilde{P}_2^* + \eta_{i3} P_3^* + \gamma_i X^* \qquad i = 1,2,3$$

(5) – (7) $$Q_i^* = \varepsilon_i P_i^* \qquad i = 1,2,3$$

where the asterisk denotes proportionate change. Thus, for example, X^* is the proportional change in X where $X = (P_1 Q_2 + \tilde{P}_2 Q_2 + P_3 Q_3)$ is total expenditure on the three goods by domestic consumers. In this model, the η_{ij} are conditional price elasticities of demand; the γ_i are conditional expenditure elasticities; and the ε_i are price elasticities of supply. The demand curves are downward sloping ($\eta_{ii} < 0$); the supply curves are non-decreasing ($\varepsilon_i \geq 0$); the goods are non-inferior ($\gamma_i \geq 0$) and substitute for one another in consumption ($\eta_{ij} > 0$ for $i \neq j$).

An increase in the duty shifts the supply curve for the dutied good to the left. To see this, write equation (1) in proportionate change form:

(8) $\tilde{P}_2^* = P_2^* + T^*$.

Substituting equation (8) into equation (6) yields the tax-burdened supply curve:

(6′) $Q_2^* = \varepsilon_2 \tilde{P}_2^* - \varepsilon_2 T^*$.

Holding constant the demand price, a 1% increase in the AD shifts the supply curve for the dutied good to the left from the initial market equilibrium point by an amount equal in proportionate terms to the dutied good's supply elasticity.

Equation (6) shows that an increase in the duty has opposite effects on the demand and supply prices. Thus, for example, if the supply of the dutied good is perfectly inelastic ($\varepsilon_2 = 0$), foreign producers bear the full incidence and $P_2^* = -T^*$. In this instance, the duty is ineffective, as it has no effect on prices in the domestic market. Thus, a necessary condition for the duty to be effective is that $\varepsilon_2 > 0$, an issue to which I shall return.

Equations (2) - (8) contain seven endogenous variables ($P_1^*, P_2^*, P_3^*, \tilde{P}_2^*, Q_1^*, Q_2^*, Q_3^*$) and two exogenous variables (X^* and T^*). Exogenous variables that affect demand and supply other than consumer expenditure and the duty are suppressed.

Tariff Pass-Through Elasticity (PTE)

Warr states (2008, p. 499) "The manner in which the landed price of imports affects domestic prices is central to trade policy analysis." The tariff pass-through elasticity (PTE) when an AD is used to raise the import price may be defined as follows:

(9) $$\frac{P_1^*}{T^*} = \left(\frac{P_1^*}{\tilde{P}_2^*}\right)\left(\frac{\tilde{P}_2^*}{T^*}\right)$$

where the first term in parenthesis indicates the *price* pass-through elasticity and the second term indicates duty incidence. In Warr's analysis the supply of the imported good is assumed to be perfectly elastic ($\varepsilon_2 = \infty$), which implies the full incidence of the duty is borne by domestic consumers, i.e., $\tilde{P}_2^*/T^* = 1$. Here I relax that assumption and add trade-diversion effects to develop a more complete expression for the PTE.

The PTE when third party effects are considered are (see appendix for derivation):[4]

(10) $$\left(\frac{P_1^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2\left(\eta_{12}(\varepsilon_3 - \eta_{33}) + \eta_{13}\eta_{32}\right)}{\Delta}$$

(11) $$\left(\frac{\tilde{P}_2^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2\left((\varepsilon_1 - \eta_{11})(\varepsilon_3 - \eta_{33}) + \eta_{13}\eta_{31}\right)}{\Delta}$$

(12) $$\left(\frac{P_3^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2\left(\eta_{32}(\varepsilon_1 - \eta_{11}) + \eta_{12}\eta_{31}\right)}{\Delta}$$

where $\Delta = (\varepsilon_1 - \eta_{11})(\varepsilon_2 - \eta_{22})(\varepsilon_3 - \eta_{33}) - \eta_{32}\eta_{23}(\varepsilon_1 - \eta_{11}) - \eta_{13}\eta_{31}(\varepsilon_2 - \eta_{22}) - \eta_{21}\eta_{12}(\varepsilon_3 - \eta_{33}) - \eta_{21}\eta_{13}\eta_{32} - \eta_{31}\eta_{12}\eta_{23}$, and equation (11) is trade incidence borne by domestic consumers.

After presuming a downward-sloping demand curve and an upward-sloping supply curve for each product in the market, I can justify the common denominator (Δ) in (10) - (12) is positive. Noted in the Armington framework, any pair of products in the group is a gross substitute ($\eta_{ij} > 0$ for $i \neq j$); I can further observe each numerator in (10) - (12) with a positive sign as well. This indicates that the AD duty imposed on the subject country should increase

[4] PTE in general means the tariff pass-through elasticity of *the domestic good's price*. In order to elaborate the relative price movements, I also calculate the PTE of the dutied good's domestic price and the non-dutied good's price. Except stated otherwise, in the paper, PTE only refers to price transmission between tariff and the price of the protected good.

domestic prices of the protected good, the dutied product, and the not-dutied good. An increase in the price of the dutied product is induced by an inward supply shift, whereas a rise in the prices of the protected and non-dutied products is induced by upward demand shifts.

In order to demonstrate the tariff pass-through process, equation (10) is restated as follows:

$$(13) \qquad \left(\frac{P_1^*}{T^*}\right)^{N=3} = \left(\frac{\eta_{12}(\varepsilon_3 - \eta_{33}) + \eta_{13}\eta_{32}}{(\varepsilon_1 - \eta_{11})(\varepsilon_3 - \eta_{33}) - \eta_{13}\eta_{31}}\right)\left(\frac{\tilde{P}_2^*}{T^*}\right)^{N=3}$$

where the first term is *price* pass-through elasticity, which is primarily determined by substitution between goods from different sources. In (10) or (13), the supply elasticity of the dutied good is a critical determinant of duty effectiveness. For example, if the supply of the dutied good is perfectly inelastic ($\varepsilon_2 = 0$), as might be true in a short-run situation, the entire incidence of the duty is borne by producers in the exporting country and the protected good's PTE = 0. This is true regardless of the size of the price pass-through elasticity (the first term in equation (13)) and whether trade-diversion effects are considered. Alternatively, if supply of the dutied good is perfectly elastic, equation (13) is reduced to:

$$(14) \qquad \left(\frac{P_1^*}{T^*}\right)^{N=3}_{\varepsilon_2=\infty} = \frac{\eta_{12}(\varepsilon_3 - \eta_{33}) + \eta_{13}\eta_{32}}{(\varepsilon_1 - \eta_{11})(\varepsilon_3 - \eta_{33}) - \eta_{13}\eta_{31}}$$

The corresponding expression when trade-diversion effects are ignored is:

$$(15) \qquad \left(\frac{P_1^*}{T^*}\right)^{N=2}_{\varepsilon_2=\infty} = \frac{\eta_{12}}{\varepsilon_1 - \eta_{11}}$$

Equations (14) and (15) represent the maximal effect of the duty on the price of the protected good. Equation (15) is identical in form to the PPE derived by Warr (2008, p. 500, equation (6)). That this elasticity has an upper limit of one can be seen by imposing homogeneity to yield:

(16) $$\left(\frac{P_1^*}{T^*}\right)_{\varepsilon_2=\infty}^{N=2} = \frac{\eta_{12}}{\varepsilon_1 + \eta_{12} + \gamma_1}.$$

In this instance, the PTE increases as the protected good becomes a closer substitute for the dutied good (larger η_{12}), as the supply of the protected good becomes less price elastic (smaller ε_1), and as the demand of the protected good becomes less expenditure elastic (smaller γ_1).

Trade-Diversion Effect

An analytical expression for the trade-diversion effect (TDE) can be obtained by subtracting equation (14) from equation (15) to yield:

(17) $$TDE = \frac{-\eta_{13}\left(\eta_{32}(\varepsilon_1 - \eta_{11}) + \eta_{31}\eta_{12}\right)}{(\varepsilon_1 - \eta_{11})\left((\varepsilon_1 - \eta_{11})(\varepsilon_3 - \eta_{33}) - \eta_{13}\eta_{31}\right)}.$$

Equation (17) is the TDE when domestic consumers bear the full incidence of the duty and thus represents the maximal effect. With this caveat in mind, the TDE is negative in sign, which means ignoring the TDE will cause the duty's effect on domestic price to be overstated. The degree of overstatement increases as: *i*) the protected good becomes a closer substitute for the non-dutied good (larger η_{13}), *ii*) the non-dutied good becomes a closer substitute for the dutied good (larger η_{32}), and *iii*) the supply of the non-dutied good becomes less price elastic (smaller ε_3). If $\varepsilon_3 = \infty$, the duty has no effect on the price of the non-dutied good and TDE = 0. Thus, the extent to which the presence of a third good undermines the effectiveness of an AD depends on substitution effects, but also on supply conditions in the market for the non-dutied good. If consumers respond to the duty by switching to the non-dutied good to a greater extent than to the protected good ($\eta_{32} > \eta_{12}$), as might be expected if the non-dutied good is less expensive than the protected good ($P_3 < P_1$), the attenuation of the PTE caused by the TDE could be empirically important.

In the Armington model, consumer preferences are assumed to be homothetic, indicating trade patterns changes only with relative price movements. Since an AD duty increases the domestic price of each good in the group, the magnitudes of the growth rate should illustrate redistribution of shares among suppliers, and consequently reveal the mechanisms of the trade-diversion effect. The differences between PTE are given by:

(18) $$\left(\frac{\tilde{P}_2^*}{T^*}\right)^{N=3} - \left(\frac{P_1^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2}{\Delta}\left((\varepsilon_1 - \eta_{11} - \eta_{12})(\varepsilon_3 - \eta_{33}) + \eta_{13}(\eta_{31} - \eta_{32})\right)$$

(19) $$\left(\frac{P_3^*}{T^*}\right)^{N=3} - \left(\frac{P_1^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2}{\Delta}\left(\eta_{32}(\varepsilon_1 - \eta_{11} - \eta_{13}) + \eta_{12}\left(\eta_{31} - \varepsilon_3 + \eta_{33}\right)\right)$$

(20) $$\left(\frac{\tilde{P}_2^*}{T^*}\right)^{N=3} - \left(\frac{P_3^*}{T^*}\right)^{N=3} = \frac{\varepsilon_2}{\Delta}\left((\varepsilon_1 - \eta_{11})(\varepsilon_3 - \eta_{33} - \eta_{32}) + \eta_{31}(\eta_{13} - \eta_{12})\right)$$

For equation (18), the increase of the domestic price of the dutied good is greater than the growth of the price of the protected good, on the *sufficient* conditions that (1) $|\eta_{11}| > |\eta_{12}|$ which is not a particularly onerous condition and (2) the non-dutied good is a closer substitute to the protected good than to the dutied good, i.e. $\eta_{31} > \eta_{32}$ (as shown later, it is always true in the Armington model given the US produces dominate the market). Considering the formula of demand elasticities, equation (19) can be simplified to equal $\frac{\varepsilon_2}{\Delta}\eta_{32}(\varepsilon_1 - \varepsilon_3)$, indicating the growth of the non-dutied good price is less than the growth of the protected good price if $\varepsilon_1 < \varepsilon_3$. Although signs of (18) and (19) depend on the reasonable pre-conditions, I can tell the sign of (20) is unequivocally positive, indicating an improved comparative advantage of the non-dutied good compared to the dutied good. In general, implementation of an AD duty will make the relative price movements favor both the protected good and the non-dutied good over the dutied good, resulting in trade diversion.

1.3 Parameterization

To assess the importance of the trade-diversion effect for the catfish AD, I need to set parameters in the structural model. For this purpose, I first follow Warr (2008) and restrict demand elasticities to conform to the "Armington assumption" (Armington 1969). In the Armington's framework, imports from different sources and domestic production are assumed to be imperfectly substitutable. The expenditure of a particular group is allocated to different suppliers based on relative price movements. This implies the demand elasticities matrix is as follows:

$$(21) \qquad \mathbf{N} = \begin{bmatrix} \eta_{11} & \eta_{12} & \eta_{13} \\ \eta_{21} & \eta_{22} & \eta_{23} \\ \eta_{31} & \eta_{32} & \eta_{33} \end{bmatrix} = \begin{bmatrix} S_1\eta - (1-S_1)\sigma & S_2(\sigma+\eta) & S_3(\sigma+\eta) \\ S_1(\sigma+\eta) & S_2\eta - (1-S_2)\sigma & S_3(\sigma+\eta) \\ S_1(\sigma+\eta) & S_2(\sigma+\eta) & S_3\eta - (1-S_3)\sigma \end{bmatrix}$$

where η = -1 is the price elasticity of demand for the three goods combined, and $\sigma > 1$ is the Armington substitution elasticity. With the maintained hypothesis that the three goods form a weakly separable group, the conditional budget shares sum to one. One advantage of the Armington framework is that the number of parameters to derive the 9 demand elasticity is reduced to three, i.e., σ, S_1, and S_2, noting that $S_1 = 1 - S_2 - S_3$.

Muhammad *et al*. (2010, p. 437) estimate the (long-run) own-price elasticity of demand for US catfish to be $\hat{\eta}_{11} = -1.42$. The sample period of the data used in Muhammad et al. is from January 1993 to December 2007. Substituting this value and S_1 = 0.83 (mean value for 1999-2007 in our sample, close to the period employed in Muhammad *et al.*) into the expression for η_{11} in equation (21), and solving for the substitution elasticity, yield $\hat{\sigma} = 3.37$. Considering the decline in S_1 during the data period, our "best-bet" estimate is $\hat{\sigma} = 2.5$. However, to gauge the sensitivity of results to substitutability between products in the same group, I also conduct alternative estimates of $\hat{\sigma} = 1.5$ and $\hat{\sigma} = 5$. Although the lower bound is outside the typical range (from 2 to 5) for Armington elasticity assumed in applied equilibrium models (Warr 2008), it is used to give the simulation results full play.

Muhammad *et al*. (2010, p. 437) estimate the long-run supply elasticity for domestic catfish to be 1.05. So, I set the "best-bet" value for $\hat{\varepsilon}_1 = 1.1$. Since Muhammad *et al*. did not estimate supply elasticities for imported catfish, I set $\hat{\varepsilon}_2$ and $\hat{\varepsilon}_3$ to 2, the value used in Kinnucan's (2003, p. 216) analysis. Also, to gauge the sensitivity of the results to import supply response, I conducted additional simulations with $\hat{\varepsilon}_2$ and $\hat{\varepsilon}_3$ set alternatively to 4 and infinity.

As discussed before, the final determination of the AD measure was issued against the targeted catfish in 2003, and this AD case remained in effect in 2009. Accordingly, I set three baseline values from the data sample sub-periods: 2002-04, 2005-07, and 2008-10. When simulating, market shares, prices and quantities are mean values in each sub-period:

Year	Price (CIF $/lb.)			Quantity (mil. lbs)			Market Share		
	P_1	P_2	P_3	Q_1	Q_2	Q_3	S_1	S_2	S_3
2002-2004	2.47	1.46	1.45	311	63	2	0.83	0.17	0.01
2005-2007	2.84	1.54	1.69	327	63	47	0.75	0.14	0.11
2008-2010	2.94	1.55	1.76	291	127	67	0.60	0.26	0.14

In 2002-04, on average, US catfish producer dominated the domestic market with a market share of 0.83, and China owned a mere share (S_3 = 0.01). After the implementation of the AD, China saw its market share rise to 0.11 at the expense of US producers in 2005-07; however, Vietnam only lost 0.03 shares. In 2008-10, Vietnam re-established itself as an important supplier with a market share of 0.26, whereas China only gained a growth of 0.03. A comparison of simulating results with the baseline values from those sub-periods should shed light on the trade-diversion effect.

Given η = -1 and the "best-bet" value of $\hat{\sigma} = 2.50$, the demand elasticities matrices (equation (21)) corresponding to the three sets of baseline market shares are:

(22) $$N = \begin{bmatrix} -1.26 & 0.25 & 0.01 \\ 1.24 & -2.25 & 0.01 \\ 1.24 & 0.25 & -2.49 \end{bmatrix}$$ (2002-04: $S_1 = 0.83$, $S_2 = 0.17$, $S_3 = 0.01$)

(23) $$N = \begin{bmatrix} -1.38 & 0.22 & 0.16 \\ 1.12 & -2.28 & 0.16 \\ 1.12 & 0.22 & -2.34 \end{bmatrix}$$ (2005-07: $S_1 = 0.75$, $S_2 = 0.14$, $S_3 = 0.11$)

(24) $$N = \begin{bmatrix} -1.60 & 0.39 & 0.21 \\ 0.90 & -2.11 & 0.21 \\ 0.90 & 0.39 & -2.29 \end{bmatrix}$$ (2008-10: $S_1 = 0.60$, $S_2 = 0.26$, $S_3 = 0.14$)

The most important elements in matrix (22) - (24) are η_{12} and η_{13}, as these elasticities are primary determinants of tariff pass-through elasticities and trade-diversion effects (see equations (10) and (17)). Since neither elasticity is very large in relation to the own-price elasticity η_{11}, I suggest that: *a*) the duty may not be very effective at raising the price of the domestic good, and *b*) trade-diversion effects may not be very important.

1.4 Simulated Tariff Pass-Through Elasticities (PTE)

In order to simulate tariff pass-through elasticities (PTE), I build up scenarios using market shares from 2002-04, 2005-07, and 2008-10, respectively. Armington elasticity ($\hat{\sigma}$) is set alternatively to 2 ("best-bet" value), 4, and 6; the supply elasticity vector is $\boldsymbol{\varepsilon}' = (1.1, 2.0, 2.0)$, our "best-bet" estimates of these parameters.

Results suggest the AD duty does not substantially affect the price of the protected good (table 2). For the considered parameter values, the PTE of the protected good (P_1^* / T^*) ranges from 0.019 to 0.13. This suggests that if *all* imports from Vietnam were assessed at 63.88%, the highest AD margin calculated by the US Department of Commerce (2003, 2009), the price of domestic catfish would rise by *at most* 8.3% (0.13 × 63.88%). This estimate assumes Armington elasticity is 5 (the upper limit) in the scenario with S_3 = 0.14. In the scenarios with S_3 = 0.01 and S_3 = 0.11, the maximum increase of the domestic catfish price is 5.6% and 4.7%, respectively.

The level of trade diversion can be evaluated by a comparison of the simulated tariff pass-through elasticities. In each case, the PTE of the domestic catfish and the non-dutied catfish are much smaller than the corresponding PTE of Vietnam's catfish. As an example, in 2002-04, for $\hat{\sigma} = 2.5$ ("best-bet" value), the domestic price of Vietnam's catfish increases by 49%; however, the prices of the protected catfish and China's catfish only increase by 5.2% and 4.1%, respectively. Actually in each case, the PTE of China's catfish is even smaller than the PTE of the protected catfish, although the differences are not substantially significant. Changes in the relative prices imply that the market share extracted from Vietnam's producers due to the AD policy is shared by the US and China, noting in the Armington framework, relative market shares are associated with relative prices through Armington elasticity.[5] The upshot is that, the market share and quasi-rents US producers enjoyed from the AD policy would have been greater without the entry of China. On the other hand, for US consumers, switching from the domestic catfish to China's catfish should increase their benefits, considering the domestically produced

[5] Our simulated results show that the market share of the non-named country is increased by 0.03% in the 2002-04 scenario, 0.5% in the 2005-07 scenario, and 1.2% in the 2008-10 scenario. Taking the initial market share of the non-dutied good into account (0.01, 0.11, and 0.14, respectively), this confirmed a positive relationship between the initial market share and the degree of trade diversion.

catfish is generally more expensive than China's catfish (table 1). Next, I evaluate the magnitudes of the welfare effects of trade diversion for all parties in the market.

1.5 Welfare Analysis

The implications of trade-diversion effects from a welfare perspective are evaluated in a partial-equilibrium setting using figure 1 as a guide. Consistent with the structural model, the diagrams depict a situation where the three goods are substitutes in consumption, but independent in production. That is, the technologies used to produce the domestic and imported goods are independent, and any specialized factors used to produce the three goods (e.g., catfish feed) are perfectly elastically supplied to each country. The latter assumption is necessary for welfare effects to have significance in the sense that they can be traced to an identifiable group (e.g., domestic producers, see Thurman 1993).

The AD shifts the supply curve for the dutied good up from S to S' in panel B by an amount equal to the *per-unit* duty, i.e., the *ad valorem* duty multiplied by the price of the dutied good in the pre-duty equilibrium. After the three markets have adjusted fully to the supply shift, the market price of the dutied good increases from P_2 to P_2', and the supply price of the dutied good decreases from P_2 to P_S. The full welfare effect is measured in the market for the dutied good (Just, Heuth, and Schmitz 2004, pp. 322-26). Specifically, the hatched area between the price lines P_2 to P_S and behind the supply curve S, labeled ΔPS_2 in panel B, represents the welfare loss to producers of the dutied good. The hatched area between the price lines P_2 to P_2' and behind the equilibrium or "total" demand curve D_T, labeled ΔCS^* in Panel B, represents the welfare effect of the supply shift on the remaining parties affected by the duty. Specifically, $\Delta CS^* = \Delta CS_1 + \Delta CS_2 + \Delta CS_3 + \Delta PS_1 + \Delta PS_3$ represents the welfare loss to domestic consumers as a result of the higher prices they must pay for domestic and imported catfish following imposition of the duty. Since the duty raises the market price of all three goods, producers of the domestic and non-dutied imported goods enjoy a welfare gain equal to the shaded areas labeled ΔPS_1 and ΔPS_3, respectively, in panels A and C.

Since a parallel upward shift in a supply curve always decreases consumer surplus, $\Delta CS^* < 0$, which implies that consumer losses outweigh producer gains. However, the duty provides tax revenue, which, depending on duty incidence, may be sufficient to compensate consumers and yield a net welfare gain for the US.

With the maintained hypothesis that supply and demand shifts are parallel, the welfare effects of the catfish duty may be approximated using the following formulas:

(25) $$\Delta PS_i = P_i Q_i P_i^* \left(1 + 0.5 Q_i^*\right) \qquad i = 1,2,3$$ (US and foreign producer impacts)

(26) $$\Delta CS^* = -P_2 Q_2 \widetilde{P}_2^* \left(1 + 0.5 Q_2^*\right)$$ ("Consumer" impact in dutied market)

(27) $$\Delta CS = \Delta CS^* - \Delta PS_1 - \Delta PS_3$$ (US consumer impact)

(28) $$\Delta TR = P_2 Q_2 \left(\widetilde{P}_2^* - P_2^*\right)\left(1 + Q_2^*\right)$$ (US treasury impact)

The P_iQ_i represents the value of the ith product in the initial equilibrium. Numerical values for the asterisked variables in equations (25) - (28) were computed by simulating the model (equations (2) – (8)) for a 35% increase in the duty. In January 2009, the US Department of Commerce ruled that catfish duties imposed in 2003 would remain in place (Martin 2009). The *de minimis* weighted-average antidumping margin in both the 2003 and 2009 USDC rulings is 36.84% (US Department of Commerce 2003, 2009). Thus, setting $T^* = 0.35$ provides a conservative estimate of welfare impact.

Distribution effects

Our first set of simulations focuses on the welfare implications of trade diversion for all parties (distribution effects). I chose the 2008-10 baseline values for prices and quantities. Simulations were run with the market share for the non-dutied good (S_3) alternatively to be 0.07, 0.14 (mean value for 2008-10), and 0.28 to mimic the observed range in this parameter over the 2002-10 period. Although the latter value is outside the range for S_3 reported in table 1, it is used to give trade diversion full play. Accordingly, the market share of the protected good (S_1) is alternatively set to be 0.67, 0.60, and 0.46, since the market share of the dutied good (S_2) is held constant at 0.26 (mean value for 2008-10). In each instance, the demand elasticities matrix is adjusted as required to ensure that demand elasticities are consistent with theory. Other parameters are set to be their "best-bet" values, the domestic supply elasticity $\hat{\varepsilon}_1 = 1.1$, the import supply elasticities $\hat{\varepsilon} = (\hat{\varepsilon}_2, \hat{\varepsilon}_3) = 2$, and Armington elasticity $\hat{\sigma} = 2.5$.

The simulating results are reported in table 3. Focusing first on the case where $S_3 = 0.14$, results indicate the largest beneficiary of the duty is the US treasury, which gains $29.2 million. Although domestic producers gain $8.1 million, this comes at the expense of domestic consumers, who lose $28.3 million. Adding together these effects yields a net gain to the domestic economy of $9.0 million. A gain occurs because import supply is sufficiently upward-sloping to permit the US to act as a monopsonist and price discriminate against foreign suppliers via the imposition of a tariff (Enke 1944). Specifically, for the considered parameter values, domestic consumers bear 51% of the duty, which means 49% of the US treasury gain comes from Vietnam producers. Vietnam producers lose $18.1 million, while China producers gain $1.5 million, for a net welfare loss to all affected parties of $7.7 million. The upshot is that the AD did more to punish foreign producers than to reward domestic producers, a common result in the AD literature for farm products (e.g., Asche 2001; Brester et al. 2002; Kinnucan 2003; Kinnucan and Myrland 2006).

Turning to trade-diversion effects, rent dissipation associated with trade diversion increases with the market size of the dutied good, but the increase is modest. An increase in the market share of the non-dutied good from 0.07 to 0.14 causes the domestic producer surplus to decline by 12%. The corresponding decline is 33%, when the market share of the non-dutied good is increased by 4 times from 0.07 to 0.28. Thus, for the considered parameter values, trade diversion has a moderate effect on rent dissipation. Trade diversion has an even less important effect on consumer surplus. When the market share of the non-dutied good increases from 0.07 to 0.28, consumers' loss is dwindled by a mere 3.2%, too small to matter. Although domestic consumers switch to the less expensive China's catfish when the domestic price of the protected good rises, their benefits from the switch are diluted by the huge loss from the rising price of the dutied-good due to an inward supply shift. This is consistent with the simulated tariff pass-through elasticities (PTE). As shown in table 2, in each scenario, the PTE of the protected good is much greater than the corresponding PTE of the protected good and the non-dutied good in magnitude.

Linking producer surplus to consumer surplus, I calculate the value of the redistribution efficiency of the AD duty, defined as the ratio of domestic producers' gain to domestic consumers' loss (in absolute value). Little of the lost quasi-rent rebounds to the benefit of domestic consumers, noting that the redistribution efficiency is reduced from 0.32 to 0.22 when S_3 is increased from 0.07 to 0.28. The upshot is that, for the considered parameter values, the trade-diversion effects are confined largely to domestic producer impacts, and even then they are less important than other factors that affect duty efficacy, namely substitution effects and import supply elasticities, as shown in our following set of simulations.

Sensitivity Analysis

The foregoing results are based on the "best-bet" values of Armington elasticity and the import supply elasticities. To gauge the sensitivity of results to the elasticity of substitution, I set $\hat{\sigma}$ alternatively to 1.5, 2.5 ("best-bet" value) and 5, and the import supply elasticities $\hat{\varepsilon} = (\hat{\varepsilon}_2, \hat{\varepsilon}_3)$ alternatively to 2 ("best-bet" value), 4 and ∞. In these simulations, market shares are set to their average values for 2008-10 (**S** = (0.62, 0.24, 0.14)), the domestic supply elasticity to its "best-bet" value $\hat{\varepsilon}_1 = 1.1$. The demand elasticity matrices corresponding to the different Armington elasticities are:

$$(29) \qquad N = \begin{bmatrix} -1.20 & 0.13 & 0.07 \\ 0.30 & -1.37 & 0.07 \\ 0.30 & 0.13 & -1.43 \end{bmatrix} \qquad (\sigma = 1.5)$$

$$(30) \qquad N = \begin{bmatrix} -1.60 & 0.39 & 0.21 \\ 0.90 & -2.11 & 0.21 \\ 0.90 & 0.39 & -2.29 \end{bmatrix} \qquad (\sigma = 2.5)$$

$$(31) \qquad N = \begin{bmatrix} -2.61 & 1.05 & 0.56 \\ 2.39 & -3.95 & 0.56 \\ 2.39 & 1.05 & -4.44 \end{bmatrix} \qquad (\sigma = 5.0)$$

Results suggest Armington elasticity and import supply elasticities are pivotal in determining both the size of the welfare impacts and their distributional consequences (table 4). An increase in Armington elasticity ($\hat{\sigma}$) has more substantial effects on US producers than on other parties. In the instance where $\hat{\varepsilon} = (\hat{\varepsilon}_2, \hat{\varepsilon}_3) = 2$, the US producer surplus is increased 1.3

times when $\hat{\sigma}$ is raised from 1.5 to 2.5. The corresponding price increase falls to 67% when $\hat{\sigma}$ is increased from 2 to 5. Import supply elasticities also have a substantial impact on producer surplus. Using $\hat{\varepsilon} = (\hat{\varepsilon}_2, \hat{\varepsilon}_3) = 2$ as the standard of comparison, if import supply is perfectly elastic, the benefits of the duty for domestic producers are overstated by a factor of 1.6 (= \$5.85 million / \$3.56 million) when $\hat{\sigma} = 1.5$, and by a factor of 2.2 (= \$30.3 million / \$13.5 million) when $\hat{\sigma} = 5$. Thus, the erroneous treatment of import prices as exogenous causes a significant upward bias in the estimated gain to domestic producers, and this bias increases as the goods become closer substitutes.[6]

Turning to distributional impacts, the national welfare gain from the duty converts to a loss for an import supply elasticity as small as 4.0. If $\hat{\varepsilon} = \infty$, the national welfare gain from a 35% duty is between -\$10.5 million and -\$25.5 million, with the latter estimate corresponding to the upper limit of the substitution elasticity ($\hat{\sigma} = 5$). When $\hat{\varepsilon} = \infty$, domestic consumers bear the full brunt of the duty, with losses ranging from -\$39.8 million to -\$49.3 million. Above all, if import supplies are perfectly elastic, the duty in essence represents a transfer from domestic consumers to domestic producers and the US Treasury, and the incidence of the duty is shifted entirely to the US consumers, indicating a null effect of AD on the welfare of foreign producers. For the redistributive efficiency of the duty (domestic producer gain / domestic consumer loss), it increases as import supplies become more price elastic, and as the goods become closer substitutes. For the considered parameter values, this ratio never exceeds 0.61 (= \$30.3 million / \$49.3 million); for "best-bet" parameter values, the ratio is 0.29 (= \$8.11 million / \$28.3 million).

[6] The study by Warr (2008) treats import price as exogenous when developing an analytical expression for the pass-through elasticity, while the study by Muhammad *et al*. (2010) treats import price as exogenous when evaluating the welfare impact of the catfish duty.

Tariff Rate and Trade Diversion

Since trade diversion makes the AD policy less effective in terms of producer surplus, net national welfare, and redistribution efficiency, a low tariff rate (i.e. AD margin) may be proposed to reduce trade diversion because it limits the benefits for the non-named countries to increase their exports to the US.[7] But, on the other hand, a low tariff rate would directly reduce the revenue gain for the government and reduce the extent to which the price of the protected good rises, provided that the tariff pass-through elasticity is constant. To examine the correlation between the tariff rate and trade diversion, I recomputed the welfare distribution with alternative AD rates. After implementing in 2003, the rate of the AD duty against Vietnam's catfish ranges from 37% to 64%. Thus, I set the tariff rate with alternative 15%, 35%, 50%, and 70%, where 50% is equal to the inverse of the "best-bet" import supply elasticity. The supply elasticities and the Armington elasticity are set to their "best-bet" values $\hat{\varepsilon} = (1.1, 2, 2)$ and $\hat{\sigma} = 2$ to permit isolation of the market-share effect. Prices and quantities are mean values for 2008-10. The market share of the non-dutied good is increased from 0.07 to 0.28 in three steps.

Results indicate producer surplus and consumer surplus (in absolute value) are increasing functions of the tariff rate (table 5). When the tariff rate is increased, consumers' loss from the increasing price of the dutied good overwhelms the positive effects of the demand outward shift of the protected good and the non-dutied good. In the instance where $S_3 = 0.14$, using $T^* = 35\%$ as the benchmark, if the tariff rate is increased to 50%, gains to domestic producers and the US treasury are improved by $3.6 million and $3.1 million, respectively; however, the loss for domestic consumers is also increased by $ 9.9 million, resulting in a decline in net national

[7] Konings, *et al*. (2001) relate the lower amount of import diversion in Europe to lower duty levels, which limit the benefits of protection for the non-named countries.

welfare (ΔW_{us}). If the tariff rate is raised to 70%, the huge negative effect on domestic consumers results in a negative net national welfare.

Finally, I focus on the national welfare gain to justify trade-diversion effects on the tariff rate. In each scenario, the national welfare gain (ΔW_{us} in table 5) is not a monotonic function of the tariff rate. In the instance where S_3 = 0.14, the national welfare is raised by 35% when the tariff rate is increased from 15% to 35%, but it is reduced by 13% when the tariff rate keeps rising to 50%. Moreover, the national welfare becomes negative when the tariff rate is set at 70%. The simulated results further demonstrate the tariff rate of 35% seems to be the "optimal tariff" in each case. In the cases with T^* = 35%, the national gain is \$9.7 million when S_3 = 0.7, \$9.0 million when S_3 = 0.14, and \$7.5 million when S_3 = 0.28. The negative relationship between the maximal national welfare and the market share of the non-dutied good coincides with the previous findings that the trade-diversion effect has a stronger consequence of dwindling the domestic producer surplus than increasing consumer surplus. In the instance with T^* = 70%, the national welfare *loss* is \$3.5 million when S_2 = 0.7, and \$8.4 million when S_2 = 0.28. This implies that the break-even point of the tariff rate at which ΔW_{us} = 0 is negatively related to the extent of trade diversion. In other words, the greater trade diversion, the faster the national welfare converges to zero when the tariff rate is growing. The upshot is that the tariff rate tends to be lower when the extent of trade diversion is greater in order to obtain an "optimal" gain in national welfare or avoid a nil or negative welfare effect.

1.6 Concluding Comments

Despite the tariff rate as high as 64%, the antidumping duty the United States imposed in 2003 on catfish imports from Vietnam failed to prevent a precipitous decline in domestic industry market share (from 84% in 2003 to 58% in 2010). It is tempting to ascribe AD ineffectiveness to the entry of China as a new foreign competitor in 2004, as consumers substitute out of the dutied good and into *both* the domestic good and the non-dutied good. This switch may also reduce the loss to consumers, considering that the non-dutied good is generally less expensive than the protected good. Study results suggest trade diversion is an epiphenomenon. Specifically, China's ability to capture 28% of the domestic market following AD implementation is estimated to have eroded the quasi-rents that domestic producers received by at most 33%. On the contrary, trade diversion cannot reduce consumers' loss substantially since the pass-through elasticity (PTE) of the dutied good (related to a supply shift) is much greater than the PTE of the protected good and the non-dutied good (related to demand shifts). The simulated duty incidence suggests, for the considered parameter values, about one half of the duty appeared as a rise in the US price of the dutied product.

Further, the extent to which trade diversion influences the effectiveness of an AD duty depends on the market share of the domestic good. In an Armington framework, a dominant market share for the domestic good, or "home bias," means that the cross-price elasticities of the protected good tend to be small. As an example, in 2005 when US producers enjoyed a market share of 0.83, the cross-price elasticities in question consistent with an Armington elasticity of 2.5 are $\eta_{12} = 0.20$ and $\eta_{13} = 0.06$.[8] Even in 2010 when US producers still dominated the market but with a smaller share of 0.58, the corresponding cross-price elasticities are $\eta_{12} = 0.87$ and $\eta_{13} = 0.15$. This indicates: (*i*) the effect of an AD on the price of the protected good (tariff pass-

[8] The complete demand elasticity matrices corresponding to the 2005 and 2010 market shares given in table 1are $N_{2005} = \begin{bmatrix} -1.26 & 0.20 & 0.06 \\ 1.25 & -2.31 & 0.06 \\ 1.25 & 0.20 & -2.44 \end{bmatrix}$ and $N_{2010} = \begin{bmatrix} -2.02 & 0.87 & 0.15 \\ 0.48 & -1.63 & 0.15 \\ 0.48 & 0.87 & -2.35 \end{bmatrix}$. It may be noted that the η_{12} element in N_{2005} matrices is close to Muhammad *et al.*'s (2010, p. 437) estimate of this parameter (0.18), even though in their study Armington restrictions were not imposed.

through elasticity, i.e. equation (10)) will tend to be small, and (*ii*) the attenuation of the tariff pass-through elasticity of the domestic good caused by trade diversion is tiny (equation (17)). Moreover, the trade-diversion effect on demand for the protected good is even weak, leading a modest effect of trade diversion on both price and volume of the protected good.

The implications of these results would seem to be straightforward. First, in terms of net national welfare, the higher the extent of trade diversion, the smaller the tariff rate, *ceteris paribus*. A high tariff rate punishes domestic consumers more, since the tariff incidence is borne by the domestic consumers and the targeted foreign suppliers, given an upward sloping import supply curve. The side-effect of a large tariff rate is to exaggerate import diversion, which has a primarily negative effect on domestic producers' welfare. On the other hand, trade diversion has little consequence of reducing consumers' loss. This implies that a tariff rate beyond some limit may reduce the net national welfare. Second, the inverse relationship between domestic industry market share and the tariff pass-through elasticity of the protected good indicates that an industry with a large market share at duty inception can expect to be disappointed in that quasi-rents will be modest. To the extent the protected industry's market share continues to decline after AD implementation (as was the case for catfish), removal of the AD will induce larger economic consequences than its insertion. This asymmetry may help to explain the tendency for ADs to remain in place long after the initial petition was filed (Bown 2007).

Chapter II. Import Demand under Price and Exchange-Rate Uncertainties

2.1 Introduction

In a study of price uncertainty, Wolak and Kolstad (1991) point out that fluctuations in exchange rates might be one of the several inherent risk sources in actual import prices. In terms of exchange risk, while some studies argue for a negative effect on agricultural trade (e.g. Anderson and Garcia 1989; Cho, Sheldon, and McCorriston 2002; Wang and Barrett 2007; and Kandilov 2008), others advocate a positive relationship (see Langley, Giugale, Meyers, and Hallahan 2000; Awokuse and Yuan 2005). More recently, by employing panel data, Erdema, Nazlioglub, and Erdemc (2010) argue that the increased variability of exchange leads to more reduction in imports of agricultural goods than in exports in Turkey. Two implications can be generated from the previous research. First, the agricultural industry is more sensitive to risk factors in that agricultural goods are typically traded with flexible pricing strategies, and they are less storable than manufactured products (Wang and Barrett). Second, since the trade effect of the exchange uncertainty is associated with properties of the market, the impact of exchange risk should be evaluated in the context of disaggregate data (McKenzie 1999). But, disaggregate agricultural trade markets, in most cases, are perfectly competitive, indicating that the import price risk may reflect all information like the exchange-rate risk. If the exchange-rate risk completely passes through into the import price risk, the trade effect of the export price risk and the exchange-rate risk should be equal. One purpose of the present paper is to test the equivalence of the exchange risk effect and the export price risk effect.

Compared to exchange-rate risk, price risk has received less attention in the literature on agricultural trade. When studying the exchange-rate risk, Klaassen (2004) claims that price is predictable due to contract constraints. Although such a viewpoint is appropriate for industrial goods, it is not applicable for agricultural trade where product differentiation is weak and firms are more numerous (Carter and Gunning-Trant 2011). A search of the literature results in only two studies that test the impact of price risk on agricultural trade flows: Seo (2001) and Muhammad (2011), which are extended from Wolak and Kolstad (1991). Taking Chinese wheat imports market as an example, Seo investigates the relationship among the expected price, the systematic risk of price, and monopolistic power of the exporters. Muhammad develops a differential demand system by incorporating import price uncertainty and finds evidence that the UK carnation importing firms are more responsive to price risk than to the expected price.

In the research of either agricultural trade or non-agricultural trade, even less attention has been given to the differences between influences of exchange risk and price risk in trade flows. Ignoring the differences between price risk effect and exchange risk may also lead to conflicting empirical evidence on the trade effect of exchange risk. The differential effects of price and the exchange rate on trade may contribute to non-equivalence of effects of exchange risk and price risk. Two notable exceptions of which I am aware discuss and evaluate the relationship between price risk and exchange risk. Cushman (1983) states that uncertain price as well as exchange rate leads to percentage changes in real exchange rate, implying uncertainty in real exchange rate should be included in the trade models. Cushman's specification is more applicable for aggregate trade rather than disaggregate commodities that are of interest in the present paper. Kawai and Zilcha (1986) explicitly take exchange rate and commodity-price uncertainties into account when examining the optimum behavior of a risk-averse international trader; however, they do not develop an empirical analysis.

The inconclusive research results on exchange volatility may be also partly attributed to the flaws of its microeconomic foundation since the normally employed import demand equation in the previous literature is not derived from a utility function or is based on a too restrictive utility function. As De Grauwe (1988) illustrated, the positive trade effect of exchange risk is plausible given a general utility function underlying the import demand equation. Given a sufficient concave utility function facing importers, an increase in exchange risk might cause firms to raise import demand in order to avoid the worst possible outcome.

The main purpose of this research is to examine whether price and exchange volatilities have detectable effects on import demand by applying a demand system upon the utility theory. This question is of interest because the optimal-decision of importers should distinguish the impacts of price risk and exchange risk owing their different properties and different availabilities of hedging instruments. In addition, the linkage between risk impact and price effect needs to be revealed theoretically in order to analyze the empirical results. Few empirical studies combine export price risk and exchange risk, although previous research has extensively explored the trade effect of the exchange risk with ambiguous results. The US salmon import market is taken as an empirical case because salmon is one of the major seafoods imported by the US, and the export earning is a crucial instrument for one of the major suppliers, Chile, to restore balance of payments.

The main contribution of the present research is to distinguish exchange risk from price risk in a differential demand system and to examine impacts of those uncertainties on import demand. Implications of the research are an alternative method to handle different sources of uncertainty in the demand system and a complementary explanation based on the underlying utility function.

This study is structured as follows. I start with the theoretical framework followed by model specifications and data. After estimating the variances of price and exchange rate, I describe the regression results for the US salmon imports demand system. The paper concludes with brief remarks and implications.

2.2 Theoretical Model

The import demands for a particular good is derived from Armington's (1969) method, which assumes consumers employ a multistage budgeting procedure for allocating expenditure among competing sources. First, total expenditure is allocated over broad groups of goods based on a weakly separable utility function or a utility tree framework. Second, expenditure on this particular good is then allocated between the domestic and imported varieties. Finally, expenditure on imports of this particular good is divided among various source countries. A complete demand system built upon a utility function can reveal determinants of trade pattern adequately.[9] This demand system can further demonstrate the degree of market power that is contingent on the substitutability of goods from different sources (Gallaway, Blonigen and Flym 1999 p. 217). As Seo (2001) noted, the research of Wolak and Kolstad (1991) is limited to only homogenous goods, causing an omission of the supplier's influence on price risk owing product differentiation.

The Rotterdam model, which is derived from an implicit utility function, has been widely used to demonstrate the agricultural trade pattern (e.g. Seale, Sparks, and Buxton 1992; Washington and Kilmer 2002; Muhammad 2009, 2011). Among them, Muhammad (2011) is the first study applying the Rotterdam model to examine the trade effect of risk variables. Different from Muhammad (2011), in the present research, I justify theoretically that uncertainty factors

[9] The conditional demand systems can be derived from either the consumer demand theory or production theory. As Washington and Kilmer (2002) point out, there are no empirical differences in estimates of conditional elasticities between consumers demand theory and production theory.

take effect through marginal utilities, and the effect depends on price effects and substitutability between goods.

Taking volatility components into account, the extended general Rotterdam model can be derived from the following utility maximization problem:

(1) $$\underset{(q)}{Max}\ u = u(q, v)$$

s.t. $p'q = y$

where u is the utility to be maximized, q is a vector of imports from different source countries, p is the corresponding import price vector, v is a vector of risk variables, and y is the conditional expenditure on the imported goods of interest.

The first order conditions of the utility maximization problem are:

(2) $$\sum_{j=1}^{n} p_j q_j = y$$

(3) $$-\lambda p_i + u_i = 0 \qquad\qquad i = 1, 2, \ldots, n.$$

where $u_i = \frac{\partial u}{\partial q_i}$ is marginal utility, λ is the Langrage multiplier.

For the maximization problem with the conditional budget constraint, a negative definite bordered Hessian matrix is the necessary and sufficient condition, implying the matrix

(4) $$U = \begin{bmatrix} 0 & u_j \\ u_i & u_{ij} \end{bmatrix}$$

is negative definite. Here, $u_{ij} = \frac{\partial u_i}{\partial q_j}$.

From the equations (2) and (3), the effects of price (p_j) on the ith product (q_i) and expenditure (y) are therefore given by:

(5) $$q_{ij} = \frac{\lambda(-q_j U_i + U_{ij})}{|U|}$$

(6) $$q_{iy} = \frac{\lambda U_i}{|U|}$$

where U_{ij} and U_i are cofactors of the matrix U, $|U|$ is determinant of this matrix. Here q_{ij} $(= \frac{\partial q_i}{\partial p_j})$ is parallel to the Slutsky equation in Phlips (1983, p. 49). I further denote $q_{ij}^* = \frac{\lambda U_{ij}}{|U|}$, which is a measure of Hicksian complementary or substituting effect (Tintner 1952).

In the spirit of Tintner (1952), which is recently applied in Brown and Lee (2010), the impact of a volatility variable (v_j) is revealed in the demand system by differentiating the first order equations with respect to one risk variable to yield:

(7) $$\begin{bmatrix} 0 & u_j \\ u_i & u_{ij} \end{bmatrix} \cdot \begin{bmatrix} -{}^{\lambda_{v_j}}\!/_{\lambda} \\ q_{i,v_j} \end{bmatrix} = \begin{bmatrix} 0 \\ -u_{i,v_j} \end{bmatrix}$$

where λ_{v_j} denotes $\frac{\partial \lambda}{\partial v_j}$, $u_{i,v_j} = \frac{\partial u_i}{\partial v_j}$ represents the effect of volatility on marginal utility of the *i*th product, and q_{i,v_j} is the effect of the volatility *j* on the *i*th product.

Next, from (7), I solve q_{i,v_j}:

(8) $$q_{i,v_j} = -\sum_{k=1}^{m} \frac{U_{ik}}{|U|} u_{k,v_j}$$

Considering the equation $\eta_{ik}^* = \frac{\lambda U_{ik}}{|U|}$, equation (8) can be restated as

(9) $$q_{i,v_j} = -\sum_{k=1}^{m} \frac{p_k}{u_k} \eta_{ik}^* u_{k,v_j}$$

After obtaining the results of effects of the exogenous variables (p, y, and v) on the demand for the *i*th good, I derive the extended Rotterdam model by using a differential approach (Theil 1977, 1980). First, the import demand equations are solutions to (2) and (3).

(10) $$q_i = q_i(y, p, v)$$

By differentiating both sides of (10) and relating the parameters to (5), (6) and (9), I obtain the extended Rotterdam model in the form:

(11) $$w_i d\ ln\ q_i = \theta_i d\ ln\ Q + \sum_{j=1}^{n} \pi_{ij} d\ ln\ p_j - \sum_{k=1}^{m} \beta_{ik} d\ ln\ v_k$$

where $Q = \frac{y}{\sum_{j=1}^{n} w_j \ln p_j}$ is real expenditure, $w_i = \frac{p_i q_i}{y}$, $\theta_i = w_i A_i = p_i \frac{\lambda u_i}{|U|}$, $\pi_{ij} = w_i \eta_i^* = \frac{p_i p_j}{y} \frac{\lambda u_{ij}}{|U|}$, and $\beta_{ik} = \sum_{j=1}^{n} \pi_{ij} \eta_{j,v_k}$. Here A_i is expenditure elasticity, η_{ij}^* is Hicksian price elasticity, and η_{j,v_k} is elasticity of marginal utility of the jth product with respect to v_k. The estimated coefficients can be explicitly converted into the corresponding elasticities.

Indicated by $\beta_{ik} = \sum_{j=1}^{n} \pi_{ij} \eta_{j,v_k}$, the net effect of the volatility depends on its impact on the marginal utility of *each product* in the group, which is further weighted by the corresponding price effect. This is consistent with what De Grauwe (1988) points out: the exchange volatility affects import demand through the consumer's marginal utility for the good, and the direction of this effect depends on the curvature of the underlying utility function. For example, in the likely case where the own-risk (v_i) only affects the ith good ($\beta_{ii} = \pi_{ii}\eta_{i,v_i}$) or the own-risk effect dominates the sign of the reduced effect of the volatility, the direction of the volatility's impact depends solely on the effect of the ith volatility on the marginal utility of the ith good (η_{i,v_i}), as the sign of π_{ii} is *a priori* negative. For example, the trade effect of volatility is positive when $\eta_{i,v_i} > 0$ due to sufficiently risk-averse behavior of importers, i.e. a more concave utility function. The opposite is true if $\eta_{i,v_i} < 0$. Hence, whether volatility exerts a positive or negative effect on trade volume depends on the direction of its effect on marginal utility. This caveat needs to be borne in mind when interpreting the empirical results.

In (11), coefficients of the risk variables are in reduced form. In order to further demonstrate how volatility factors affect the market, with simple manipulation, I restate equation (11) as follows:

(12) $w_i d \ln q_i = \theta_i d \ln Q + \sum_{j=1}^{n} \pi_{ij}(d \ln p_j - \sum_{k=1}^{m} \eta_{j,v_k} d \ln v_k)$

where the parameters of volatility are expressed in structural forms. This expression reflects that change in the *j*th "effective" price is the actual price change minus the summation of changes in marginal utility of the relevant *j*th product as a result of changes in all volatility variables in the demand system. If changes in *k*th volatility decrease (increase) marginal utility of the *j*th good, demand for the *i*th good would be more (less) sensitive to changes in the *k*th price given the ith good and the *j*th good are substitutable (complementary) to each other.[10] In terms of uncertainty in international trade, Balg and Metcalf (2010) and Bergin (2004) posit a risk-averse firm would attach a risk premium as an extra markup to cover the costs of exchange-rate fluctuations; Wolak and Kolstad (1991) postulate that input-price risk premium is the percentage above the current expected market price a firm would pay for riskless input supply.

When implementing (12) to the empirical research of agricultural trade, I decompose the import price risk into export price risk and exchange risk to obtain the following theoretical model:

(13) $w_i d\ ln\ q_i = \theta_i d\ ln\ Q + \sum_{j=1}^{n} \pi_{ij}(d\ ln\ p_j - \sum_{k=1}^{m} \gamma_{jk}\ d\ ln\ v(p_k^*)$ - $\sum_{l=1}^{o} \lambda_{jl} dln\ v(e_l))$

where $v(p_k^*)$ stands for risk of the export price measured in exporting country's currency, and $v(e_l)$ represents the exchange risk.

The underlying utility function with the budget constrain represented by equation (1) indicates the general restrictions on the demand system, namely:[11]

(14) $\sum_{j=1}^{n} \pi_{ij} = 0$ (homogeneity)

(15) $\pi_{ij} = \pi_{ji}$ (symmetry)

(16) $\sum_{i=1}^{n} \theta_i = 1, \sum_{i=1}^{n} \pi_{ij} = 0$ (adding-up)

[10] The interpretation of the risk factor exerting its role through changes in "effective prices" is identical to that in the advertising-augmented trade model (e.g. Duffy 1995)
[11] Coefficients of risk factors satisfy the adding-up restrictions through the price effects. See Brown and Lee (2002) for more detail.

Besides the general constraints, other specific restrictions can be placed on the effects of the risk variables upon the econometric tests against the unrestricted model A represented by equation (13). This results in a reduction of the parameter spaces and efficiency of regression results. Duffy (1987) assumes no cross effect of preference variables, implying the following specific restrictions:

(R1) $\gamma_{jk} = 0$ for $j \neq k$ (zero cross-price risk effect)

(R1') $\lambda_{jl} = 0$ for $j \neq l$ (zero cross-exchange risk effect)

If Duffy's restriction (R1) cannot be rejected, a stronger restriction (Theil 1980) that own-risk effects are identical can be further tested (for elaboration, I suppose there are 5 price risk variables and 4 exchange volatility variables in the demand system):

(R2) $\gamma_{11} = \gamma_{22} = \gamma_{33} = \gamma_{44} = \gamma_{55} = \bar{\gamma}$ (constant own-risk effect)

(R2') $\lambda_{11} = \lambda_{22} = \lambda_{33} = \lambda_{44} = \bar{\lambda}$ (constant own-risk effect)

Given R1 and R2 cannot be rejected, one main concern in the present paper is the equivalency of own-price risk and exchange own-risk effect:

(R3) $\bar{\gamma} = \bar{\lambda}$ (equivalency of price risk and exchange risk effects)

Lastly, I test the hypothesis of risk neutrality

(R4) $\bar{\gamma} = 0$ (price-risk neutrality)

(R4') $\bar{\lambda} = 0$ (exchange-risk neutrality)

If R4 is rejected, the conventional Rotterdam model exclusive of risk factors is preferred.

Different from the general restrictions that, as common treatment in the literature, can be imposed in the regression to be consistent with the demand theory, the specific restrictions (R1-R4) are imposed only if they are compatible with the data.

The above specific hypotheses are summarized as follows

Models for testing

Model symbol	Restrictions
Model A	Equation (13) with homogeneity and symmetry constraints imposed
Model B	Zero cross-risk effect (R1 and/or R1’)
Model C	A constant own-risk effect (R2 and/or R2’)
Model D	Equivalency of price risk and exchange risk effects (R3)
Model E	Risk neutrality, i.e. conventional demand system (R4)

2.3 Model Specifications and Data

The forgone theoretical model is proposed to highlight the effects of export price risk and exchange risk on the US import demand for salmon, which can be differentiated by sources.

Since the introduction of salmon aquaculture in the early 1980s, the international salmon trade is growing with Chile, Canada, Norway, and the UK being the main suppliers. Among importers, the US has traditionally been the world’s major market for salmon. Currently, Atlantic salmon is the second most imported seafood product in the US only after shrimp. Chile and Canada were the leading countries exporting farmed Atlantic salmon to the US throughout the past twenty years. In 1995, America imported salmon at the value of $277 million, of which about 42% came from Chile and 49% from Canada (see table 1). The US salmon import surged almost 4.6 times from 1995 to 2008; however, the market was still highly concentrated with more than 85% of total imports from Chile and Canada. During the sample period, the prices of salmon from different sources had different levels of volatility despite the stable relative price ratios (see figure 1). Subsequently, the trade effect of the importers’ attitudes to price risk deserves more attention. Taking Chile as an example, the great fluctuation of price and exchange rate did not reduce its dominant position in this market (see table 1), implying the importers might not be sensitive to the volatility of Chile's salmon price.

In this study, US salmon import data on value (CIF in US dollar) and quantity (kilogram) are from the US International Trade Commission (USITC) where import salmon is represented by 52 HTSUS-10 codes according to different species and forms of salmon. Among them Atlantic salmon (fresh, frozen, fillet fresh, and fillet frozen) represented by 11 HTSUS-10 codes consistently accounted for about 85% of total imports during the sample period. The data period is from January 1995 to December 2008, totally 156 observation.[12] Data on exchange-rates are obtained from the US Department of Agriculture – Economic Research Service (ERS).

The demand system for US import salmon contains five equations distinguished by import sources, namely Chile, Canada, Norway, the UK, and the rest of the world (ROW) that is an aggregation of farmed Atlantic salmon imported from countries not specified. For a particular supplier, the monthly import price is obtained by dividing the total import (CIF) value in the US dollar by quantity (kilogram). For ROW, the monthly aggregate value and quantity are used to calculate the import price. Dividing the import price (in the US dollar) by the corresponding monthly average exchange rate (foreign currency per US dollar) yields the export price in foreign currency. After creating the import price and export price, the variances of prices (and exchange rates) can be estimated. Since the US dollar is used as the representative currency for ROW, only the export price risk is measured for ROW. Subsequently, upon the theoretical model represented by equation (13), the empirical model estimated is

(17) $\overline{w_{\iota,t}}\, \Delta ln\, q_{i,t} = \alpha_i + \theta_i \Delta\, ln\, Q_t + \sum_{j=1}^{5} \pi_{ij} [\Delta\, ln\, p_{j,t} - \sum_{k=1}^{5} \gamma_{jk}\, \Delta\, ln\, v(p_{k,t}^{*})$

[12] The sample period is ended in 2008 in order to avoid structural changes in the US salmon import market induced by the 2008 economic recession and the outbreaks of infectious salmon anemia (ISA) that has been released since later 2007 (see Asche, Hansen, and Tveteras 2009).

$- \sum_{l=1}^{4} \lambda_{jl} \Delta \ ln \ v(e_{l,t})] + \varepsilon_{i,t.}$

where i denotes suppliers (Chile = 1, Canada = 2, Norway = 3, the UK = 4, and ROW = 5), t stands for the time subscript (monthly), w_i is budget share, q_i represents import volume, p_i is import price (in US dollar), $v(p_k^*)$ stands for variance of export price (p_k^*, in foreign currency), $v(e_l)$ represents variance of the real exchange-rate $v(e_l)$, and ε_i is the i.i.d. error term. The intercept a_i is included in each demand equation to account for autonomous shifts in demand due to taste change or other trend-like phenomena. Following Theil (1980), changes in real expenditure are replaced with a Divisia volume index ($\Delta \ ln \ Q_t = \sum_{j=1}^{5} \overline{w_{\iota,t}} \Delta \ ln \ q_{j,t}$), and finite logarithmic changes are employed to replace infinitesimal changes in the theoretical model. However, the variables are 12-month differenced in order to account for seasonality in demand (Lee 1988) and also to save the degree of freedom. Hence, $\Delta lnx_t = lnx_t - lnx_{t-12} \approx dlnx_t$. Similarly, $\overline{w_{\iota,t}}$ is the arithmetic mean of the expenditure share of the ith good in t-12 and t.

2.4 Measurement of Uncertainty

The generalized autoregressive conditional heteroskedasticity (GARCH) model is employed to measure the price and exchange-rate volatilities. The GARCH method can fully reveal the characteristics of time-series data (Kandilov 2008) and can explicitly test whether the movement in the conditional variance of price or exchange rate over time is statistically significant (Pattichis 2003). Those advantages lend the GARCH method to wide applications in the agricultural trade literature (for example, Wang and Barrett 2007; Kandilov 2008; and Erdema, Nazlioglub, and Erdemc 2010).

For a typical AR(1)-ARCH(1,1) model in which the residuals are from a AR(1) process, the volatility is obtained by jointly estimating the following equations:[13]

(18) AR(1): $E_{i,t} = \pi_0 + \pi_1 E_{i,t-1} + \Sigma_{i,t}$

(19) GARCH(1, 1) $V(E_i)_t = \delta_0 + \delta_1 \Sigma^2_{i,t-1} + \delta_2 V(E_i)_{t-1} + \Theta_{i,t}$

where, E_i is the differential logarithm price or bilateral exchange-rate variable, $V(E_i)$ represents the conditional volatility, and $\Sigma_{i,t}$ and $\Theta_{i,t}$ are error terms of AR(1) and GARCH (1,1) processes, respectively. For the GARCH(1,1) process to be well-defined, restrictions on the estimated coefficients include $\delta_0 > 0$, $\delta_1 + \delta_2 < 1$, and $|\delta_1| < 1$.

Price and exchange variance estimates by country are depicted in figure 2-4. For each major supplier, the import price is more volatile than the corresponding export price. The exchange-rate variances seem to track each other reasonably well. On the other hand, the exchange rates in general fluctuate less than either export prices or import prices. This renders equivalence of the export price risk and the exchange-rate risk effects suspect. For Chile and Canada, which successfully captured the lion's share of the US salmon imports, their prices and exchange rates were no less volatile than those of other suppliers.

2.5 Regression Results

The demand system of the US salmon import market contains five equations distinguished by exporting sources: Chile, Canada, Norway, the UK, and ROW. In estimation, one equation (ROW) is dropped from the system to avoid singularity. The relevant coefficients can be recovered on the basis of demand constraints. The preliminary estimation using the Seemingly Unrelated Regression (SUR) indicated the evidence of first-order autocorrelation in some

[13] As Pattichis (2003) points out, the order of the AR process has little impact on the GARCH models.

equations. Accordingly, the General Method of Moments (GMM) approach was employed to produce the Newey-West estimator in order to correct autocorrelation as well as heteroskedasticity.

Model selection and Hypothesis tests

For purposes of comparison, the Rotterdam model inclusive of the import price risk variables is first estimated, and the results are displayed in table 2.[14] The results in general reflect that the differential-based demand system is compatible with the data at a satisfactory level in that all 5 estimated expenditure coefficients and 4 own-price coefficients are significant and have the correct signs in agreement with the demand theory. In the case of Chile, the major supplier, the own-price effect is negative but not significant. Most cross-price effects are significant with a positive or negative sign. The negative cross-price elasticities between Chilean and Canadian salmon indicate those two varieties are complementary to each other. This contradicts the fact that there were opposite trends of market shares for Chile and Canada during the sample period. For variance estimates, in the cases of Chile, Canada, and Norway, coefficients of the export price volatilities are significantly different from zero, ranging from -0.6 to -1.3. However, the insignificant own-price effect in Chile's equation and the unexpected substitutability between Chilean salmon and Canadian salmon render the estimated import price variance effects suspect. This leads us to the estimation of the extended Rotterdam model, where the import price risk is decomposed into the export price risk and the exchange risk.

Before turning to the empirical results of the extended Rotterdam model, specific restrictions of the demand system are justified by a comparison of models with different

[14] The general demand restrictions and Duffy's restriction (zero cross-risk effects) are imposed when estimating the model.

economic hypotheses.[15] According to the likelihood ratio (LR) test results (table 3), I fail to reject Model C against Model B where the Duffy's restriction is imposed, indicating the support of Theil's hypothesis, i.e. a constant exchange risk effect. The results may relate to the common properties of data generating processes for exchange rates, as demonstrated in figure 4. Rejection of Model C' against Model B (p-value = 0.03) indicates different responses of marginal utility to changes in export price risk due to differential price signals. The information related to price should be more available for imports from major suppliers. This is further evidenced by the testing results of the hypothesis that effects of price variances for major suppliers (i.e. Chile, Canada, and Norway) are identical (p-value = 0.23). Consequently, the hypothesis of a constant export price variance effect is treated as a maintained hypothesis in the present paper. Next, I test the equivalence of import price risk and exchange risk effects by testing Model D against Model C where restrictions of a constant export price risk effect and a constant exchange risk effect are imposed. The LR test results show Model D is rejected with a p-value = 0.023. Subsequently, the conventional Rotterdam model exclusive of risk factors (Model E) is tested against Model C. Model E is firmly rejected ($p < 0.0001$), indicating: (1) importers are not risk neutral, reflecting risk factors should explain the observed salmon trade pattern in addition to traditional variables like relative prices, and (2) the empirical results from the conventional Rotterdam model may be misleading due to misspecification. According to the above test results, the remaining discussion will rely on estimates of Model D and the results are reported in table 4, where regression results of the conventional Rotterdam model (Model E) are also listed to highlight the differences.

Estimates of Model D are overall satisfactory in that all marginal share estimates (θ_i) are significantly positive and all own-price coefficients (π_{ii}) are negative and significant. Chile and

[15] I start from Model B due to the parameterizing problem. Following Phlips (1983, p55), all models have homogeneity and symmetry imposed though both properties are rejected in preliminary tests.

Canada's regression results have the best explanatory powers with R^2 equaling 0.48 and 0.55, respectively. With respect to expenditure and price effects, the estimated coefficients of Model D are substantially different from the counterparts of the conventional Rotterdam model (Model E).

In model D, both the constant export price risk effect and the constant exchange-rate risk effect are significant and are negative in sign, implying the US salmon importers are in general risk-averse. The magnitude of the exchange risk coefficient is 2 times the magnitude of the export price risk coefficient (-0.13 vs. -0.07). This indicates that the marginal utility of the imported salmon is more responsive to the exchange risk than to export price risk. Compared to the export price risk, the exchange risk is far beyond the control of the agricultural traders. The estimated coefficients of variances of the export price risk and the exchange risk are substantially smaller than the estimated import price risk effects in the alternative Rotterdam model. Because in the Rotterdam model the statistical significance of parameters have less economic meaning than elasticities, the remaining discussion will focus on conditional expenditure elasticities, Hicksian price elasticities, and risk elasticities, which are estimated at the means of budget shares (table 5).

Price and Expenditure Elasticities

For Chile, Canada, and Norway, the expenditure elasticities are close to one, in the range between 0.98 and 1.16, suggesting consumer preferences for salmon from these sources are homothetic. When consumer preferences are homothetic, the market share only responds to changes in the relative price. This further implies that price and price risk play a crucial role in this market. On the other hand, imports from runners-up in this market (i.e. the UK and ROW) are less sensitive to changes in the total import expenditure. For example, a one-percentage growth in the conditional expenditure would improve imports from UK and ROW by 0.4% and

0.2%, respectively. The uneven distribution of benefits from the rising expenditure can explain the stable market share of Chile in 1995-2008 (table 1).

The estimated own-price elasticities in the range between -0.3 in Chile and -2.4 in Norway indicate that importers respond differently to changes in prices of salmon from different sources. The import demand is less sensitive to changes in major suppliers' prices and is more sensitive to changes in other suppliers' prices. Considering aggregate data on frozen and fresh salmon were employed in the regression, those findings are somewhat in accordance with the previous research. Xie, Kinnucan, and Myrland (2009) estimate the world demand for fresh salmon is slightly elastic at 1.03 and for frozen salmon at 0.37. The greatest own-price elasticity for Norway (-2.4) is not unexpected, because fresh salmon accounted for a great proportion of imports from Norway during the data period.

In most cases, the cross-price elasticities are significant and positive, suggesting competition between any pairs of salmon in the group of interest. The only exception is salmon from Norway and the UK, where the negative cross-price elasticities indicate complementary relations between these two goods. Import demand for Chile and Canada are less sensitive to prices of salmon from other suppliers. For example, a one-percentage increase in the price of Norway's salmon would raise imports from Chile by 0.12% and raise imports from Canada by 0.29%. On the contrary, import demand for salmon from the runners-up is more sensitive to changes in prices of salmon from the major suppliers. Taking Norway as an example, a one-percentage increase in prices of Chile and Canada's salmon would improve imports from Norway by 1.4% and 2.2%, respectively. These findings are consistent with the properties of the market where Chile and Canada jointly had the lion's share of the market throughout the sample period.

Volatility Elasticities

The estimates of uncertainty variables are central to this research. Since I cannot reject the Theil's restriction that either the price risk or the exchange risk takes a role in the marginal utility via "adjusted" price with a negative constant, the strength of the response in demand for the *i*th salmon with respect to changes in the risk factor of the *j*th salmon solely depends upon substitutability between the these two varieties. For example, in Canada's equation, an increase in variance of Chile's salmon price would expand Chile's "adjusted" price, which consequently effects the demand for Canada's salmon to some extend, depending on the magnitude of import demand elasticity of Canada's salmon with respect to Chile's price.

Uncertainty from the own-currency realignments exerts a significantly negative effect on imports from Chile (-0.04), Canada (-0.08), Norway (-0.32), and the UK (-0.11). These results are compatible with the findings in Wang and Barrett (2007) and Kandilov (2008). Thus, the estimate results reject De Grauwe's hypothesis (1988) that an increase in the exchange risk would cause firms to import more in order to avoid the worst possible outcome. Compared to exchange risks, export price risks have less substantial impacts on the trade flows. The variances of the own-price have negative effects in all cases, with the derived elasticities ranging from -0.02 in Chile and -0.18 in Norway. With regard to the export price risk elasticities, the different response of the import demand is likely to be explained by market shares, transportation costs, and availability of information. This is evidenced by a greater tolerance for the price volatilities of salmon from Chile and Canada, which are major suppliers in this market, and at the same time, have a shorter distance to the US than any other suppliers, including Norway and the UK.

Different from own-risk elasticities, which have consistently negative signs, the cross-risk effects are ambiguous, depending on substitutability between products. Demand for salmon from Chile is less significantly influenced by the price volatilities of Canada's salmon and Norway's salmon, with a modest magnitude of 0.005 and 0.009, respectively. Demand for Canada's salmon is more sensitive to Chile's price volatility than the response of Chile's salmon demand to Canada's price volatility (0.008 vs. 0.005). The same conclusion can be made on the comparison of exchange-rate volatility elasticities. The different responses to risk factors may be correlated with changes in the market where Chile gained an 11% increase of shares from 1995 to 2008, and Canada lost 13%.

Two arguments further shed light on trade effects of the risk variables. First, demand elasticities with respect to risk factors are generally small in absolute values, since the impacts of risk variables are weighted by price effects, which are, in most cases, are not strong. Second, the substantial difference between export price risk and exchange-rate risk may explain why, in some cases, the Rotterdam model inclusive of the import price risk failed to capture significant effects of own-price variance.

2.6 Conclusions

Analogous to the treatment of preference variables in the literature (Brown and Lee 2002, 2010), I augmented risk factors into the Rotterdam model in which risk factors affect marginal utilities via "adjusted prices". This methodology is appealing since it is commonly recognized that risk-averse firms are supposed to add a proportional markup on the realized or expected prices. When the extended Rotterdam model inclusive of risk factors is employed to evaluate the trade pattern, one hypothesis deserved more attention: the equivalence of export price risk and exchange risk effects. In the literature, few endeavors have been made to examine the combined effect of export-price risk and exchange risk, especially regarding agricultural trade where trade participators are presumed to be more risk-averse.

The empirical evidence given above appears to sustain the conjecture that export price risk and exchange risk take significant effect on the US import demand for farmed Atlantic salmon, although they differ from each other. The difference between effects of export price risk and exchange risk may be associated with abilities of importers to control the uncertainty in the market. Despite being significant in most cases, the elasticities of price risk and exchange risk are tiny. This can be explained by price inelastic demand curves and weak substitutability between products from different source countries since, theoretically, risk factors take effect on trade flows by changing marginal utility, which is further weighted by price effects. Those findings are consistent with the features of the US farmed salmon import market, where the two largest suppliers (Chile and Canada) accounted for about 85% of the total salmon throughout the sample period, although their prices and exchange-rates had a great degree of fluctuation. In general, relative price advantages, availability of substitutes, and consumer preferences are likely to be the main factors resulting in the observed trade pattern.

Chapter III. A Risk-Augmented Cointegrating Import Demand System

3.1 Introduction

The main purpose of this paper is to examine the role of price risk in the allocation of import expenditures across exporting sources by taking the US codfish market as an empirical application. The agricultural industry is more sensitive to price risk in that agricultural goods are less storable than manufactured products and are traded with flexible prices (Wang and Barrett 2007), and agricultural product differentiation is weak and firms are more numerous (Carter and Gunning-Trant 2010). Trade data from the US International Trade Committee (USITC) reveals that China replaced Canada as the major supplier of cod to the US after 2004. In addition to the traditional variables like relative prices, I want to evaluate the extent to which the import price risk can explain the observed trade pattern.

In the research of agricultural trade, little attention has been paid to the influence of price uncertainty in trade flows. Seo (2001) and Muhammad (2011) are two exceptions. Taking Chinese wheat imports market as an example, Seo investigates the relationship between the expected price, the systematic risk of price, and monopolistic power of the exporters. Muhammad develops a differential demand system inclusive of import price uncertainty and finds the evidence that the UK carnation importing firms are, for the most part, risk-averse. In a study of price uncertainty, Wolak and Kolstad (1991) point out that fluctuations in exchange rates might be one of the inherent risk sources in actual import prices. The significant relationship between exchange rate risk and import demand for agricultural commodities is verified in Langley et al. (2000); Cho, Sheldon, and McCorriston (2002); and Kandilov (2008). Considering the numerous trade firms, the US codfish market can be presumed to be reasonably competitive, indicating that the import price expressed in US dollars should contain all the relevant information such as exchange volatility.

In the above trade literature, few studies are based on a complete demand system with the exception of Muhammad (2011). This lack of attention to demand interrelationships may cause biased empirical results. As Alessie and Kapteyn (1991 p. 404) state, "One obvious omitted factor in micro-studies is the interdependence of preferences." In a well-integrated international market like codfish, the trade impact of the price risk of a product from a particular country can be offset by the price risk of other similar goods through a substitute or complementary effects. Moreover, the system-wide approach to demand system facilitates the joint testing of theoretical restrictions due to the underlying utility function (Duffy 1987), and evaluating the preference variable like the price risk that is the main concern in the present paper.

Different from Muhammad (2011), where a differential approach in line with the Rotterdam demand model is applied, in the present paper I build risk factors into the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer (1980). The extended AIDS model inclusive of risk factors can distinguish the "competitive effect" and "baseline effect" of price risks, which is important to demonstrate the risk preference of importers, as I will see later. Another difference of the present paper compared with Muhammad (2011) is that I account for the stationarity of data generating process when estimating the empirical model.

On the basis of a great deal of compelling reasons and evidence from previous research, the price terms may be endogenous in the typical demand system for agricultural goods. Furthermore, the price volatility can also be deemed to be jointly determined with prices and expenditure shares, considering contracts between importers and leading suppliers can strongly stabilize the price. Therefore, in the present paper, the empirical model is estimated by applying a cointegrating-based Vector Error Correction Model (VECM) approach, in which the "normalization" of the demand equation provides a useful framework to allow for endogeneity of either price or volatility. Duffy (2003) investigates the advantages of cointegrating long-run estimates compared to static equilibrium estimates from other regression methods like the Seemingly Unrelated Regression (SUR). As Reziti and Ozanne (1999) state, the cointegration technique can further improve the estimation of the theoretical model with regard to the adjusting mechanism of the market. This view is cited and echoed by Granger (1999, p. 16):

> "The classical approach to constructing a model starts with a sound, internally consistent, economic theory which provides a tight specification for the empirical model. This model is then estimated and interpreted. Unfortunately, this strategy towards modeling has not always proved to be a successful one. Models produced in this way often do not fit the data in various important directions. As one pair of applied economists put it, 'a recurring problem in empirical studies of consumer and producer behavior is that the regularity properties implied by microeconomic theory have more often than not been rejected' (Rezili and Ozanne (1997)), who then go on to say 'such rejection means that empirical work loses a good deal of its theoretical credibility.' They point out that a major problem is 'the failure of static equilibrium theory to account for dynamic aspects of static equilibrium theory to account for dynamic aspects of consumer and producer behavior' and show how the introduction of dynamics into an equilibrium model, by use of a stricture known as an error-correcting model, leads to clear improvements."

In what follows, the US codfish import market is briefly overviewed. Next, I discuss the theoretical framework. Afterwards, the empirical models are established upon the theory, followed by the measurement of price volatility and estimation methods. The latter includes estimation results and analysis. The final section consists of summary and implications.

3.2 Background

The US codfish import market is selected to implement the empirical study. For the last 20 years, this market has been relatively concentrated with Canada, China, and Iceland as the major suppliers; however, expenditure shares allocated by US importers to the main suppliers varied dramatically over recent decades. Canada and Iceland have historically been the largest exporters to the US, but since 2004 China has dominated this market. At the same time, the total (CIF) value of codfish import reduced by 45 percent in the period from 1989 to 2010. The huge variation in this market makes the codfish an insightful case to the role of price risk in addition to traditional determinants (expenditure and prices) in the observed trade pattern. Further, the suppliers include a developing country (China) and a developed country (Canada) that have different exchange rate regimes and different transportation costs, leading to different movements of prices. Kandilov (2008) claims that trade effects of exchange volatility are remarkably larger on agricultural exports from the developing countries than from the developed countries.

Codfish, or cod, is a species of deep-sea fish common to the North Pacific Ocean and North Atlantic Ocean. According to USITC, from 1989 to 2010, the US has imported codfish from 80 countries, among which the leading countries include Canada, China, and Iceland. On average, the annual imports (CIF) of codfish from those countries account for about 70% of total codfish imports. In the USITC database, there are 25 HTS-10 codes for codfish responding to different species and forms. Among them fillet frozen, fillet fresh, fresh, frozen, and dried codfish represented by 8 HTS-10 codes consistently accounts for about 90 percent of total imports. For the top suppliers, Canada has dominated dried codfish during the sample period. Canada and Iceland shared the fresh cod market, and China explored and controlled the fillet frozen codfish market where other top suppliers have not substantially entered. In this paper, the major species and forms of codfish are aggregated to construct the demand system upon a two-stage budgeting process.

Over the last twenty years, US codfish import volumes have experienced dramatic changes. In total, the US imported 474 million US dollars value of codfish in 1989; however, the total value sharply declined to 262 million US dollars in 2010. This sharp decline happened in the early years of the period, and the variation of total imports has become relatively weak since 1994. The low fluctuation of total imports does not suggest a stable market share for individual suppliers. As shown in table 1, imports of codfish from China grew faster after 1999, resulting in China's dominant role in this market since 2004. Between 1989-1999, on average, the annual CIF value of Canadian codfish accounted for 41% of the total codfish imports and Chinese codfish only 2.6%. However, in the period of 2000-2010, the positions for these two countries were reversed in the market (17% vs. 44%).

The observed trade pattern can be explained by changes in importers' budget and relative prices (table 1, figure A1). The four price series have a decidedly upward movement and seem to track each other reasonably well. However, the price of China's cod increased a lesser degree with a comparison of prices of Canada and Iceland's cod. Figure A1 further implies an interesting fact that there is a negative relationship between expenditure share and variation of price. Before 2004 when Canada dominated the codfish market, the movement of its price was weak; however, Canada's price fluctuated strongly when its expenditure share was decreasing after 2004. For China, the degree of price variations has become significantly smaller since China was changing from a runner-up to a leader in this market after 2004. Two implications can be generated from this observation. First, the variance of price may be one of the determinants to explain the trade pattern. Second, expenditure share and movement of price might jointly determine each other, indicating the plausibility of applying the cointegrating technique in the empirical analysis.

3.3 The Theoretical Model

Wolak and Kolstad (1991) postulate that an input-price risk premium (or cost of risk) is the percentage above the current expected market price a firm would pay for riskless input supply. In terms of elements of price risk like exchange fluctuations, Balg and Metcalf (2010) and Bergin (2004) posit a risk-averse firm would attach a risk premium as an extra markup to cover the costs of uncertainty. Therefore, a change in price risk would influence the "effective" price of the good ("competitive effect"). On the other hand, risk-averse importers are perceived to define a baseline plan due to maintenance of cooperation and diversification. The baseline expenditure pattern can be modified by the information about price risk ("baseline effect"), though it might be independent of price. For a risk-averse firm, the competitive effect of risk is expected to be negative, but the baseline effect may be positive or negative depending on the firm's attitudes to risk. If the firm is sufficiently risk-averse, the positive baseline effect may be greater than the negative competitive effect and dominates the sign of combined effect of price risk. As De Grauwe (1988) postulates, it is likely that an increase in price risk might cause firms to import more to avoid the worst possible outcome. Starting from a traditional AIDS model, I demonstrate how to incorporate the competitive effect and the baseline effect of price risk into this demand system.

Following Deaton and Muellbauer (1980), the importer's expenditure function is first defined in the form:

(1) $\log c(u, p) = a(p) + u\, b(p)$

where $c(\cdot)$ is a cost (expenditure) function, u is the utility function, p is a vector of price, and

(2) $a(p) = \log P = a_0 + \sum_k a_k \log p_k + 0.5 \sum_k \sum_l \gamma^*_{kl} \log p_k \log p_l$

(3) $b(p) = \beta_0 \prod_k p_k^{\beta_k}$.

where a_0, a_k, β_0, β_k, and γ_{kl}^* are parameters.

Applying Shephard's lemma and defining $\gamma_{kl} = 0.5(\gamma_{kl}^* + \gamma_{lk}^*)$, the general AIDS demand system with the expenditure share (w_i) as the dependent variable can be derived from specification (1) - (3),

(4) $w_i = a_i + \beta_i \log (x/P) + \sum_j \gamma_{ij} \log p_j$

where a_i can be explained as the baseline portion of imports from the ith country, and x is the total expenditure.

Risk-averse importers should take the risk components (v) into account when allocating expenditure across suppliers to maximize utility or minimize expenditure. First, the competitive effect of risk factors implies that the utility function can be modified by incorporating multiplicative scaling factors.

(5) $u = u(q^*)$

where $q_i^* = q_i m_i$ and m_i is a scaling factor. As common in the literature on preference variables in consumption literature (for example, Duffy 1995), m_i is endogenized by assuming that it varies with the level of price volatility of the ith product in a constant elasticity formulation, i.e. $m_i = v_i^{\delta_i}$ where δ_i is negative and is greater than unity in absolute value if importers are risk-averse. Since the budget constraint for importers is not effected by importers' attitudes to price risk, I can define an "effective price" (or "adjusted price") as $p_i^* = p_i/m_i$, indicating the expenditure function can be specified as

(6) $c(u, p, v) = c(u, p^*)$

Accordingly, (1) can be restated as:

(7) $\log c(u, p^*) = a(p^*) + u\, b(p^*)$,

where $p*$ is a vector with the components like p_i^*.

Second, the baseline effect means price risk takes effect on the baseline expenditure shares, which can be performed via adjusting the price index since the baseline imports are originally related to this index (Duffy 1995). Consequently, decomposing price risk effect into the competitive effect and the baseline effect indicates an updated price index:

(8) $\log P^* = a_0 + \sum_k a_k \log p_k^* + \sum_k \sum_j \theta_{kj} \log p_k^* \log v_j + 0.5 \sum_k \sum_l \gamma_{kl}^* \log p_k^* \log p_l^*$

Hence, replacing equation (1) and (2) with (6) and (7) results in the risk-augmented AIDS demand system in the form:

(9) $w_i = a_i + \sum_{ij} \theta_{ij} \log v_{ij} + \beta_i \log (x/P^*) + \sum_j \gamma_{ij} (\log p_j - \delta_j \log v_j)$

Rearranging terms in (9) yields the theoretical model:

(10) $w_i = a_i + \beta_i \log (x/P^*) + \sum_j \gamma_{ij} \log p_j + \sum_{ij} \lambda_{ij} \log v_j$

where $\lambda_{ij} (= \theta_{ij} - \gamma_{ij}\delta_j)$ is the combined effect of price risk. For risk-averse importers, the competitive effect of own-risk is negative since δ_i is negative and γ_{ii} is also negative due to the law of demand. Considering $|\delta_i| > 1$, if $\theta_{ii} \leq 0$, the parameter of own-risk in equation (10), i.e. λ_{ii}, should be greater than the own-price effect (γ_{ii}) in absolute value. Therefore, the baseline effect is positive as long as the estimated γ_{ii} is bigger than λ_{ii} in absolute value. The opposite is true if γ_{ii} is smaller than λ_{ii} in absolute value. For cross effects, the sign of γ_{ij} ($i \neq j$) depends on substitutability between goods i and j, indicating an undetermined sign of the cross-risk effect.

3.4 Empirical Model and Data

The empirical model of the US codfish import demand is specified by incorporating an error term into (10) to obtain:

(11) $w_{i,t} = a_i + \beta_i \log (x_t/P_t^*) + \sum_{j=1}^{4} \gamma_{ij} \log p_{i,t} + \sum_{j=1}^{3} \lambda_{ij} \log v(p_{j,t}) + u_{i,t}$

where i represents supplier (Canada = 1, China = 2, Iceland = 3, and ROW = 4); t stands for the time subscript (monthly); u_i is an i.i.d. error term; and P^*, as a common in the literature (e.g. Duffy 2003), is measured in the Stone form:

(12) $\log P_t^* = \sum_{j=1}^{4} w_{j,t} \log p_{j,t}$.

All other terms and variables are as previously defined.

Only price volatilities of cod from the top 3 suppliers, i.e. Canada, China, and Iceland, are taken into account in the demand system. Except for the major suppliers, there were lots of runners-up in the US codfish market, so the fluctuation of the price of ROW was smoothed during the sample period, as confirmed in figure A1. Furthermore, as discussed in the latter sector, the statistical test rejected the existence of a conditional volatility of the ROW price.

To satisfy properties of the demand system, namely homogeneity, symmetry, and adding-up, the following restrictions are imposed:

(13) $\sum_{j=1}^{4} \Upsilon_{ij} = 0$ (homogeneity)

(14) $\gamma_{ij} = \gamma_{ji}$ (symmetry)

(15) $\sum_{i=1}^{4} \alpha_i = 1$, $\sum_{i=1}^{4} \beta_i = \sum_{i=1}^{4} \Upsilon_{ij} = \sum_{i=1}^{4} \lambda_{ij} = 0$ (adding-up)

After estimating the demand system represented by (11), the estimate parameters can be used to derive the Marshallian elasticities

(16) $E_{ij}^{x} = \beta_i / \overline{w_i}$ (income elasticity)

(17) $E_{ii} = -1 + \gamma_{ii} / \overline{w_i} - \beta_i$ (own-price elasticity)

(18) $E_{ij} = \gamma_{ij} / \overline{w_i} - (\beta_i \overline{w_j}) / \overline{w_i}$ (cross-price elasticity)

(19) $E_{ij}^{v} = \lambda_{ij} / \overline{w_i}$ (price-risk elasticity)

where $\overline{w_i}$ is the expenditure share of commodity i in the base year (to be discussed in detail).

Monthly cod import data from January 1989 to December 2010 are employed to estimate the risk-augmented AIDS model represented by (11). The data are from the US International Trade Committee (USITC) and the US National Marine Fishery Sources (NMFS). Import values are in the US dollar on a cost-insurance-freight (CIF) basis and import quantities are measured in unit of kilograms, resulting in the unit of prices in US$ per kilogram (dividing value by quantity for each month). For a particular month, expenditure share of the *i*th country (w_i) is calculated by dividing import value for country i by total import value in this month. Further, price and expenditure time series are normalized to one (with respect to the base year 2004) in order to minimize the estimating problems due to the Stone Index approximation (Pashardes 1993, Moschini 1995).

3.5 Price Volatility

Before estimating the import demand system, the moments of import price distribution must be estimated. The methods of measuring risk have evolved over time, and they can fall roughly into two categories: the unconditional volatility represented by the moving standard deviation and the conditional volatility represented by the generalized autoregressive conditional heteroskedasticity method (GARCH). While a clearly dominant method has not yet emerged in the empirical research (Bahmani-Oskooee and Hegerty 2007), the conditional volatility has become more popular after the advent of cointegrating analysis because: (*i*) the GARCH model accounts well for the heavy tails of the distribution of the original variable (Kandilov 2008); (*ii*) this method allows for time-varying conditional variance (i.e. volatility clustering) in the original variable (Wang and Barrett 2007); and (*iii*) the GARCH model can test whether the movement in the conditional variance of a variable over time is statistically significant or not. These advantages have led GARCH-type models to be very efficient in evaluating price uncertainty (Rezitis and Stavropoulos 2011). Given the correlation between Canada's price, China's price, and Iceland's price, the Multivariate-GARCH method (M-GARCH) rather than the univariate GARCH method is applied in the present paper.

One condition for the implement of the GARCH-type model is the existence of an ARCH effect in the price dynamics. I first test the individual and joint ARCH effects of the three codfish import prices by applying the Lagrange Multiplier (LM) method.[16] The sample value is larger than the LM critical value at the 5% level of significance for each individual test and the joint test (Table A1). Thus, the M-GARCH method is employed in the present paper to obtain the proxy for price risk.

The M-GARCH model is expanded from the univariate GARCH approach by accounting for the conditional correlation between innovations from different price data generating processes. In the paper, I focus on uncertainty in the prices of cod from Canada, China, and Iceland, indicating a 3-dimensional vector GARCH (1,1) process such that

(20) $V_{\mathrm{t}} = a_0 + A\, E_{t-1}^{(2)} + B\, V_{t-1}$

where V_{t} is a vector (3×1) of conditional price variance, $E_{t-1}^{(2)}$ is a vector (3×1) of the squared error terms from mean regressions, and others are parameters to be estimated. Given different assumptions of the conditional correlations, there are many types of M-GARCH specification in the literature. In the present paper, I apply the DCC-GARCH model of Robert Engle (2002) since it presumes time-varying conditional correlations, and is less restrictive.

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[16] For an individual ARCH test, I first regress one price variable on its lags to obtain the residual. Then I regress residual on m lags, and assess joint significance of coefficients of lag variables. If the coefficients are different from zero based on LM test then the null of conditional homoscedasticity can be rejected.

Figure 1 presents the estimated conditional variance for each price series by applying the DCC-GARCH approach (table A2). Canada and China's prices had high degrees of variation throughout the period. Consistent with the pattern of price movements reflected in figure 1, Canada's price became more volatile in the latter period; however, China's price had more fluctuation in the early period. For Iceland, except for several extreme observations, the price volatility had a considerably low degree of fluctuation. Extreme observations also emerged in the estimated Canadian and Chinese price variances. In order to avoid distortion impacts of extreme observations when justifying the impact of price volatility, I compute the mean values of price volatility variables for 2000-2002 when China's share began to rise and for 2008-2010 when China finally dominated the market. By a comparison of those two periods, Canada's price volatility increased as high as 5.7 times, and Iceland's price volatility increased 7.4 times. On the contrary, China's price volatility reduced by 70%. Intuitively, price uncertainty could be one of the crucial determinants explaining the cod trade pattern during the sample period.

3.6 Econometric Procedure

In a multivariate time series context, the first step is to test the stationarity of each variable. Properties of the data series generating processes dictate the methods to verify a common stochastic trend among the data series. Given the existence of cointegration among variables, the long-run relation between variables can be estimated by applying the Vector Error Correction Model (VECM). If any linear combination of these variables fails to result in common stochastic trends, the SUR estimates of the demand system may be appropriate after transforming variables to be stationary according to the properties of the data generating processes.

Prior to the cointegration test, the stationarity properties of variables are evaluated by implementing the Augmented Dickey Fuller (ADF) test. The order of augmentation in the ADF test is determined by the Akaike Information Criterion (AIC). If the null hypothesis cannot be rejected, I can conclude that the data series is integrated of order one, I(1). The outcomes of the ADF test suggest (table A3), at the conventional percent level, the null hypothesis cannot be rejected for almost all variables, indicating most variables included in the models are integrated of order one.

After evaluating the stationarity of each variable, attempts are made to test for the presence of common stochastic trends among the variables. As most variables included in the demand system are I(1), the Johansen procedure is implemented to test for cointegration in the import demand expression. This method is based on an unrestricted vector autoregressive regression (VAR) approach, and hence all variables incorporated in the model are assumed to be endogenous. The test results produced from the Johansen trace and eigenvalue tests are reported in table 2. Both the trace test and the eigenvalue test indicate five cointegration relationships between variables in the demand system, which contradicts the expected ranks of three cointegration spaces upon the demand theory. With a further inspection of the test results, the sample values of the null hypothesis $r = 5$ in the eigenvalue test and the null hypotheses $r = 5$ and $r = 4$ in the trace test are only marginally larger than the critical value at the 5% level. On balance, therefore, the predictions from the theoretical framework and the moderate support from the econometric test results ensure the three cointegrating relationships between variables in the demand system.

Since the Johansen test, which is based on an unrestricted VAR approach, just confirms the cointegrating space, economic constraints need to be imposed on VAR to obtain the identified long-run relationship. At the same time, the short-run adjusting mechanism is controlled by an error correction process. Thus, the VECM approach is employed to estimate the demand system represented by equation (11). Subsequently, I restate the demand system in a VCEM form:[17]

(21) $\Delta Y_t = \alpha\beta`Y_{t-1} + \Gamma_1\, \Delta Y_{t-1} + \ldots + \Gamma_p\, \Delta Y_{t-g} + \mu D_t + U_t$

where Y_t is a 11×1 vector of all variables in the demand system, i.e. $w_{1,t}$, $w_{2,t}$, $w_{3,t}$, $\log(x_t/P_t^{*}*)$, $\log p_{1,t}$, $\log p_{2,t}$, $\log p_{3,t}$, $\log p_{4,t}$ $\log v(p_{1,t})$, $\log v(p_{2,t})$, and $\log v(p_{3,t})$; D_t is a vector of dummy variables utilized to control seasonal adjustments in the trade pattern; g represents the number of lags for the first difference variables. Here, α is the loading matrix (11×3), matrix β (11×3) has the long-run coefficients, Γ_i (11×11) is short run parameters matrix, and μ is a matrix of coefficients of dummy variables. As usual, U_t is a vector of error term.

Considering the existence of three cointegrating relationships in the unrestricted VAR model, I need 3×3 = 9 constraints to identify each demand equation (Iootty, Pinto Jr, and Ebeling 2009). Hence, besides three normalizing constraints, three homogeneity constraints and three symmetry constraints are imposed on VECM when estimating the long-run economic demand relationship.

3.7 Regression Results

Table 3 reports the estimated long-run coefficients (β*'s*) of the US codfish import demand system expressed in the VECM specification, i.e. equation (21). For the purpose of comparison, the SUR estimates of the static demand system are also regressed (Table A4). The VECM

[17] The Akaike criterion proposed a lag order of 3 while the Schwarz information criterion and Hannan-Quinn criterion proposed a lag order of 1, so the lag order of 1 is chosen in order to avoid unstable estimate results after considering 11 variables in the system.

estimated coefficients are more significant and larger in magnitude when compared to the SUR estimates. Moreover, there is strong evidence of autocorrelation in the residues of SUR regression. Therefore, the remaining discussion will rely on the regression results from VECM.

The regression results overall are satisfactory in that most of the estimated parameters are statistically significant. Consistent with the trade theory, the estimated coefficients of own-price are significant with negative signs for all countries, suggesting that an improvement in competitiveness (a lower relative price) yields an increase in expenditure share. The 3 own-risk effects, in the range between -0.11 and -0.18, are significantly negative, implying the competitive effect dominates the sign of the risk factor. What is an interesting finding is that, in each case, the coefficient of own-price is greater than the corresponding coefficient of own-risk, indicating, as discussed in the previous sector, the baseline effect of the price risk is positive. Both positive and negative coefficients of cross-risk factors are captured in the regression, and 9 out of the 12 cross effects are substantially different from zero. Moreover, when the cross-price effect is positive (negative), the corresponding cross-risk effect is also positive (negative). Given a positive expenditure effect, the positive cross-price effect indicates the two goods are substitutes for each other. Thus, a rise in the fluctuation of one good's price would decrease its marginal utility and consequently improve the marginal utility of its substitute, leading to a positive effect of the cross-price variance, and vice versa for complementary goods.

Considering the lack of economic meaning of the estimated coefficients in the demand system, I derive expenditure elasticities (E_i^x), price elasticities (E_{ij}), and volatility elasticities (E_{ij}^v) by employing suppliers' expenditure shares in the base year, 2004. The estimated elasticities are displayed in Table 4.

Imports from Canada, China, and Iceland are more sensitive to changes in total imports. For example a one-percentage increase in the conditional expenditure would increase imports from Canada and China by 2.0% and 4.5%, respectively. The expenditure elasticity of Iceland's cod is negative (E_3^x = -4.5), suggesting that imports from Iceland would increase when the total expenditure decreases. Since 1995, when China entered into the US codfish market substantially (expenditure share > 1), the total expenditure has kept a rising trend until 2008. Hence, China's dominating market role is first due to changes in import expenditure of codfish.

In general, US codfish demand is strongly sensitive to changes in price. China's codfish has the greatest price elasticity (E_{22} = -11.7), followed by Iceland (E_{33} = -8.5), and Canada (E_{11} = -3.9). This suggests, taking China as an example, a one-percentage decrease in China's price would increase cod imports from China by 11.7%, holding other determinants constant. Cross-price elasticities reflect the pattern commonly found in trade literature. Most of the off-diagonal elements are positive, implying gross substitutes between cod from different sources. A one-percentage increase in Canada and Iceland's prices would raise imports from China by 1.1% and 6.0%, respectively. On the other hand, a one-percentage increase in China's price only benefits Canada by a 2.5% increase of import volume, and a one-percentage increase in Iceland's price would further drop Canada's imports by 0.96%.

Turning to price volatility elasticities (E_{ij}^v), the main concern in the present paper, all own-risk elasticities are statistically significant with expected negative signs. This means the competitive effect of the risk factor dominates the sign of the risk factor, though the baseline effect of the risk factor is positive. For example, a 1% increase in Canada's price volatility is estimated to decrease the import demand for Canada's cod by 0.51%, given the other variables are constant. For Iceland, a 1% increase in price volatility should reduce import demand for

Iceland's cod by 0.85%. The US importers are less sensitive to the fluctuation of China's price, considering the smallest own-risk elasticity (E_{22}^{v} = -0.32). The own-risk elasticities are much smaller than the corresponding own-price elasticities, -0.51 vs. -3.88 in the case of Canada, -0.32 vs. -11.7 in the case of China, and -0.85 vs. -8.52 in the case of Iceland. I further evaluate the extent to which price and risk may have contributed to the rising market share of China at the expense of Canada. During the sample period, the great annual growth rate of China's share happened in 2005. China's share increased 18% from 2004 to 2005, while Canada lost 9%. In 2005, the price of Canada's cod was increased by 0.59%, resulting in a 2.3% decrease of volume of cod imported from Canada. For China, the price of cod was increased by 5.6%, shrinking imports from China by 66%, *ceteris paribus*. Unlike changes in price, Canada's price volatility in 2005 was as large as 1.4 times of the magnitude in 2004. On the contrary, China's price was relatively stable as the magnitude of volatility decreased by 29% in 2005. Considering the estimated own-risk elasticities, the increased price volatility would reduce imports from Canada by 69%; however, the declined volatility would increase demand for China's cod by 9.3%, which offset the negative effect of the rising price in the same year.

The import demand for codfish is also sensitive to the changes in the cross-risk factors. A one-percentage increase in volatility of China's cod price and Iceland's cod price would increase imports from Canada by 0.13% and 0.36%, respectively. Imports of China's cod are only sensitive to Canada's price volatility (E_{21}^{v} = 1.28). Codfish from Iceland only responds negatively to Canada's price volatility, though the corresponding cross-price elasticity is not significant. Taking the annual growth rates (2004-2005) of Canada and China' price volatilities into account (136% and -29%, respectively), *ceteris paribus*, changes in Canada's price volatility would increase cod imports from China 1.7 times, which completely offset the negative effect of the rising own-price. In contrast, China's relatively stable price would further reduce cod imports from Canada by 3.9%.

Overall, the inclusion of risk factors in the demand system can adequately track the observed trade pattern. In the case of China, influences of the low volatility of own-price and the high volatility of cross-prices offset the negative effect of the rising own-price and contribute substantially to the rising trend of codfish imported from China during the sample period. By way of comparison, for Canada, in spite of a low degree of changes in the annual average price, the high fluctuation of price and the strong substitutability between Canada and China's cod are the main reasons explaining the downward trend of cod imports from Canada during the sample period.

3.8 Summary and Implications

In the present paper, I developed a risk-augmented Almost Ideal Demand System (AIDS) model to explore the extent to which risk factors explain the observed trade pattern. In the extended model, the risk effect on import demand is decomposed into the competitive effect and the baseline effect. Taking nonstationarities in the data and endogeneity into account, I employ a cointegrating-based Vector Error Correction Model (VECM) approach to estimate the long-run responses of imports to the changes in the one or several determinants.

The resulting model is utilized to the US codfish market and the estimate results of the demand system inclusive of price volatility adequately revealed formulation of the observed trade pattern. First, the estimated VECM long-run responses of importers to expenditure, price, and volatility are more significant and larger in magnitude when compared to the Seemingly Unrelated Regression (SUR) estimates, implying the static model may understate the responses of import to changes in expenditure, price, and volatility. On the basis of the VCEM regression results, it can be argued that in the long run, the price volatility exerts significant effects on importers' strategic decision to allocate expenditure among different source suppliers. For example, the low fluctuation of China's price and the high fluctuation of the competing suppliers' prices contributed substantially to China's rapid rise as a major supplier of codfish imported to the US after 2004; whereas, there was a huge negative effect of the rising price on cod imports from China due to the extremely elastic demand curve.

The policy implications of these results for agriculture trade would likely seem to be straightforward. Exporting countries interested in a target market should attempt to stabilize the price. However, as Wolak and Kolstad (1991) state, much of price uncertainty is induced by factors, which are out of the control of suppliers. Furthermore, the negative relationship between price uncertainty and expenditure share limits abilities of runners-up to reduce price uncertainty. Consequently, perhaps a more reasonable implication is related to trade policy designing which, in most cases, is essentially based on price analysis. For example, regarding the antidumping policy, the relationship between the antidumping duties and price risk is investigated in Blonigen (2004), which is cited in Carter and Gunning-Trant (2011, p. 99) as follows: "Substantial price volatility in agricultural markets often leads to higher AD margins compared with those of manufacturing, particularly when the product is highly perishable."

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Appendix 1: Tables and Figures for Chapter I

Table 1.1 Prices, Quantities, and Market Shares for Domestic and Imported Frozen Catfish Fillets, United States, 1999-2010

Year	Price (CIF $/lb.)[a]			Quantity (mil. lbs)			Market Share		
	P_1	P_2	P_3	Q_1	Q_2	Q_3	S_1	S_2	S_3
1999	2.76	1.96	1.22	120	7	1	0.95	0.04	0.005
2000	2.82	1.66	1.48	120	19	1	0.91	0.09	0.005
2001	2.61	1.39	1.37	115	30	1	0.87	0.12	0.004
2002	2.39	1.45	1.17	131	46	1	0.82	0.18	0.002
2003[b]	**2.41**	**1.41**	**1.51**	**125**	**38**	**1**	**0.84**	**0.15**	**0.005**
2004	2.62	1.53	1.68	122	44	3	0.82	0.17	0.011
2005	2.67	1.46	1.52	124	36	10	0.83	0.13	0.04
2006	2.92	1.60	1.74	118	50	33	0.72	0.16	0.12
2007	2.92	1.58	1.81	104	36	40	0.70	0.13	0.17
2008	2.91	1.57	1.76	103	53	44	0.65	0.18	0.17
2009	2.96	1.55	1.72	96	86	42	0.58	0.27	0.15
2010	2.97	1.52	1.79	98	108	28	0.58	0.32	0.10

[a]1 = US, 2 = Vietnam, 3 = Rest of World (China mainly). Vietnam and ROW prices are measured by dividing CIF value by quantity.
[b]Year in which antidumping duties were imposed on imports from Vietnam.
Sources: Hanson and Sites (2009), NOAH Fisheries (2011), and USITC (2011)

Table 1.2 Simulated Duty Pass-through Elasticities[a]

Armington Elasticity	Price Pass-Through	Duty Pass-Through		
	P_1^*/P_2^{*b}	$\widetilde{P}_2^* / T^{*b}$	P_1^* / T^{*b}	P_3 / T^{*b}
	2002-2004 ($S_3 = 0.01$)			
$\sigma = 1.5$	0.038	0.588	0.022	0.017
$\sigma = 2.5$	0.106	0.486	0.052	0.041
$\sigma = 5.0$	0.240	0.361	0.087	0.076
	2005-2007 ($S_3 = 0.11$)			
$\sigma = 1.5$	0.033	0.586	0.019	0.014
$\sigma = 2.5$	0.092	0.480	0.044	0.035
$\sigma = 5.0$	0.211	0.350	0.074	0.064
	2008-2010 ($S_3 = 0.14$)			
$\sigma = 1.5$	0.058	0.597	0.035	0.026
$\sigma = 2.5$	0.155	0.507	0.079	0.063
$\sigma = 5.0$	0.325	0.399	0.130	0.113

[a]The supply elasticities set to $\varepsilon_1 = 1.1$, $\varepsilon_2 = 2.0$, and $\varepsilon_3 = 2.0$
[b]Computed using text equations (10) - (12) and (13).

Table 1.3 Distributional Effects of a 35% Catfish Antidumping Duty Under Alternative Assumptions About the Market Share of the Non-Dutied Good

Group	Formula	Units	Market Share of Non-Dutied Good[b]			Percent Change
			$S_3 = 0.07$	$S_3 = 0.14$	$S_3 = 0.28$	
U.S. Producers	ΔPS_1	Mil $	9.15	8.14	6.10	-33.4
U.S. Consumers	$\Delta CS = \Delta CS_1 + \Delta CS_2 + \Delta CS_3$	Mil $	-28.6	-28.3	-27.7	-3.16
U.S. Treasury	ΔTR	Mil $	29.2	29.2	29.1	-0.19
Total U.S. Impact	$\Delta W_{US} = \Delta PS_1 + \Delta CS + \Delta TR$	Mil $	9.69	8.96	7.48	-22.8
Vietnam Producers	ΔPS_2	Mil $	-18.1	-18.1	-18.2	0.28
China Producers	ΔPS_3	Mil $	0.77	1.51	3.00	289
Total Impact	$\Delta W_{US} + \Delta PS_2 + \Delta PS_2$	Mil $	-7.66	-7.67	-7.69	0.36
Incidence of Duty:						
U.S. Consumers	$\tilde{P}_2^* / T^*$	%/100	0.508	0.507	0.506	-0.35
Vietnam Producers	P_2^* / T^*	%/100	-0.492	-0.493	-0.494	0.36
Redistribution Efficiency	$\Delta PS_1 / \Delta CS_1$		0.32	0.29	0.22	-31.2

[a]For details, see text equations (25) - (28) and attendant discussion.

[b]Except for market shares, other parameters are set be "best-bet" values. Market share of the dutied good is held constant at $S_2 = 0.26$, the mean value for 2008-10. The middle value for the non-dutied good, $S_3 = 0.14$, is the mean value of this parameter over the 2008-10 period.

Table 1.4 Sensitivity of Welfare Effects of the Catfish Antidumping Duty to the Armington Elasticity (σ) and Import Supply Elasticities ($\varepsilon = \varepsilon_2 = \varepsilon_3$)[a]

Group	$\sigma = 1.5$			$\sigma = 2.5$			$\sigma = 5$		
	$\varepsilon = 2$	$\varepsilon = 4$	$\varepsilon = \infty$	$\varepsilon = 2$[b]	$\varepsilon = 4$	$\varepsilon = \infty$	$\varepsilon = 2$	$\varepsilon = 4$	$\varepsilon = \infty$
U.S. Producers	3.56	4.42	5.85	8.11	10.6	15.2	13.5	18.7	30.3
U.S. Consumers	-27.0	-32.3	-39.8	-28.3	-35.0	-44.3	-30.3	-39.3	-49.3
U.S. Treasury	32.0	28.8	23.4	29.2	24.1	13.7	25.8	17.4	-6.54
Total U.S. Impact	8.50	0.89	-10.5	8.94	-0.37	-15.4	8.96	-3.20	-25.5
Vietnam Producers	-15.4	-9.24	0.00	-18.1	-11.2	0.00	-21.1	-13.5	0.00
China Producers	0.62	0.49	0.00	1.53	1.40	0.00	2.80	3.11	0.00
Total Impact	-6.27	-7.86	-10.5	-7.67	-10.2	-15.4	-9.36	-13.6	-25.5
Incidence of Duty:									
U.S. Consumers	0.60	0.75	1.00	0.51	0.67	1.00	0.40	0.56	1.00
Vietnam Producers	-0.40	-0.25	0.00	-0.49	-0.33	0.00	-0.60	-0.44	0.00
Redistribution Efficiency	0.13	0.14	0.15	0.29	0.30	0.34	0.45	0.48	0.61

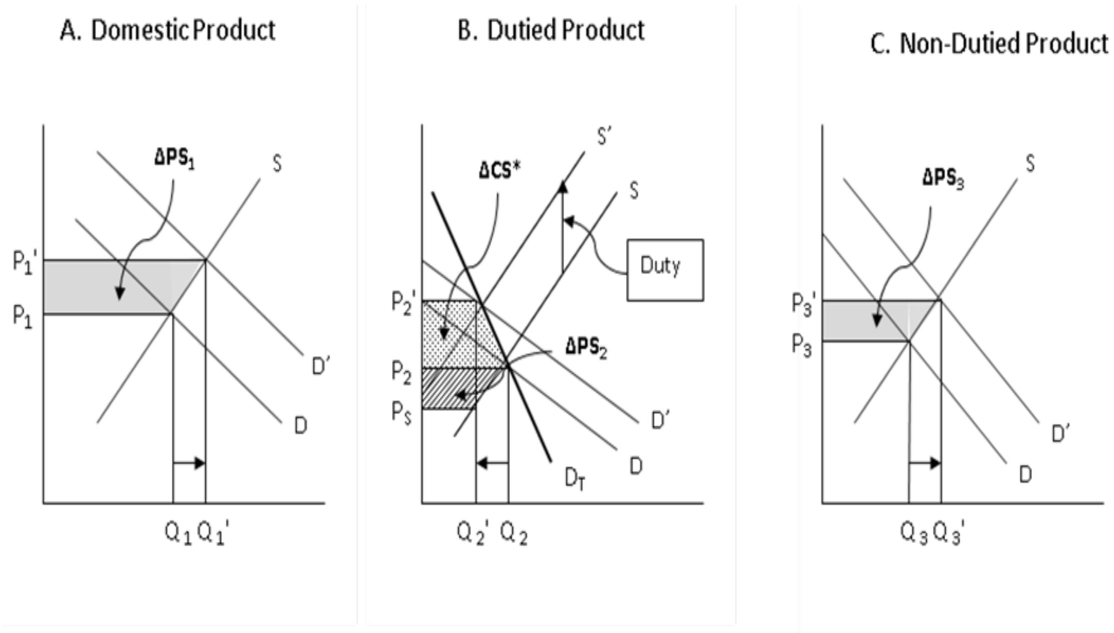
[a]The domestic supply elasticity is set to $\varepsilon_1 = 1.1$ and market shares are set to $S_1 = 0.60$, $S_2 = 0.26$ and $S_3 = 0.14$, the mean values for 2008-10.
[b]"Best-bet" estimates.

Table 1.5 Sensitivity of Trade-Diversion Effects to Antidumping Duty Rates[a]

Group	$S_3 = 0.07$[b]				$S_3 = 0.14$[b]				$S_3 = 0.28$[b]			
	T* = 15%	T* = 35%	T* = 50%	T* = 70%	T* = 15%	T* = 35%	T* = 50%	T* = 70%	T* = 15%	T* = 35%	T* = 50%	T* = 70%
U.S. Producers	3.89	9.15	13.2	18.6	3.46	8.14	11.7	16.5	2.59	6.10	8.76	12.4
U.S. Consumers	-13.2	-28.6	-38.6	-49.8	-13.1	-28.3	-38.2	-49.2	-12.8	-27.7	-37.3	-48.0
U.S. Treasury	16.3	29.2	32.3	27.7	16.3	29.2	32.3	27.6	16.2	29.1	32.2	27.5
Total U.S. Impact	6.96	9.69	6.83	-3.52	6.65	8.96	5.77	-5.05	6.04	7.48	3.62	-8.14
Vietnam Producers	-8.69	-18.12	-23.58	-28.71	-8.70	-18.1	-23.6	-28.7	-8.72	-18.2	-23.6	-28.8
China Producers	0.33	0.77	1.11	1.58	0.64	1.51	2.18	3.08	1.27	3.00	4.33	6.13
Total Impact	-1.41	-7.66	-15.6	-30.7	-1.41	-7.67	-15.7	-30.7	-1.41	-7.69	-15.7	-30.8
Incidence of Duty:												
U.S. Consumers	0.508	0.508	0.508	0.508	0.507	0.507	0.507	0.507	0.506	0.506	0.506	0.506
Vietnam Producers	-0.492	-0.492	-0.492	-0.492	-0.493	-0.493	-0.493	-0.493	-0.494	-0.494	-0.494	-0.494
Redistribution												
Efficiency	0.29	0.32	0.34	0.37	0.26	0.29	0.31	0.34	0.20	0.22	0.23	0.26

[a]The supply elasticities set to $\varepsilon_1 = 1.1$, $\varepsilon_2 = \varepsilon_3 = 2.0$, and $\sigma = 2.5$, all "best-bet" values.
[b]Market share of the dutied good is held constant at $S_2 = 0.26$, the mean value for 2008-10.

Figure 1.1 Welfare Analysis of an Antidumping Duty



$$\Delta CS^* = \Delta CS_1 + \Delta CS_2 + \Delta CS_3 + \Delta PS_1 + \Delta PS_3$$

Appendix: Derivation of the Pass-Through Elasticity

To derive the PTE, I first reduce the structural model (equations (2) – (8)) to three equations by solving for equilibrium in each market to yield:

(A1) $$(\varepsilon_1 - \eta_{11})P_1^* - \eta_{12}\tilde{P}_2^* - \eta_{13}P_3^* = 0$$

(A2) $$-\eta_{21}P_1^* + (\varepsilon_2 - \eta_{22})\tilde{P}_2^* - \eta_{23}P_3^* = \varepsilon_2 T^*$$

(A3) $$-\eta_{31}P_1^* - \eta_{32}\tilde{P}_2^* + (\varepsilon_3 - \eta_{33})P_3^* = 0.$$

The equilibrium in matrix form is:

(A4) $$\begin{bmatrix} (\varepsilon_1 - \eta_{11}) & -\eta_{12} & -\eta_{13} \\ -\eta_{21} & (\varepsilon_2 - \eta_{22}) & -\eta_{23} \\ -\eta_{31} & -\eta_{32} & (\varepsilon_3 - \eta_{33}) \end{bmatrix} \begin{bmatrix} P_1^* \\ \tilde{P}_2^* \\ P_3^* \end{bmatrix} = \begin{bmatrix} 0 \\ \varepsilon_2 \\ 0 \end{bmatrix} T^*$$

Cramer's rule is applied to equation (A4) to obtain the reduced-form elasticities P_1^*/T^* and $\tilde{P}_2^*/T^*$ (and P_3^*/T^*). Inserting these expressions into

(A5) $$\frac{P_1^*}{T^*} = \left(\frac{P_1^*/T^*}{\tilde{P}_2^*/T^*}\right)\left(\frac{\tilde{P}_2^*}{T^*}\right)$$

gives the PTE when N = 3. The corresponding expression when N = 2 is obtained by the setting the structural elasticities in equations (A1) – (A3) that have a 3 in the subscript to zero, and repeating the above steps.

Appendix 2: Tables and Figures for Chapter II

Table 2.1 US Salmon Imports and Market Shares by Sources

Year	Total Imports (mill. US $)	Value Shares (%)				
		Chile	Canada	Norway	UK	ROW
1995	277	42.1	49.0	5.5	1.9	1.5
1996	316	46.8	44.6	3.6	3.0	1.9
1997	428	43.2	50.1	2.5	2.1	2.1
1998	549	48.9	42.2	3.6	3.4	1.8
1999	689	40.5	39.0	9.6	7.0	3.9
2000	794	53.4	32.1	6.9	5.1	2.5
2001	879	51.4	39.8	4.7	3.1	1.0
2002	947	52.1	39.5	4.7	3.0	0.8
2003	1,072	58.7	28.3	5.2	6.3	1.5
2004	1,032	62.6	26.7	3.8	5.5	1.5
2005	1,184	62.1	29.2	3.3	3.8	1.6
2006	1,471	60.1	28.8	4.7	4.4	2.0
2007	1,579	58.9	26.5	5.6	6.4	2.7
2008	1,562	57.0	28.8	4.0	6.1	4.1
Average		52.7	36.0	4.8	4.4	2.1

Note: Data are obtained from USITC.

Table 2.2 GMM Estimates of U.S. Import Demand System for Salmon, Rotterdam Model Inclusive Import Price Risk, 1995-2008 Monthly Data (1 = Chile, 2 = Canada, 3. Norway, 4 = United Kingdom, 5 = ROW)

Variable	Coef.	Eq 1	Eq 2	Eq 3	Eq 4	Eq 5
Trend	α_i	0.0021	-0.0166 [a]	0.0019	0.0055[c]	0.007 [a]
		(0.0063)	(0.0063)	(0.0019)	(0.0028)	(0.0015)
Δ *ln* Q	θ_i	0.5414 [a]	0.421 [a]	0.029 [a]	0.0187 [c]	-0.010 [a]
		(0.0255)	(0.0302)	(0.0107)	(0.0109)	(0.008)
Δ *ln* p_1	π_{i1}	-0.0072				
		(0.0158)				
Δ *ln* p_2	π_{i2}	-0.0313[b]	-0.179 [a]			
		(0.0142)	(0.0304)			
Δ *ln* p_3	π_{i3}	0.0016	0.102 [a]	-0.0549 [a]		
		(0.0077)	(0.0168)	(0.0111)		
Δ *ln* p_4	π_{i4}	0.0307 [a]	0.0559 [a]	-0.028 [a]	-0.047 [a]	
		(0.010)	(0.011)	(0.0063)	(0.0071)	
Δ *ln* p_5	π_{i5}	0.0062	0.0524 [a]	-0.0208 [a]	-0.0116 [a]	-0.026 [a]
		(0.005)	(0.0059)	(0.0045)	(0.0041)	(0.0035)
Δ *ln* v(p_1)	Υ_1	-1.315[b]				
		(0.515)				
Δ *ln* v(p_2)	Υ_2	-1.0135[a]				
		(0.168)				
Δ *ln* v(p_3)	Υ_3	-0.616[a]				
		(0.223)				
Δ *ln* v(p_4)	Υ_4	-0.165				
		(0.154)				
Δ *ln* v(p_5)	Υ_5	0.180				
		(0.153)				

Note: Numbers in parentheses are asymptotic standard errors. [a] indicates significance at the $p < 0.01$ level; [b] indicates significance at the $p < 0.05$ level; [c] indicates significance at the $p < 0.10$ level.

Table 2.3 Tests of Theoretical Restrictions

Model	Against	Number of Restrictions	LR Statistic	*p*-value
Model B				
Model C	Model B	4	10.75	0.029
Model C′	Model B	3	3.96	0.265
Model D	Model C	1	5.14	0.023
Model E	Model C	2	55.10	<0.0001

Note: Model B restricts the cross-effect of price and exchange risks are zero (Duffy's restriction). Model C and Model C′ impose an identical non-zero effect of all own-price risk and an identical non-zero effect of all own-exchange risk (Theil's restriction), respectively. Model D imposes the equivalence of the non-zero own-price and non-zero own-risk effects. Model E imposes the additional restriction that the identical non-zero effects of risk factor are equal to zero. For each model, symmetry and homogeneity are imposed.

Table 2.4 GMM Estimates of U.S. Import Demand System for Salmon, Rotterdam Model 1995-2008 Monthly Data (1 = Chile, 2 = Canada, 3 = Norway, 4 = United Kingdom, 5 = ROW)

		Model D (with export price and exchange-rate variances)					Model E (without risk factors)				
Variable	Coef.	Eq 1	Eq 2	Eq 3	Eq 4	Eq 5	Eq 1	Eq 2	Eq 3	Eq 4	Eq 5
Trend	α_i	0.003	-0.015[a]	0.002[a]	0.006	0.004[a]	0.006	-0.016	-0.0002	0.005	0.005[a]
		(0.003)	(0.004)	(0.002)	(0.002)	(0.0007)	(0.009)	(0.01)	(0.003)	(0.004)	(0.002)
$\Delta\, ln$ Q	θ_i	0.514[a]	0.419[a]	0.048[a]	0.015[c]	0.004	0.503[a]	0.447[a]	0.041[b]	0.014	-0.005
		(0.017)	(0.023)	(0.009)	(0.008)	(0.0048)	(0.036)	(0.043)	(0.017)	(0.014)	(0.009)
$\Delta\, ln\, p_1$	π_{i1}	-0.16[a]					-0.213[b]				
		(0.033)					(0.06)				
$\Delta\, ln\, p_2$	π_{i2}	0.039[a]	-0.223[a]				0.121[b]	-0.28[a]			
		(0.021)	(0.028)				(0.058)	(0.075)			
$\Delta\, ln\, p_3$	π_{i3}	0.065[a]	0.105[a]	-0.117[a]			0.067[b]	0.064[a]	-0.09[a]		
		(0.019)	(0.016)	(0.012)			(0.027)	(0.024)	(0.019)		
$\Delta\, ln\, p_4$	π_{i4}	0.047[a]	0.03[a]	-0.034[a]	-0.037[a]		0.023[c]	0.049[b]	-0.027[a]	-0.038[a]	
		(0.007)	(0.011)	(0.005)	(0.007)		(0.013)	(0.021)	(0.008)	(0.012)	
$\Delta\, ln\, p_5$	π_{i5}	0.0085	0.05[a]	-0.019[a]	-0.007[b]	-0.033[a]	0.002	0.046[a]	-0.013	-0.007	-0.028[a]
		(0.006)	(0.006)	(0.004)	(0.0029)	(0.002)	(0.011)	(0.011)	(0.006)	(0.005)	(0.004)
Price Risk	$\bar{\gamma}$	-0.073[a]									
		(0.024)									
ER Risk	$\bar{\lambda}$	-0.131[a]									
		(0.024)									

Note: Numbers in parentheses are asymptotic standard errors. [a] indicates significance at the $p < 0.01$ level; [b] indicates significance at the $p < 0.05$ level; [c] indicates significance at the $p < 0.10$ level.

Table 2.5 Conditional Demand Elasticities (1 = Chile, 2 = Canada, 3 = Norway, 4 = United Kingdom, 5 = ROW)

	Expenditure Elasticities	Hicksian Price Elasticities				
Eqn	E_i^y	E_{i1}^*	E_{i2}^*	E_{i3}^*	E_{i4}^*	E_{i5}^*
1	0.975[a]	-0.303 [a]	0.073[c]	0.124 [a]	0.090 [a]	0.016
	(0.033)	(0.063)	(0.04)	(0.036)	(0.014)	(0.012)
2	1.163 [a]	0.107[c]	-0.619 [a]	0.291 [a]	0.084 [a]	0.139 [a]
	(0.063)	(0.059)	(0.077)	(0.044)	(0.03)	(0.018)
3	0.993 [a]	1.350 [a]	2.165 [a]	-2.41 [a]	-0.701 [a]	-0.393 [a]
	(0.188)	(0.395)	(0.331)	(0.25)	(0.101)	(0.082)
4	0.344[b]	1.085 [a]	0.691 [a]	-0.776 [a]	-0.851 [a]	-0.160 [a]
	(0.182)	(0.165)	(0.245)	(0.111)	(0.149)	(0.068)
5	0.194 [a]	0.412 [a]	2.422 [a]	-0.920 [a]	-0.339 [a]	-1.574[b]
	(0.234)	(0.302)	(0.311)	(0.191)	(0.143)	(0.124)

Note: Numbers in parentheses are asymptotic standard errors. [a] indicates significance at the $p < 0.01$ level; [b] indicates significance at the $p < 0.05$ level; [c] indicates significance at the $p < 0.10$ level.

Table 2.6 Conditional Demand Elasticities (1 = Chile, 2 = Canada, 3 = Norway, 4 = United Kingdom, 5 = ROW) (cont)

	Expenditure Elasticities	Hicksian Price Elasticities				
Eqn	E_i^y	E_{i1}^*	E_{i2}^*	E_{i3}^*	E_{i4}^*	E_{i5}^*
1	0.975[a]	-0.303[a]	0.073[c]	0.124[a]	0.090[a]	0.016
	(0.033)	(0.063)	(0.04)	(0.036)	(0.014)	(0.012)
2	1.163[a]	0.107[c]	-0.619[a]	0.291[a]	0.084[a]	0.139[a]
	(0.063)	(0.059)	(0.077)	(0.044)	(0.03)	(0.018)
3	0.993[a]	1.350[a]	2.165[a]	-2.41[a]	-0.701[a]	-0.393[a]
	(0.188)	(0.395)	(0.331)	(0.25)	(0.101)	(0.082)
4	0.344[b]	1.085[a]	0.691[a]	-0.776[a]	-0.851[a]	-0.160[a]
	(0.182)	(0.165)	(0.245)	(0.111)	(0.149)	(0.068)
5	0.194[a]	0.412[a]	2.422[a]	-0.920[a]	-0.339[a]	-1.574[b]
	(0.234)	(0.302)	(0.311)	(0.191)	(0.143)	(0.124)

Note: Numbers in parentheses are asymptotic standard errors. [a] indicates significance at the $p < 0.01$ l indicates significance at the $p < 0.05$ level; [c] indicates significance at the $p < 0.10$ level.

Figure 2.1 Import US Dollar Prices of Salmon by Sources

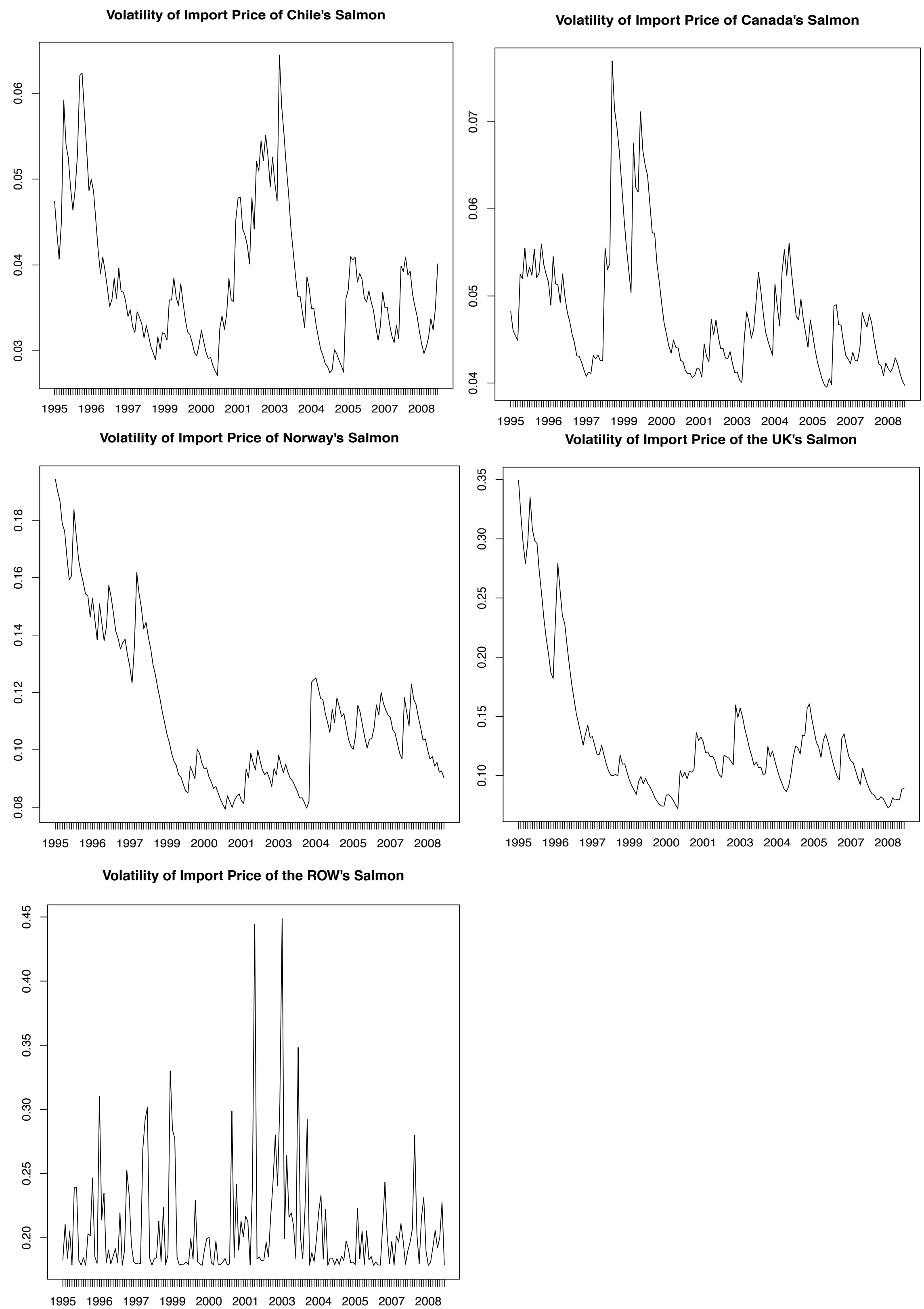


Figure 2.2 Conditional Variance Estimates of Import Prices (in US dollar): January 1995 - December 2008

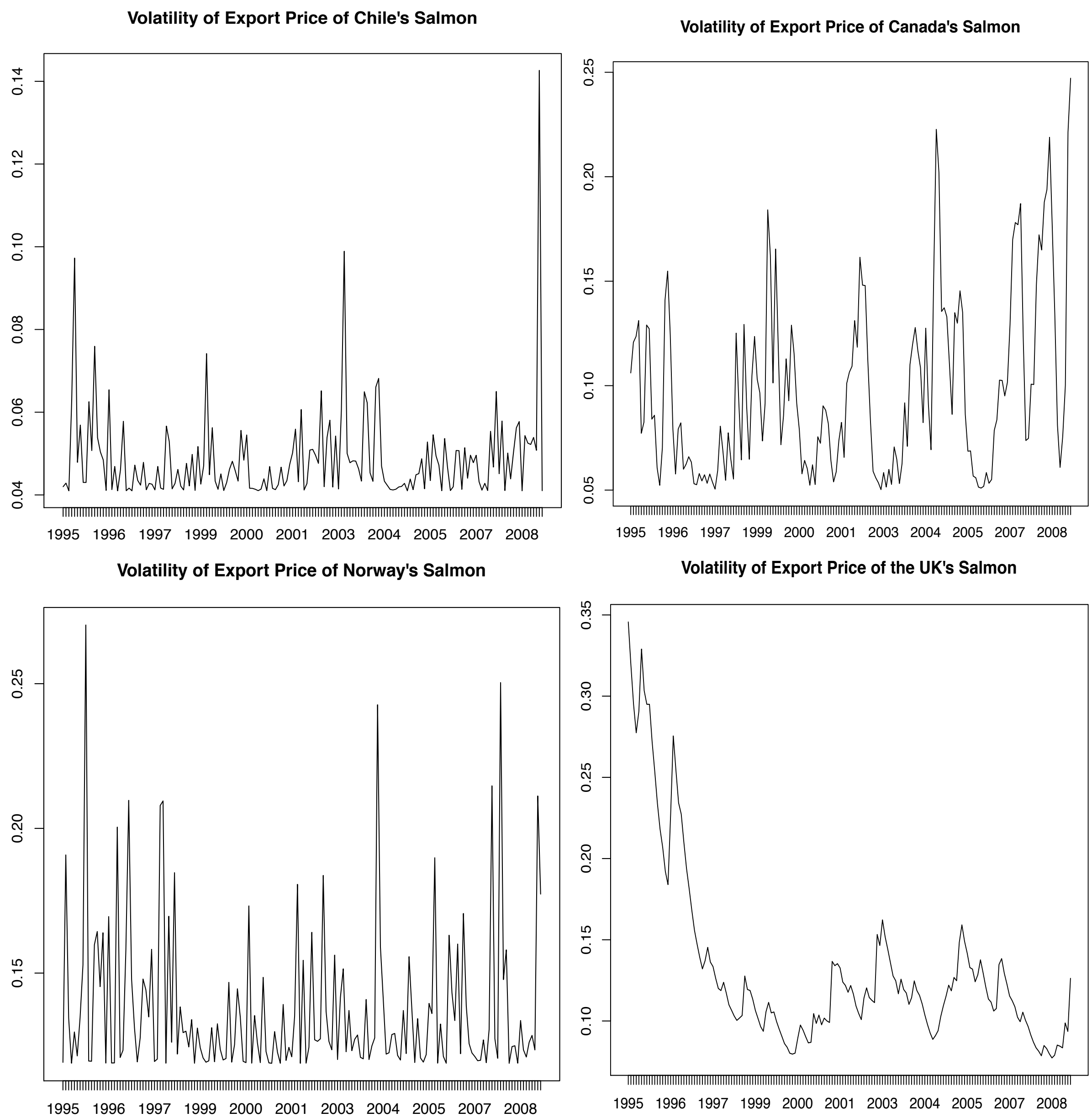


Figure 2.3 Conditional Variance Estimate of Export Prices (in Foreign Currency): January 1995 - December 2008

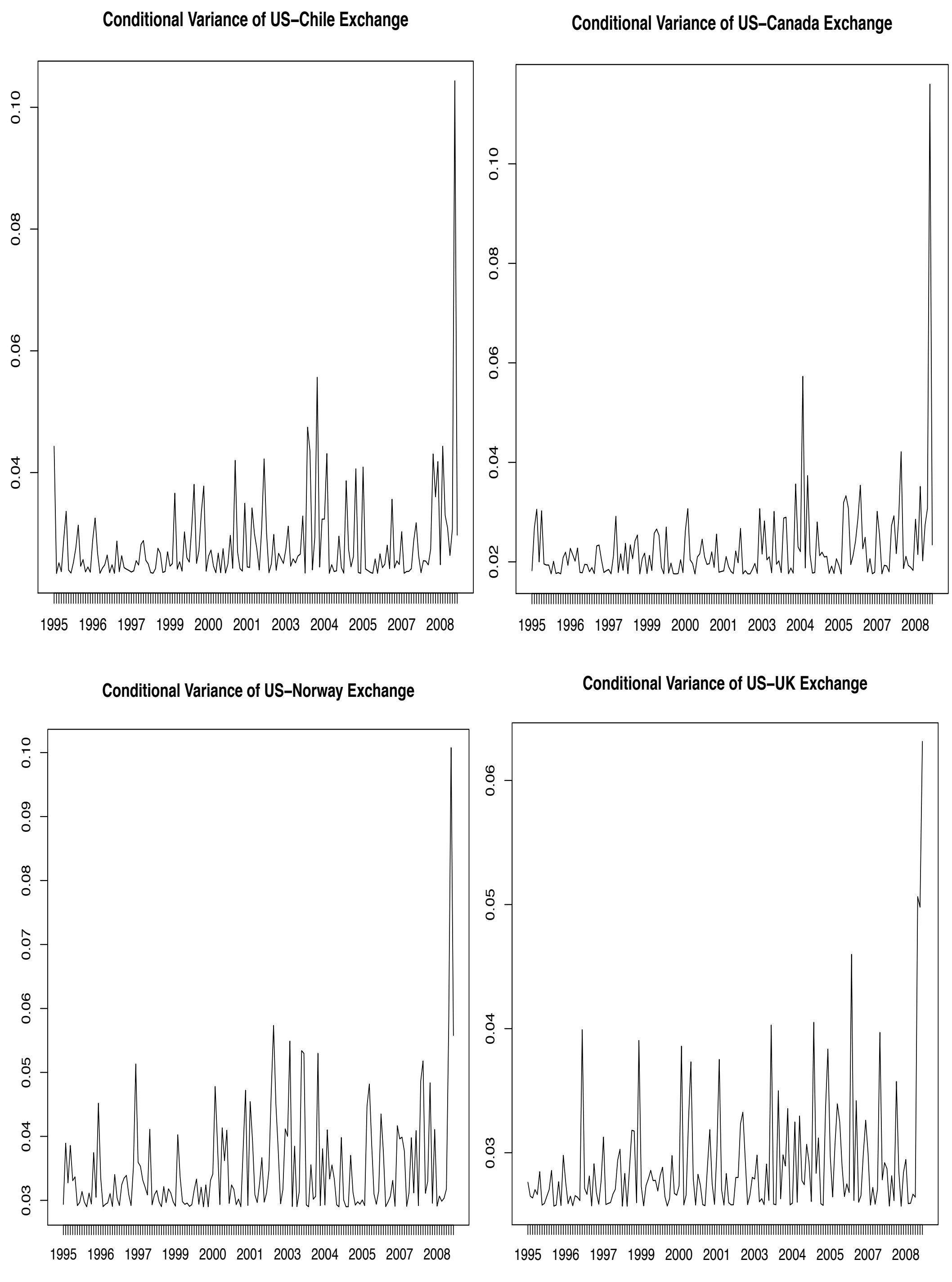


Figure 2.4 Conditional Variance Estimate of Bilateral Exchange Rate: January 1995 - December 2008

Appendix 3: Tables and Figures for Chapter III

Table 3.1 US Cod Imports, Expenditure Shares, Prices by Sources

YEAR	Imports (mill. US $)	Expenditure shares (%)				Prices (US$ / Kilogram)			
		Canada	China	Iceland	ROW	Canada	China	Iceland	ROW
1989	474	61.2	0.2	21.0	17.5	1.4	2.8	2.1	1.3
1990	501	68.9	0.4	15.3	15.4	1.3	2.0	1.6	1.4
1991	510	59.3	0.7	18.4	21.6	1.6	1.8	2.0	1.7
1992	375	49.5	1.6	21.2	27.6	1.5	1.5	1.8	1.6
1993	291	36.2	1.9	36.4	25.4	1.5	1.6	1.7	1.5
1994	262	26.2	1.0	37.0	35.7	1.4	1.9	1.7	1.5
1995	281	24.5	3.1	33.6	38.7	1.3	1.5	1.8	1.6
1996	255	32.9	3.0	37.1	27.0	1.3	1.8	1.7	1.6
1997	318	30.0	5.4	33.6	31.1	1.4	1.7	1.8	1.6
1998	313	30.1	3.6	35.4	31.0	1.5	1.5	2.1	1.9
1999	379	30.5	7.8	36.0	25.6	1.8	1.9	2.2	2.2
2000	355	29.6	13.8	30.3	26.2	1.8	2.2	2.2	2.1
2001	290	28.5	15.7	29.4	26.4	1.7	2.1	2.1	2.0
2002	337	24.3	20.0	26.2	29.5	1.7	2.3	2.2	2.0
2003	315	24.5	23.5	23.3	28.7	1.8	2.3	2.3	2.1
2004	325	21.0	33.1	20.7	25.2	1.9	2.3	2.3	2.2
2005	327	11.8	51.1	15.6	21.5	1.9	2.5	2.7	2.4
2006	349	9.5	61.0	13.6	16.0	2.1	2.9	3.0	2.5
2007	355	8.6	71.1	8.2	12.2	2.3	3.1	3.2	2.7
2008	332	9.7	70.7	6.1	13.5	2.5	3.4	3.0	2.8
2009	250	11.8	63.4	9.5	15.3	2.4	2.7	2.6	2.2
2010	262	10.6	64.1	10.0	15.3	2.2	2.3	2.8	2.4

Sources: USITC and value data are in CIF measure.

Table 3.2 Johansen Cointegration Test Results

Eigenvalue Method				Maximum Eigen Method			
H_o:	H_a:	Eigenvalue	0.05 Critical value	H_o:	H_a:	Max-Eigen Statistic	0.05 Critical value
r = 0	r = 1	545.6*	358.7	r = 0	r ≥ 1	138.3*	79.23
r = 1	r = 2	407.2*	307.2	r ≤ 1	r ≥ 2	93.2*	73.47
r = 2	r = 3	314.0*	260.8	r ≤ 2	r ≥ 3	73.6*	67.77
r = 3	r = 4	240.3*	216.6	r ≤ 3	r ≥ 4	62.1*	61.59
r = 4	r = 5	178.2*	178.0	r ≤ 4	r ≥ 5	56.5*	55.81
r = 5	r = 6	121.6	141.7	r ≤ 5	r ≥ 6	46.2	49.97
r = 6	r = 7	75.4	108.6	r ≤ 6	r ≥ 7	30.3	43.62
r = 7	r = 8	45.0	81.25	r ≤ 7	r ≥ 8	20.5	37.83
r = 8	r = 9	24.5	56.28	r ≤ 8	r ≥ 9	15.9	31.56
r = 9	r = 10	8.56	35.46	r ≤ 9	r ≥ 10	6.79	24.97
r = 10	r = 11	1.77	17.8	r ≤ 10	r ≥ 11	1.77	17.80

*: significant at the 5% level. Critical values are from Pesaran, Shin, and Smith (2000)

Table 3.3 VECM Estimates of US Import Demand System for Codfish, AIDS Model, 1989-2010 Monthly Data (1 = Canada, 2 = China, 3 = Iceland, 4 = ROW)

Variable	Eq 1	Eq 2	Eq 3	Eq 4
constant	0.321***	-0.090	0.503***	0.267***
	(0.045)	(0.174)	(0.121)	(0.033)
log (x/P^*)	0.212***	1.168***	-1.151***	-0.229***
	(0.068)	(0.269)	(0.186)	(0.05)
log p_1	-0.559***	0.603***	-0.157	0.113***
	(0.108)	(0.086)	(0.1)	(0.058)
log p_2	0.603***	-3.134***	2.229***	0.303***
	(0.086)	(0.342)	(0.237)	(0.062)
log p_3	-0.157	2.229***	-1.820***	-0.252***
	(0.1)	(0.237)	(0.188)	(0.057)
log p_4	0.113**	0.303***	-0.252***	-0.164***
	(0.058)	(0.067)	(0.059)	(0.048)
log $v(p_1)$	-0.106**	0.421***	-0.248***	-0.067***
	(0.03)	(0.117)	(0.081)	(0.021)
log $v(p_2)$	0.028*	-0.106*	0.062	0.016
	(0.016)	(0.061)	(0.043)	(0.012)
log $v(p_3)$	0.075***	0.149	-0.178***	-0.047***
	(0.025)	(0.098)	(0.068)	(0.018)

Note: Numbers in parentheses are asymptotic standard errors. Single asterisk (*) indicates significance at the $p < 0.10$ level; double asterisk (**) indicates significance at the $p < 0.05$ level; triple asterisk (***) indicates significance at the $p < 0.01$ level.

Table 3.4 Derived Demand Elasticities (1 = Canada, 2 = China, 3 = Iceland, 4 = ROW)

Eqn	Expenditure Elasticities	Marshallian Price Elasticities				Price Volatility Elasticities		
	E_i^x	E_{i1}	E_{i2}	E_{i3}	E_{i4}	E_{i1}^v	E_{i2}^v	E_{i3}^v
1	2.01***	-3.88***	2.54***	-0.96**	0.29	-0.51***	0.13*	0.36***
	(0.324)	(0.516)	(0.412)	(0.474)	(0.274)	(0.141)	(0.077)	(0.121)
2	4.54***	1.08***	-11.67***	6.01***	0.03	1.28***	-0.32*	0.45
	(0.814)	(0.262)	(1.036)	(0.717)	(0.202)	(0.355)	(0.186)	(0.297)
3	-4.48**	0.41	12.42***	-8.52***	0.17	-1.18***	0.30*	-0.85**
	(0.886)	(0.474)	(1.127)	(0.897)	(0.28)	(0.387)	(0.202)	(0.324)
4	0.08	0.64***	1.51***	-0.82***	-1.43***	-0.27***	0.06	-0.19***
	(0.2)	(0.233)	(0.248)	(0.226)	(0.192)	(0.086)	(0.047)	(0.072)

Note: Numbers in parentheses are standard errors computed via the Delta method. Note: Numbers in parentheses are asymptotic standard errors. Single asterisk (*) indicates significance at the $p < 0.10$ level; double asterisk (**) indicates significance at the $p < 0.05$ level; triple asterisk (***) indicates significance at the $p < 0.01$ level.

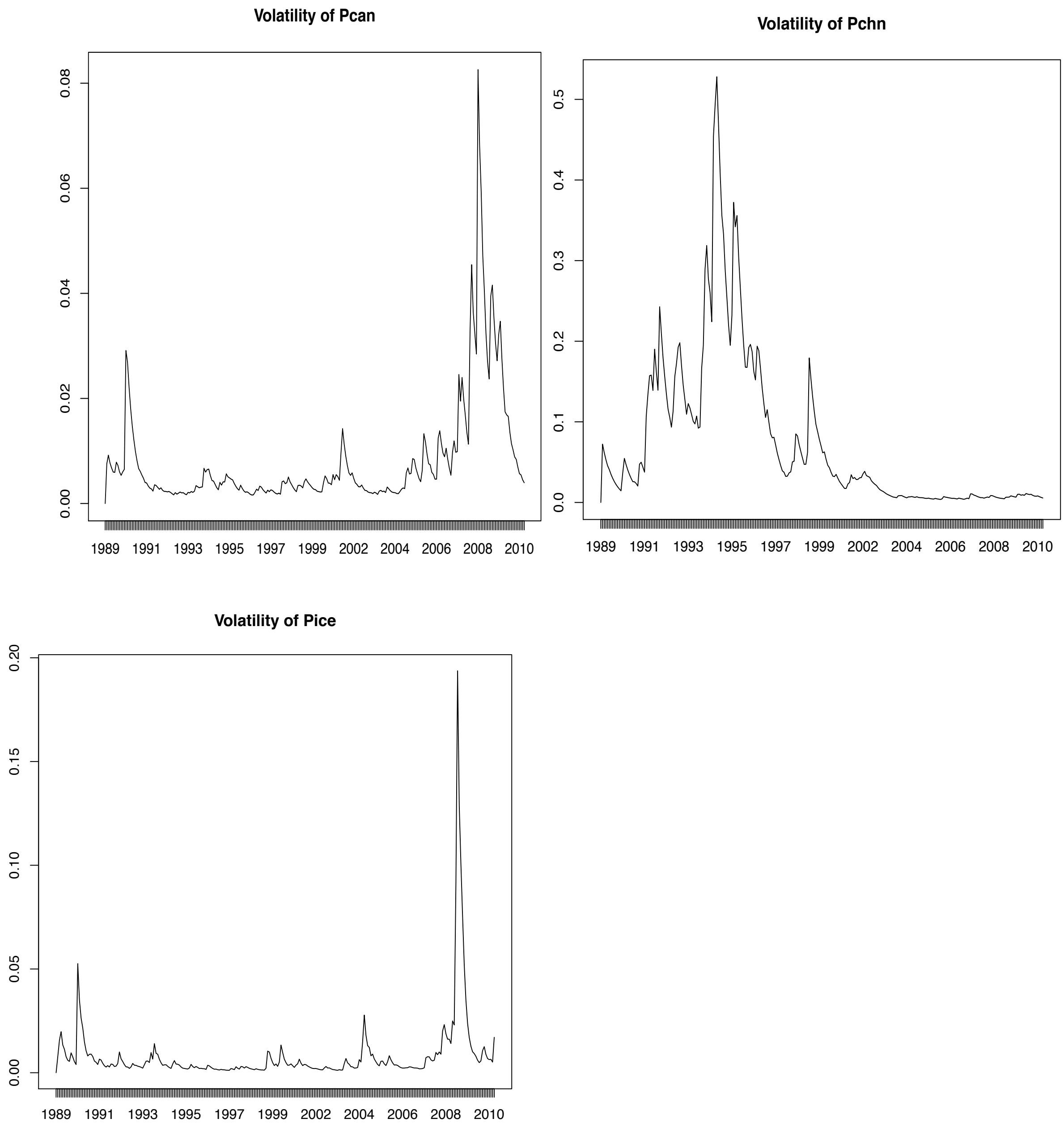


Figure 3.1 M-GARCH Estimated Conditional Price Volatility

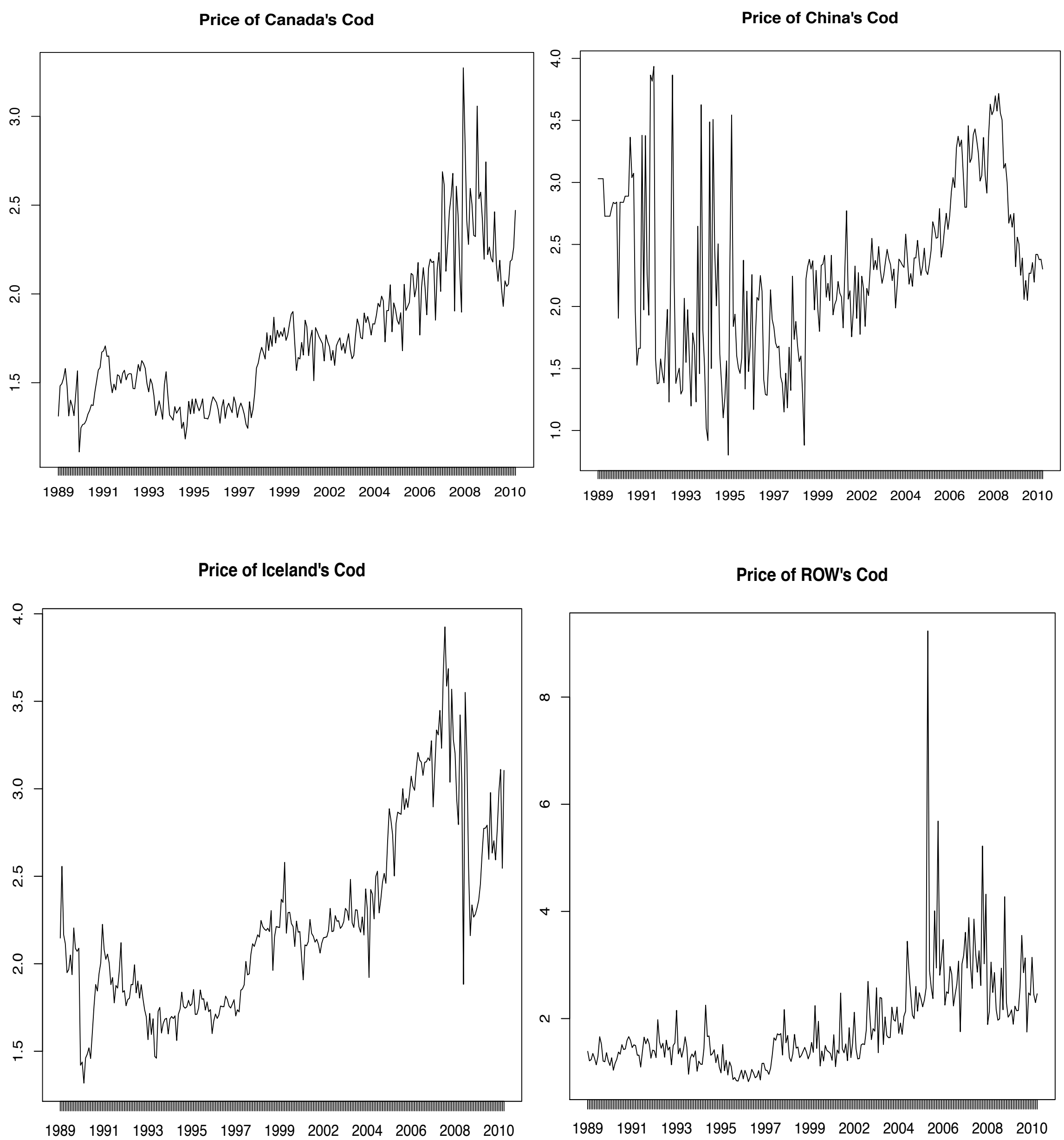


Figure 3.A1 US Import Cod Prices by Sources

Table 3.A1 ARCH Test for Price (1 = Canada, 2 = China, 3 = Iceland)

	Chi-Squared	DF	*P* - value
$\Delta \log p_1$	59	16	<0.00001
$\Delta \log p_2$	56	16	<0.00001
$\Delta \log p_3$	75	16	<0.00001
Joint test	312	180	<0.00001

Note: (1) the null hypothesis is no ARCH effect; (2) the lag length is 16 in the univariate test, and 5 in the multivariate test.

Table 3.A2 M-GARCH Estimates of Conditional Variance of Price (1 = Canada, 2 = China, 3 = Iceland)

	a_i	A_{i1}	A_{i2}	A_{i3}	B_{i1}	B_{i2}	B_{i3}
Equation							
1	0.0001	0.201	0.000002	0.00003	0.7739	0.000003	0.0302
	(0.0002)	(0.0292)	(0.0591)	(0.0693)	(0.0503)	(0.0212)	(0.0886)
Equation							
2	0.0001	0.00000001	0.1474	0.0001	0.000003	0.8546	1E-10
	(0.0648)	(0.0704)	(0.0003)	(0.0158)	(0.0212)	(0.0009)	(0.1432)
Equation							
3	0.0002	0.0003	0.0001	0.333	0.0458	1E-10	0.637
	(0.001)	(0.0011)	(0.0158)	(0.0228)	(0.0003)	(0.1432)	(0.1984)

Note: Numbers in parentheses are standard errors computed.

Table 3.A3 Unit Root Tests (1 = Canada, 2 = China, 3 = Iceland, 4 = ROW)

Variable	ADF_c[a]		$ADF_{c,t}$[b]	
w_1	-1.88	(5)[c]	-2.77	(5)
w_2	0.35	(8)	-1.72	(8)
w_3	-1.14	(8)	-1.90	(8)
w_4	-3.99	(2)	-4.60	(2)
log (x/P^*)	-1.9	(7)	-6.83	(1)
log p_1	-0.86	(8)	-2.58	(8)
log p_2	-2.64	(6)	-3.39	(6)
log p_3	-1.28	(2)	-3.64	(2)
log p_4	-1.39	(5)	-2.68	(5)
log $v(p_1)$	-2.25	(2)	-2.93	(1)
log $v(p_2)$	-1.04	(1)	-3.13	(1)
log $v(p_3)$	-3.61	(1)	-3.90	(1)
Critical Value (5%)	-2.87		-3.42	

[a] ADF_c test is in the form: $\Delta Y_t = \eta_0 + \eta_1 Y_{t-1} + \sum_{k=i}^{n} \theta_i \Delta Y_{t-i} + e_t$;
[b] $ADF_{c,t}$ test is in the form: $\Delta Y_t = \eta_0 + \eta_1 t + \eta_2 Y_{t-1} + \sum_{k=i}^{n} \theta_i \Delta Y_{t-i} + e_t$;
[c] The number of lagged difference terms in the test is given in bracket for each variable. The maximum order is 8.

Table 3.A4 SUR Estimates of US Import Demand System for Cod, AIDS Model, 1989-2010 Monthly Data (1 = Canada, 2 = China, 3 = Iceland, 4 = ROW)

Variable	Eq 1	Eq 2	Eq 3	Eq 4
constant	0.279***	0.212***	0.251***	0.258***
	(0.018)	(0.018)	(0.014)	(0.033)
log (x/P^*)	0.321***	-0.237***	-0.041*	-0.043*
	(0.027)	(0.027)	(0.02)	(0.018)
log p_1	-0.093*	0.046*	0.016	0.031
	(0.056)	(0.027)	(0.045)	(0.025)
log p_2	0.046*	-0.055*	0.017	-0.008
	(0.027)	(0.029)	(0.021)	(0.02)
log p_3	0.016	0.017	0.021	-0.054
	(0.045)	(0.021)	(0.05)	(0.036)
log p_4	0.031	-0.008	-0.054**	0.031
	(0.025)	(0.017)	(0.02)	(0.02)
log $v(p_1)$	-0.009	0.089***	-0.051***	-0.029***
	(0.012)	(0.012)	(0.009)	(0.008)
log $v(p_2)$	0.027***	-0.094***	0.037***	0.030***
	(0.007)	(0.007)	(0.005)	(0.005)
log $v(p_3)$	0.040***	-0.015	-0.010	-0.015*
	(0.01)	(0.01)	(0.008)	(0.007)

Note: Numbers in parentheses are asymptotic standard errors. Single asterisk (*) indicates significance at the $p < 0.10$ level; double asterisk (**) indicates significance at the $p < 0.05$ level; triple asterisk (***) indicates significance at the $p < 0.01$ level. The estimated coefficients of dummy variables are suppressed from the report.