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A DYNAMIC MONTHLY MODEL OF U.S. PORK PRODUCT MARKETS: TESTING FOR AND DISCERNING THE ROLE OF HEDGING ON PORK-RELATED FOOD COSTS

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ABSTRACT

This paper extends prior econometric research and applies cointegrated VAR (CVAR) modeling methods to estimate a monthly system of U.S. upstream/downstream pork-based food product markets, while incorporating an empirical link to a relevant pork futures market.

The study uses the CVAR to develop hypotheses that, if accepted, suggest the existence of hedging activity; tests such hypotheses that the data statistically accepts; imposes the hedging-suggesting restrictions into the estimated model; and then shows how the final cointegrated parameters illuminate the role and policy implications of hedging in the workings and dynamic interactions of the modeled pork product markets. Analysis of all cointegrated parameters suggests how hedging can act as a cushion on the effects of sharp upstream episodes of market volatility on U.S. downstream pork-related food cost patterns.

As well, results suggest that because U.S. demand for upstream product markets is a function of relative own/futures price, financial policies working through futures price are as effective policy levers as commodity-focused policies working through upstream commodity prices in influencing the modeled markets and managing related pork-based food cost patterns.

Keywords: cointegrated vector autoregression; hedging; comparative policy analysis; U.S. pork-related product prices.

JEL codes: C32, L66, G13, Q18

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INTRODUCTION: STUDY JUSTIFICATION, PURPOSE, AND REVIEW OF RELATED LITERATURE

This study extends prior research summarized below that demonstrates the cointegrated vector autoregression (cointegrated VAR) model's policy-analytic usefulness in addressing issues relevant to farm, resource/energy, and food market issues – particularly issues regarding commodity-based food costs and related patterns of food inflation. Such extensions arise from the application of cointegrated VAR modeling methods to a monthly U.S. system of pork product markets. This system extends from an upstream slaughter market to a number of downstream markets for relatively more processed or value-added pork products, and includes a linkage to a U.S. pork futures market. This is perhaps the literature's first linkage to a pork futures market for a monthly cointegrated VAR model of a U.S. pork product market system. And as demonstrated below, it is a linkage of increasing importance and interest to research, policymaking, financial market and agribusiness agents, as the trading pool for pork futures positions rapidly expands in size and in the number of non-traditional speculative traders that hold futures positions as a profit-generating asset class. In so doing, we accomplish the following goals:

- We demonstrate the important and policy-relevant dynamic influences and interactions among U.S. pork-related food product markets and the position taking of hedging agents. In turn, results illuminate hedging's effects on U.S. pork-based food costs and patterns of pork-based food inflation of such risk management activity.
- The second is a methodological goal related to the first goal's achievement. For perhaps the first time in the literature, we demonstrate how Johansen and Juselius' rich hypothesis testing capabilities afforded by the cointegrated VAR model may be harnessed to suggest hedging activity at a point in an upstream/downstream product market system – here of U.S. pork-based product markets. We then demonstrate how such statistically accepted restrictions that suggest hedging activity are imposed on the estimated model, and how the finally estimated model is interpreted to discern the role and policy implications of hedging in the workings and dynamic interactions of the modeled markets.
- And third, the series of hypothesis tests and parameter estimates from the cointegrated VAR model of a monthly upstream/downstream system of U.S. pork product markets provides cointegrating relationships with a rich set of economically interpretable estimates and market-propelling parameters of use in comparatively and empirically assessing market impacts of specific policies and market events. As a consequence, a number of comparative implications emerge for policies and events with a financial focus working through the futures market as opposed to those with a commodity focus.

Commodity futures markets generally have undergone two sorts of changes in recent years. Not only have the trading volumes of such contracts dramatically risen, but the trading pools, once primarily comprised of traditional traders who focused on price discovery and risk management, have also expanded noticeably. According to Auerlich et. al., such pool expansion has admitted increasing numbers of non-traditional speculative traders from hedging funds and other financial entities that hold commodity futures positions as a profit-generating asset class. These changes in the pools of traders have elicited increasing concerns over the effects on supply, demand, and levels of price volatility, not only in the futures markets themselves, but also on the contracts' underlying commodity markets, as well as on the markets for more processed products downstream that use such commodities as production inputs.

Changes in the CMEGroup's lean hog pork contract have generated concerns over effects on U.S. pork producers, processors, and pork-based food product markets further downstream, and ultimately on patterns of U.S. pork-related food inflation. According to the Commodity Futures Trading Commission's *Commitments of Traders* reports (CFTC, COT 2011), total open interest in the CMEGroup's pork contract rose 179% from 82,102 to 231,784 during the 2004-2010 period. And while the proportions of trader classes holding these positions have remained relatively constant during this period (at 19% for non-commercial trades and within the 52-55% range for commercial trades), numbers of both classes of trades have clearly increased. Consequently, the following seem very important and well-placed efforts: capturing a futures market linkage to a monthly model of such pork product markets; uncovering the extent to which agents in the U.S. upstream/downstream pork market complex use futures to manage risk and in which markets; assessing the comparative empirical market effects of alternative policies/events emanating from such different pork markets; and using the results to address patterns of U.S. pork-related food cost inflation.

Prior Literature

Hua demonstrated how cointegrated VAR methods can establish the existence, and illuminate the empirical nature, of global policy transmission mechanisms from macroeconomic (macro) policies to commodity markets through price. He applied cointegrated VAR modeling methods to 1970-1995 data that focused on five non-petroleum global commodity prices and three macro policy levers that included a world industrial price index (as a growth proxy), an exchange rate index of the U.S. dollar relative to major world currencies, and a global interest rate proxy. His cointegrated VAR model estimations uncovered five cointegration vectors or CVs between each of the five prices and the three macro policy levers. Analysis and interpretation of the CVs suggested strong statistical evidence of the global macro/commodity policy transmission mechanisms reflected by the CVs. Such results and interpretations further suggested that policies promoting global growth likely stimulate world commodity demand, and in turn, commodity prices. Finally, policies appreciating relative U.S. dollar values and/or reducing global interest rates tend to augment global commodity demand and prices.

Lambert and Miljkovic demonstrated the usefulness of cointegrated VAR methods to identify policies that most effectively address and manage U.S. food price inflation issues from upstream agricultural input markets to aggregate food prices. They focused on the debate of whether the 2006-2008 surge in U.S. food inflation was induced by farm price and manufacturing wage increases or by surges in energy costs and consumer incomes. Their cointgrated VAR model of U.S. food prices, farm prices, fuel prices, wages, and consumer income was estimated with 1989-2009 data, and yielded two CVs that were analyzed and interpreted. Results suggested that farm prices and manufacturing wages, rather than energy costs and consumer incomes, were the relatively more effective determinants of, and hence the better policy tools with which to manage, U.S. food price inflation.

Babula and Lund applied methods similar to those of Lambert and Miljkovic, although they focused on a more disaggregate system of U.S. pork-related input/output product markets. These pork-based markets included the upstream farm market, the wholesale market for processed pork, and the wholesale sausage market for the 1989-2006 period. A quarterly cointegrated VAR of these U.S. pork product markets yielded a single CV that emerged as a U.S. demand for pork as an input. Analysis, statistical inference, and interpretation of the cointegrated VAR model results provided a policy-relevant empirical depiction of how the pork-based markets work and dynamically interact, as well as updated empirical estimates of market-propelling parameters and price transmissions. Results demonstrated how policies and market events upstream influence related downstream pork-related food product markets, along with insights on how these policies and events address and manage downstream patterns of pork-based food costs.

Cointegrated VAR Modeling, Modeled Markets, and Data Resources

As is well known, economic time series often fail to meet conditions of weak stationarity (a.k.a. stationarity and ergodicity) required of valid inference, and in some cases unbiased estimators, when applying regression to time-ordered data (Granger and Newbold, pp. 1-5). And while often individually non-stationary, they can form vectors with stationary linear combinations, whereby the series move in tandem and behave in a stationary manner as a group known as an error-correcting cointegrated system (Johansen and Juselius).

A search of U.S. market data resources provided the following endogenous monthly variables (denoted by parenthetical labels) with adequate numbers with which to conduct this study:

- U.S. commercial pork slaughter (QSLAUGHT): millions of pounds in Federally inspected and other plants, based on packers' dressed weights, from the U.S. Department of Agriculture, Economic Research Service (USDA/ERS 2012).
- U.S. slaughter pork price (PSLAUGHT): This is the U.S. producer price index or PPI for slaughter hogs, farm products group, series number WPU0132. This was obtained from the U.S. Department of Labor, U.S. Bureau of Labor Statistics (Labor, BLS 2012).
- Pork Futures Price (PFUTURES): This is a series of the monthly averages of the daily settlement prices for the CMEGroup's lean hog contract downloaded from Bloomberg.[1] The roll methodology was chosen as the front month contract with roll

[1] See the following link accessed by an author on February 14, 2012: www.bloomberg.com. Ticker: LH1 Comdty. Further, the futures contracts used in PFUTURES have a final settlement price that is based on the information obtained from "National Daily Direct Hog Prior Day Report-Slaughtered Swine" released by the U.S. Department of Agriculture for the two-day period ending on the contract's last day of trading. The settlement price is derived from data concerning all producer-sold negotiated and swine or pork market formula barrows and gilts purchased on a lean value direct basis for which the head count, average net price and average carcass weight" are reported in the above report. See CME Rule book at the following link accessed by an author on December 22, 2011: http://www.cmegroup.com/rulebook/CME/II/150/152/152.pdf

into the next nearest contract on the first business day of the front month contract's expiration month.

- U.S. wholesale price for processed pork meat (PPROC): This is reflected by the U.S. PPI for pork (processed or cured), not canned or made into sausage, series no. PCU3116123116121 (Labor, BLS 2012).
- U.S. wholesale price of sausage (PSAUSAGE): This price is reflected by the U.S. PPI for sausage and similar products (except canned) made from purchased products, series no. PCU311616124 (Labor, BLS 2012).
- U.S. wholesale price of ham (PHAM): Given the lack of a U.S. PPI for ham with adequate observations for this analysis, PHAM is reflected by the U.S. consumer price index for ham, U.S. city average, series no. CUUR0000SEFD02 (Labor, BLS 2012).

It is important to note that the chosen CMEGroup pork futures price and its underlying contract are considered the most "national or macro-relevant" pork futures market contract representation as could be located to match the nationally-surveyed nature of the four PPI series. As well, the CMEGroup pork futures contract price was chosen for its location at a pricing point similar to the PSLAUGHT point that assigns value to U.S. slaughtered pork (QSLAUGHT), but at an average forward horizon of 70 days.[2] Further, PFUTURES (i) represents the most widely traded U.S. pork futures contract, (ii) represents a stream of futures price indications of all of the contract's delivery months throughout the year, and (iii) is considered the most widely and nationally monitored "bell-weather" U.S. futures pork price indicator based on observed levels of open interest.

Data are monthly, modeled in natural logarithms, not seasonally adjusted, and shown below to be non-stationary or integrated of order-1 [I(1)]. Following prior work, we crafted a sample within the liberalized North American trade regime that was initiated by the January, 1989 implementation of the Canadian/U.S. Free Trade Agreement, was followed by the January, 1994 implementation of the North American Free Trade Agreement or NAFTA, and resulted in this study's monthly 1989:01 – 2011:12 sample period (see Babula and Lund, p. 109).

Following Juselius and Toro and Juselius (chs. 1-4), we examined the logged levels and differences to assess the data's non-stationarity properties. Such examinations led to formulation of specification implications of these properties that utilize inherent stores of information to avoid compromised inference, and in some cases, biased estimates (Granger and Newbold). Incorporating statistically supported specification implications in turn results in a statistically adequate underlying VAR model (and algebraically unrestricted VEC) with which the cointegration properties of the six endogenous variables can be exploited.

As noted in prior research, monthly quantities for downstream processed pork, sausage, and ham markets are not available, since such data are often considered proprietary business

[2] The estimate that PFUTURES prices unprocessed product at/near the slaughter point at a time-point that is an average 70 days forward of PSLAUGHT arises from a number of factors. First, as close as could be obtained, PSLAUGHT defined above provides a current national upstream price for unprocessed slaughtered pork at/near the farm gate. And second, the 70-day average estimate uses the assumptions that (i) a monthly average price presents approximately half a month (15 days), (ii) the contract will have one or two months until the delivery month, and (iii) average settlement occurs about 15 days into the delivery month, so as to realize the following summation: 15+(30 or 60)+15. Since PFUTURES is based on futures contracts that are more often a single month out rather than two, the average 70-day period of futures being forward of PSLAUGHT emerged.

information and are excluded from the public domain (Babula and Lund). We consequently followed prior research procedure and specified the U.S. pork-related markets represented above as reduced form price equations as Goodwin, Mackenzie, and Djunaidi did for broiler product markets, and as Bessler and Akleman did for pork and beef product markets.

The Statistical Model: The Levels VAR and Unrestricted VEC Equivalent[3]

Sims (1980) and Bessler (1984) note that a VAR model posits each endogenous variable as a function of k lags of itself and of each of the system's remaining endogenous variables. The above pork-related variables render the following six-equation VAR model in logged levels:

$$\begin{aligned}
X(t) = {} & a(1,2)*QSLAUGHT(t-1) + \ldots + a(1,k)*QSLAUGHT(t-k) + \\
& a(2,1)*PSLAUGHT(t-1) + \ldots\ldots \quad + a(2,k)*PSLAUGHT(t-k) + \\
& a(3,1)*PFUTURES(t-1) + \ldots\ldots \quad + a(3,k)*PFUTURES(t-k) + \\
& a(4,1)*PPROC(t-1) + \ldots\ldots\ldots \quad + a(4,k)*PPROC(t-k) + \\
& a(5,1)*PHAM(t-1) + \ldots\ldots\ldots \quad + a(5,k)*PHAM(t-k) + \\
& a(6,1)*PSAUSAGE(t-1) \ldots\ldots\ldots \quad + a(6,k)*PSAUSAGE(t-k) + \\
& a(c)*CONSTANT \quad + a(s)*SEASONALS + \gamma(t)
\end{aligned} \tag{1}$$

In (1), X(t) = QSLAUGHT(t), PSLAUGHT(t), PFUTURES(t), PPROC(t), PHAM(t), and PSAUSAGE(t). The asterisk denotes the multiplication multiplier; the t refers to current time period-t; and $\gamma(t)$ is a vector of white noise residuals. The a-coefficients are ordinary least squares regression estimates with the first parenthetical digit denoting the six endogenous variables as ordered in X(t)'s definition, and the second reflecting the lagged value. The lag structure, k=3, was proscribed by results from the application of Tiao and Box's (1978) lag search procedure. The a(c) denotes the intercept on a constant of 1.0. Equation 1 also includes a vector of 11 centered seasonal variables and a number of other binary variables discussed below. A TREND or TRD is also included.

It is well known that a levels VAR of a lag order-k can be equivalently written more compactly as an unrestricted vector error correction (unrestricted VEC) model (Juselius, pp. 59-63; Johansen and Juselius):

$$\Delta x(t) = \Gamma(1)*\Delta x(t-1) + \ldots + \Gamma(k-1)*\Delta x(t-k+1) + \Pi*x(t-1) + \Phi D(t) + \varepsilon(t) \tag{2}$$

The endogenous variable number, p, is six. The $\varepsilon(t)$ are white noise residuals, the delta is the difference operator, while the x(t) and x(t-1) are p by 1 vectors of the endogenous variables in current and lagged levels. The $\Gamma(1), \ldots, \Gamma(k-1)$ terms are p by p matrices of short run regression coefficients, and Π us a p by p long run error correction term to account for endogenous levels.

The $\Phi D(t)$ is a set of deterministic variables, including an array of binary (dummy) variables that will be added to address stationarity and specification issues as the analysis unfolds below. The error correction (EC) term is decomposed as follows:

[3] This section draws heavily on the seminal articles by Johansen and Juselius and the book by Juselius.

$$\Pi = \alpha^*\beta' \quad (3)$$

The α is a p by r matrix of adjustment coefficients (r is the number of cointegrating relationships or the rank of Π discussed below). The β is a p by r vector of cointegrating parameters.

The error correction or EC term retains the levels-based and other long run information: linear combinations of non-differenced and individually I(1) levels variables (under cointegration); permanent shift binaries to capture more enduring effects of policy/market events (presented below); and a linear trend. The term [$\Gamma(1)^*\Delta x(t-1) \ldots \Gamma(k-1)\Delta x(t-k+1)$, $\Phi D(t)$] collectively comprises the model's short run/deterministic component (hereafter denoted short run component) that includes the permanent shift binaries in differenced form, observation-specific outlier binaries (introduced below), and seasonal binaries.

We followed prior related research (see Babula and Lund) and initially restricted non-differenced permanent shift binary variables to the levels-based error-correction space to account for implementations of the North American Free Trade Agreement in 1994 and the Uruguay Round Agreement in 1995 (denoted NAFTA and URUGUAY). Two additional permanent shift binaries were also considered: one to account for the recession that began in December, 2007 and terminated in mid-2009 (RECESS) and another to account for the extraordinary fall in U.S. pork slaughter price in mid-1997 through late 1998 (DIST9798) that is discussed later.

We followed Juselius' (ch. 6) method of identifying and including extraordinarily influential effects of month-specific 'outlier" events through specification of "outlier" binaries. When a potentially includable outlier was identified with a "large" standardized residual, an appropriately specific variable was included in differenced form as part of equation 2's short run component, and retained if a battery of diagnostics (discussed below) moved favorably to suggest enhanced specification.[4]

Table 1's battery of diagnostic values for the levels VAR (and its unrestricted VEC algebraic equivalent) before and after efforts focused on enhanced specification suggest clear benefits from such efforts. The trace correlation, a goodness of fit indicator, increased 92% to 0.655. While serial correlation was initially an issue, the finally estimated levels VAR after specification efforts generated evidence that serial correlation was no longer an issue.

Doornik-Hansen (D-H) values test the null that the estimated model's residuals behave normally. The D-H values for the estimated system greatly improved after specification such

[4] We followed Juselius' (ch. 6) analysis to identify outlier binaries in equation 2's short run component. An observation-specific event was judged as potentially "extraordinary" if its standardized residual exceeded 3.0 in absolute value. Such a rule for outliers was designed based on the effective sample size of 276 observations using the Bonferoni criterion: INVNORMAL $(1\text{-}1.025)^T$ where T=276 and INVNORMAL is a function for the normal distribution that returns the variable for the c-density function of a standard normal distribution (Estima). The Bonferoni variate had an absolute value of 3.7. Having realized that there were some month-specific events with potentially extraordinary effects with absolute standardized residual values of about 3.0, we opted to follow recent research and chose a more conservative Bonferoni absolute value criterion of 3.0 rather than 3.7 (Babula and Lund). Observations with absolute standardized residual values of 3.0 or more were thereby considered as potential outliers, and we specified an appropriately defined variable for relevant observations for the sequential estimate procedure. Sixteen binaries were ultimately included. Due to space limitation considerations, we do not report the binaries as they are part of the estimated model's short run component that is not a focus of this study on long run cointegration relationships. The binaries are available from the authors on request.

that the initially non-normally behaving system of estimated residuals ultimately achieved strongly normal behavior.

Table 1. Mis-specification Tests for the Unrestricted VEC: Before and After Specification Efforts

Test and/or equation	Null hypothesis and/or test explanation	Prior efforts at specification adequacy	After efforts at specification adequacy
Trace correlation	system-wide goodness of fit: large proportion desirable	0.342	0.655
LM Test for serial correlation (lag 2)	Ho: no serial correlation by lag-2. Reject for p-values of 0.5 or less	116.3 (p=0.000000, see note)	47.2 (p=0.1004)
Doornik-Hansen test, system-wide normality	Ho: modeled system behaves normally. Reject for p-values below 0.05.	334.3 (p=0.000000, see note)	11.91 (p=0.51)
Doornik-Hansen test for normal residuals (univariate)	Ho: equation residuals are normal. Reject for values above 9.2 critical value		
ΔQSLAUGHT		1.80	2.51
ΔPSLAUGHT		152.02	0.45
ΔPFUTURES		21.05	1.56
ΔPPROC		37.41	.045
ΔPHAM		30.37	1.06
ΔPSAUSAGE		34.02	2.65
Skewness (kurtosis) univariate values	Skewness: ideal is zero; "small" absolute value acceptable kurtosis: ideal is 3.0; acceptable range is 3.0-5.0.		
ΔQSLAUGHT		-0.191 (3.09)	-0.218 (2.90)
ΔPSLAUGHT		-1.62 (15.8)	.036 (3.10)
ΔPFUTURES		-.042 (4.63)	0.036 (3.1)
ΔPPROC		0.45 (5.44)	0.087 (3.1)
ΔPHAM		-0.77 (3.68)	-0.147 (2.90)
ΔPSAUSAGE		0.48 (5.33)	0.231 (2.90)

Note.—The p-value "p=0.000000" implies that the value was adequately small so as not to register at the 6^{th} decimal place.

The univariate D-H values suggest noticeable improvement from a setting before specification efforts where five of the six estimated equations generated non-normally behaving residuals to a point after such efforts when all six equations generated evidence at the 5% level that was clearly insufficient to reject the hypothesis of normally behaving estimated residuals. Finally, Table 1 suggests that the finally estimated and statistically statistically adequate model displayed skewness and kurtosis indicators that fell within literature-accepted ranges.

Cointegration: Testing For and Imposing an Appropriate Reduced Rank

The endogenous variables are shown below to be I(1). Juselius (p. 80) notes that cointegrated variables are driven by common trends, and stationary linear combinations called cointegrating vectors or CVs (also known as cointegrating relations). The Π-matrix in equation 3 is a p by p (here 6 by 6) matrix equal to the product of two p by r matrices: β of error correction estimates that under cointegration combine into $r < p$ stationary CVs of the six individually non-stationary pork-related prices, and α of adjustment coefficients (betas and alphas respectively). Under cointegration, the rank of β'x(t) is reduced despite the non-stationarity of x(t)'s six series.

Table 2 provides nested trace test results and evidence for rank determination. Evidence at the 5% significance level is sufficient to reject the first three null hypotheses, suggesting that r is not 2 or less.

Table 2. Nested Trace Test Statistics and Reduced Rank Determination

Null Hypothesis	Trace Value	95% Fractile	Result
Rank or $r \leq 0$	173.05	124.65	Reject null that $r \leq 0$.
Rank or $r \leq 1$	123.29	95.76	Reject null that $r \leq 1$.
Rank or $r \leq 2$	78.48	70.86	Reject null that $r \leq 2$.
Rank or $r \leq 3$	47.35	49.97	Fail to reject that $r \leq 3$.
Rank or $r \leq 4$	26.20	32.93	Fail to reject null that $r \leq 4$.
Rank or $r \leq 5$	7.02	19.65	Fail to reject null that $r \leq 5$.

Notes. -- As recommended by Juselius (2006), CATS2–generated fractiles are increased by 4*1.8 or 7.2 to account for the four permanent shift binary variables restricted to lie within the cointegration space. As recommended by Juselius (ch. 8) and programmed by Dennis, trace values are corrected with Bartlett's small sample adjustment.

Evidence at the 5% level is insufficient to reject the fourth null that r is 3 or less, and given the nested nature of these tests, this suggests that reduced rank is three. This and other evidence thereby suggests that three cointegrating relations error-correct the modeled system.[5]

[5] Juselius (ch. 8) , Juselius and Toro, and Juselius and Franchi strongly recommend against sole reliance on trace test evidence in determining the EC space's reduced rank, here r=3. We followed these recommendations and consulted two other sources of evidence: patterns of characteristic roots in the companion matrices generated under relevant assumptions of r, as well as analysis of the significance patterns of the estimated adjustment or α-coefficients for the potentially included CVs. If r=3 is an appropriate choice, then one expects p-r or 3 unit roots in the companion matrix under r=3 with the (p-r+1)st or fourth root being substantially sub-unity. Under r=3, the companion matrix's first three characteristic roots were indeed unity, with the fourth valued at 0.87, a value notably below 1.0, so as to support the choice of r=3. We do not report the results of the characteristic unit root patterns under other r-assumptions, although all generally and collectively support the choice of three for the EC space's reduced rank. These other results are available from the authors on request. The second source of evidence considered was an analysis of the statistical significance patterns of the α-estimates. A CV that actively participates in, and that should be included in, the VEC model's EC mechanism should display statistically significant adjustment coefficients, a reflection that the CV is endogenously participating and adjusting to the error correction process (Juselius, pp. 137-144). The estimated unrestricted VEC model's first three CV's indeed contain statistically significant adjustment coefficient estimates, as demonstrated by the estimate's pseudo t-value having an absolute value of 2.6 or more at the five percent level (Juselius, p. 142). More specifically, the first CVs generated the statistically significant α-estimates as follows: on ΔQSLAUGHT (t = -4.2), ΔPSLAUGHT (t = 3.3), ΔPSAUSAGE (t = 2.8), and ΔPHAM (t = -5.2) in CV1; on

Hypothesis Tests on the Three Unrestricted Cointegrating Relations

One begins with the three unrestricted cointegrating relationships (not reported due to space considerations) that emerged from Johansen and Juselius' reduced rank estimation of equation 2 after having imposed a rank of r =3 on the EC space. One conducts a sequential series of hypothesis tests (detailed below) on the EC space, and then re-estimates the system with the statistically-supported restrictions imposed. Hypothesis tests on the beta take the form:

$$\beta = H^*\varphi \tag{4}$$

Above, β is p1 by r vector of coefficients included in the cointegration space,[6] and H is a p1 by s design matrix, with s being the number of unrestricted or free beta coefficients. The φ is an s by r matrix of unrestricted beta coefficients. The hypothesis test value or statistic is:

$$2\ln(Q) = T^*\sum [(1-\lambda_i^*) / (1-\lambda_i)] \text{ for } i= 1,2,3 \ (=r) \tag{5}$$

Asterisked (non-asterisked) eigenvalues (λ_i = 1, 2, 3) are generated with (without) the tested restrictions imposed. Following Juselius (chs. 10-12), two groups of hypotheses are tested. The first group consists of six system-based and rank-dependent stationarity (unit root) tests on the six endogenous variables. Juselius, Juselius and Toro, and Juselius and Franchi recommend this approach over univariate unit root tests (e.g. Dickey-Fuller and Phillips-Perron tests) for such multivariate models as ours. Six system-based and rank-dependent likelihood ratio tests programmed in CATS2 by Dennis were conducted to test the null hypotheses that each of the endogenous variables is stationary. Evidence suggested that all six endogenous variables are non-stationary or I(1) in logged levels.[7]

The second group of tested hypotheses are those that emerge from perusing the three rank-restricted CVs' parameter estimates. Juselius (chs. 10-12) recommends the following four-step hypothesis testing/re-estimation procedure, which permits one to incorporate economic and econometric theory, market knowledge/expertise, and prior research results into the cointegration space through the imposition of statistically supported parameter restrictions:

- After perusing the rank-restricted CV estimates, one formulates an economically/econometrically viable hypothesis,

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ΔQSLAUGHT (t = 4.03), ΔPFUTURES (t = -3.1), ΔPSLAUGHT (t = -3.4), and ΔPHAM (t = -4.8) in CV2; and on ΔPFURURES (t = 3.8) in CV3.

[6] The p1 equals 11: it is the sum of p=6 endogenous variables plus the five deterministic variables restricted to the cointegration space (four permanent shift binaries and a trend).

[7] More specifically, equation 4 is re-written as) $\beta^c = [b,\varphi]$. Let p1 be the new dimension of 11 to reflect the six endogenous variables and the five deterministic variables restricted to lie in the EC space. The β^c is a p1 by r or 11 by 3 matrix with one of the variable's levels restricted to a unit vector; b is a p1 by 1 or 11 by 1 vector with a unity value corresponding to the variable the stationarity of which is being tested; and φ is a p1 by r-1 or 11 by 2 matrix of the remaining unrestricted CV parameter estimates. Given the rank of 3, then the test values and parenthetical p-values for the six stationarity tests are as follows with the null of stationarity rejected for p values below 0.05: 11.51 (0.0093) for QSLAUGHT; 15.6 (0.0014) for PSLAUGHT; 11.40 (0.0097) for PFUTURES; 11.34 (0.010) for PPROC; 21.30 (0.00010) for PSAUSAGE; and 9.20 (0.027) for PHAM.

- The hypothesis is tested using equations (4) and (5),
- If statistically supported, the hypothesis is incorporated into the EC space through imposition of restrictions,
- With the supported restrictions imposed, the EC space is re-estimated using Johansen and Juselius' reduced rank estimator, and the newly emergent CV parameter estimates are re-examined for additional and viable testable hypotheses.

This process should be iteratively repeated until the evidence accepts the cumulative set of statistically accepted restrictions; offers no substantial further hypotheses to test; and renders a set of finally restricted CVs accepted by the data and that are ultimately reported below. In so doing, it is possible to restrict the cointegrating relations so as to sometimes attribute structural properties and interpretations to the CVs, as is the case with our three finally restricted CVs (Juselius). Table 1 summarizes the seven sets of hypothesis tests that are discussed below.

As is well known, the set of three CVs that emerge after imposition of the reduced rank of r = 3 do not meet the rank condition of identification. In Table 3's first set of tested conditions, TS-1, we began with hypotheses based on theory, market knowledge, and prior research, and that rendered the CVs compliant with the rank condition of identification ("identifying conditions"). These hypotheses included the following restriction that, if statistically accepted, is shown below to suggest the existence of hedging activity: -β(PFUTURES) = β(PSLAUGHT). This restriction was chosen for testing because in the CV2 that emerged from the estimation after imposition of the reduced rank of 3 on the EC space, the β-estimates on futures and slaughter prices were oppositely signed and of nearly equal absolute values.

This suggests that QSLAUGHT (CV2's normalizing variable) is dependent on both prices rather than on slaughter price alone. This test and its interpretation are explained and rationalized below. TS-1 restrictions were tested using equations (4) and (5); were then accepted by the data; and were imposed on the CVs that were then re-estimated using Johansen and Juselius' reduced-rank estimator.

At this point, the CVs were identified and CV2 began to emerge as a processor demand for U.S. slaughter pork, with QSLAUGHT and PSLAUGHT negatively related. In turn, a new set of estimates then yielded further testable hypotheses that were added to TS-1 to render test set 2 or TS-2 in Table 3.

This procedure [bulleted items above] was iteratively implemented five more times, with accepted restrictions collectively imposed. Space limitations precluded the reporting of estimation results after each of the seven sets of hypothesis tests. Rather, Table 3 reports seven sets of hypothesis tests/test results and imposed restrictions that evolved and were successively accepted. The collective set of restrictions that were statistically accepted across the reduced-rank estimations comprise TS-7. The testing of TS-7 generated a chi-square test value of 17.95 that, with a p-value of 0.16 in excess of 0.05, suggests that the data clearly accepted the TS-7 restrictions at the 5% significance level.

Imposition of TS-7's collective restrictions and the final (seventh) reduced rank estimation rendered the three finally-restricted cointegrated relations that constitute the model's three cointegrating relations or CVs and that are presented below as equations 6, 7, and 8.

Table 3. Sequential Hypothesis Tests on Beta Cointegration Parameters of U.S. Pork-Based Markets

Tested restrictions, restriction no.	Explanation, reasoning	Test values, test results, interpretations, analysis of estimates. Fail to reject (i.e. accept) restrictions for p>0.01 (or 0.05)
Test Set-1, TS-1: Identifying restrictions to comply with rank conditionand Test for Hedging		
4 on CV1: β(PSAUSAGE) = β(PHAM) = β(QSLAUGHT) = 0 β(NAFTA) = 0 4 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = 0 -β(PFUTURES) = β(PSLAUGHT) 2 on CV3: β(QPORK) = β(PFUTURES) = 0	Identifying restrictions. Near-zero value in rank-restricted CV1. Identifying restrictions. Identifying restriction constituting an effective hedging test. Identifying restrictions.	Chi-sq. (df=4) = 9.41, p = 0.052. Results: Evidence strongly accepts TS-1 restrictions at 5% level as p exceeds 0.05. Analysis/new restrictions: The following t-values are insignificant: -0.36 on β(RECESS) and -0.64 on β(DIST9798) in CV1. Add zero restrictions on these coefficients in TS-2.
Test set-2, TS-2: TS-1 restrictions plus β (RECESS)) = β(DIST9798) = 0 on CV1		
6 on CV1: β(PSAUSAGE) = β(PHAM) = β(QSLAUGHT) = 0 β(NAFTA) = 0 β(RECESS) = β(DIST9798) = 0 4 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = 0 β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) 2 on CV3: β(QPORK) = β(PFUTURES) = 0	See TS-1. Insig. β t-values, analysis of TS-1 results. See TS-1. See TS-1.	Chi-sq. (df=6) = 9.67, p=0.14. Results: Evidence accepts TS-2 restrictions as p far exceeds 0.05. Analysis/new restrictions: Following t-values are insignificant: 0.87 on β(URUGUAY) in CV1 and 0.67 on β(DIST9798) in CV3. Add zero restriction on this coefficient in TS-3.
Test set 3, TS-3: TS-2 restrictions plus β(URUGUAY) = 0 in CV1 and β(DIST9798)= 0 in CV3.		
7 on CV1: β(PSAUSAGE) = β(PHAM) = β(PCHOPS) = 0 β(NAFTA) = β(RECESS) = β(DIST9798) = 0 β(URUGUAY) = 0 4 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = 0 β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) 3 on CV3: β(QPORK) = β(PFUTURES) = 0 β(DIST9798) = 0	See TS-1 and TS-2. Insig. β t-values, analysis of TS-2 results See TS-1 and TS-2. See TS-1 and TS-2. Insig. β t-value, analysis of TS-2 results.	Chi-sq. (df=8) = 10.62, p=0.22. Results: Evidence strongly accepts TS-3 restrictions as p far exceeds 0.05. Analysis/new restrictions: T-value of -1.1 on 3.0 on β(SAUSUAGE) is insignificant in CV3. Add zero restriction on this coefficient in TS-4.

Tested restrictions, restriction no.	Explanation, reasoning	Test values, test results, interpretations, analysis of estimates. Fail to reject (i.e. accept) restrictions for p>0.01 (or 0.05)
Test Set-4, TS-4: TS-3 restrictions plus β(PSAUSAGE)=0 in CV3.		
7 on CV1: β(PSAUSAGE) = β(PHAM) = β(PCHOPS) = 0 β(NAFTA) = β(RECESS) = β(DIST9798) = 0 β(URUGUAY) = 0 4 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = 0 β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) 4 on CV3: β(QPORK) = β(PFUTURES) = 0 β(DIST9798) = 0 β(PSAUSAGE) = 0	See TS-1 through TS-3. See TS-1 through TS-3. See TS-1 through TS-3. Insig. β t-value, analysis of TS-3 results.	Chi-sq. (df=9) = 11.28, p=0.26. Results: Evidence strongly accepts all cumulative restrictions comprising TS-4, as p far exceeds 0.05. Analysis/new restrictions: T-values of -1.7 on β(URUGUAY) and of 1.7 on β(NAFTA) in CV3 are insignificant. Add these as zero restrictions in TS-5.
Test Set-5, TS-5: TS-4 restrictions plus β(URUGUAY) = β(NAFTA)=0 in CV3.		
7 on CV1: β(PSAUSAGE) = β(PHAM) = β(PCHOPS) = 0 β(NAFTA) = β(RECESS) = β(DIST9798) = 0 β(URUGUAY) = 0 4 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = 0 β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) 6 on CV3: β(QPORK) = β(PFUTURES) = 0 β(DIST9798) = 0 β(PSAUSAGE) = 0 β(URUGUAY) = β(NAFTA) = 0	See TS-1 through TS-4. See TS-1 through TS-4. See TS-1 through TS-4. Insig. β t-values, analysis of TS-4 results.	Chi-sq. (df=11) = 13.73, p= 0.248. Results: Evidence strongly accepts all cumulative restrictions comprising TS-5, as p far exceeds 0.05. Analysis/new restrictions: T-value -1.45 on β(NAFTA) in CV2 is insignificant. Add this as a zero restriction in TS-6.
Test Set-6, TS-6: TS-5 restrictions plus β(NAFTA) = 0 in CV2.		
7 on CV1: β(PSAUSAGE) = β(PHAM) = β(PCHOPS) = β(NAFTA) = β(RECESS) = β(DIST9798) = 0 β(URUGUAY) = 0 5 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) β(NAFTA) =0. 6 on CV3: β(QPORK) = β(PFUTURES) = 0 β(DIST9798) = β(PSAUSAGE) = 0 β(URUGUAY) = β(NAFTA) = 0	See TS-1 through TS-5. See TS-1 through TS-5; Insig. β t-value, analysis of TS-5 results. See TS-1 through TS-5.	Chi-sq. (df=12) = 15.27, p= 0.227. Results: Evidence strongly accepts all cumulative restrictions comprising TS-6, as p far exceeds 0.05. Analysis/new restrictions: T-value -1.995 on β(DIST9798) in CV2 is insignificant. Add this as a zero restriction in TS-7.

Table 3. (Continued)

Test Set-7, TS-7: TS-6 Restrictions plus plus β(DIST9798) = 0 in CV2.		
7 on CV1: β(PSAUSAGE) = β(PHAM) = β(PCHOPS) = 0 β(NAFTA) = β(RECESS) = β(DIST9798) = 0 β(URUGUAY) = 0 6 on CV2: β(PSAUSAGE) = β(PHAM) = = β(PPROC) = β(NAFTA) = 0 -β(PFUTURES) = β(PSLAUGHT) β(NAFTA) =0. β(DIST9798) = 0 6 on CV3: β(QPORK) = β(PFUTURES) = 0 β(DIST9798) = 0 β(PSAUSAGE) = 0 β(URUGUAY) = β(NAFTA) = 0	See TS-1 through TS-6. See TS-1 through TS-6. Insig. β t-value, analysis of TS-5 results. See TS-1 through TS-6.	Chi-sq. (df=13) = 17.95, p= 0.0.16. Results: Evidence strongly accepts all cumulative restrictions comprising TS-7, as p far exceeds 0.05. Analysis/new restrictions: Analysis is complete. No more restrictions need to be added. See equations 6-8 for finally-restricted cointegration parameter estimates.

As demonstrated below, these relations provide some structural economic content, and they may be economically interpreted for policy implications, particularly equation 7 or CV2.

(6) PSLAUGHT =	1.02*PFUTURES (13.54)	+0.743*PPROC (3.72)	-0.001*TREND (-4.05)		
(7) QSLAUGHT =	-0.42(PSLAUGHT – PFUTURES) (11.50)		-0.02*URUGUAY (-2.22)	+0.03*RECESS (+2.47)	+0.001*TRD (14.92)
(8) PHAM =	+0.01*PSLAUGHT (4.91)	+0.12*PPROC (1.81)	+0.031*RECESS (2.70)	+0.001*TRD (9.79)	

Juselius (p. 142) notes the frequently-encountered difficulty in attributing structural economic interpretations to emergent cointegrating relations:

> "It is . . . important to note that the cointegration rank is not in general equivalent to the number of theoretical equilibrium relations derived from an economic model . . . Thus cointegration between variables is a statistical property of the data that only exceptionally can be given a direct interpretation as an economic . . . relation. The reason for this is that a theoretically meaningful relation can be (and often is) a linear combination of several 'irreducible cointegrated relations."

We first focus on and emphasize equation 7 or CV2 because it emerged from the finally-restricted cointegration space as one of the theoretical relations noted by Juselius above. More specifically, equation 7 is shown to have emerged as a U.S. processor demand for slaughtered pork as a productive input.

Meanwhile, CV1 or equation 6 and CV3 or equation 8 emerged as reduced form product price transmission mechanisms. These will likely remain irreducible relations and without full structural interpretations until the complete downstream information sets that are currently devoid of downstream quantities shall become as complete as the slaughter market's information set. Yet CV1 and CV3 also have policy interpretations and implications, as noted below for research, policymaking, and agribusiness agents.

Processors' Input Demand for Slaughtered Pork and Indications of Hedging

Equation 7 emerges as a monthly U.S. processor demand for slaughtered pork as an input that displays similarities with the quarterly demand relation estimated by Babula and Lund (p. 128), although this latter demand did not incorporate a linkage to the futures market. Recall that after estimation with an imposed reduced rank of 3, CV2 began to emerge as a demand with a negative relationship between slaughter quantity (its normalizing variable) and its price, PSLAUGHT. As well, this CV2 contained β-estimates for slaughter and futures prices that were equal in absolute value but oppositely signed. This suggests that the two prices may have oppositely-signed effects of similar absolute magnitudes on CV2's normalizing variable, QSLAUGHT, and can be interpreted as an initial signal that hedging activity may exist at the U.S. pork slaughter point. Consequently, we added the testable hypothesis that β(PSLAUGHT) = -β(PFUTURES) in CV2 as one of TS-1's identifying restrictions. As explained below, this tests for indications of hedging activity at the modeled system's slaughter pricing point. We further demonstrate that the restriction was accepted by the data and incorporated into the emergent processor demand for slaughtered pork. We then interpret the finally-restricted CV2 cointegrated parameters and demonstrate how they rationalize the observed market adjustments during the extraordinary episode of U.S. pork product market volatility that occurred in late-1998 and early-1999. We demonstrate how the then-current and futures prices interacted and influenced U.S. processor demand for slaughter pork.

Hedging's Role in Processor Demand for Slaughtered Pork

As expected for a demand, equation 7's processor demand for slaughtered pork is negatively related to slaughter price. And perhaps more interestingly, such demand is positively and equally dependent on futures price. That is, QSLAUGHT depends as much on PFUTURES that prices slaughtered product an average 70 days forward as on its own-price of PSLAUGHT. Insofar as data were modeled in natural logarithms, equation 7's slaughter/futures price difference term implies that processor demand for slaughtered pork depends on the relative slaughter/futures price. This implication is reinforced by the notable statistical strength of the slaughter/futures price pseudo-t values (11.5) that far exceed Juselius' (p. 142) noted critical values at the 5% significance level (+2.6 or -2.6).

At first glance, equation 7's slaughter/futures price terms may suggest that concurrently equal movements could or would be mutually offsetting with no effect on processor demand for slaughtered pork. However, such a precise offset is unlikely, since the modeled slaughter and futures prices do not define (and are not intended to define) the CMEGroup pork futures contract's underlying basis. Rather, PSLAUGHT is a nationally-surveyed PPI for slaughtered pork intended to capture national food price patterns at the U.S. pork slaughter point, and is not the contract-specific cash price used to settle the CMEGroup's pork futures contract that in turn generates PFUTURES. So while the two prices are expected to qualitatively move in tandem, there is no hard expectation that a related event or policy shock should generate equal percent changes in the two prices.

Risk-managing hedging activity is implied by equation 7's slaughter/futures price term. As slaughter price increases relative to futures price, processor demand for slaughtered pork at the current pricing point, PSLAUGHT, is relatively more expensive than at the futures pricing point, PFUTURES, at an average horizon of 70 days forward. As PSLAUGHT/PFUTURES rises, there is a willingness of some processors to postpone current demand and shift some of the total demand towards the relatively cheaper futures pricing point an average of 70 days forward. This postponement of demand at the current price cushions or softens the PSLAUGHT-induced drop in current demand with a countervailing rise in future demand at PFUTURES as agents hedge by taking CMEGroup contract positions.

That is, the negative effect on processor demand from the PSLAUGHT increase appears cushioned by an increase in demand at the now relatively cheaper futures pricing point (PFUTURES) as hedgers "re-shuffle" overall demand across the two alternative price horizons.

Likewise, as slaughter price declines relative to futures price, processor demand for slaughtered pork at the current price, PSLAUGHT, is relatively cheaper than at the futures pricing point 70 days forward. As PSLAUGHT/PFUTURES falls, there is a PSLAUGHT-induced increase in current demand that is cushioned or offset by a decline in demand at the futures price 70 days forward, as agents hedge through taking CMEGroup futures positions. That is, the positive effect on processor demand from the PSLAUGHT decline appears cushioned by a decline in demand at the now relatively more costly futures pricing point as hedgers re-shuffle overall demand across the two alternative price horizons.

It is important to note that events that shock PFUTURES also have similarly reasoned effects on processor demand through changes in the relative slaughter/futures price. Reasoning on the demand effects from changes in the relative price would be motivated by changes in PFUTURES as just provided for changes in PSLAUGHT above.

Verification of Equation 7's Suggested Hedging: The Late-98/Early-99 Episode

During the latter half of 1998, U.S. pork prices plummeted 76.5% at the slaughter point and 48.2% at the futures point. The resulting decline in the relative slaughter/futures price rendered processor demand for slaughtered pork relatively cheaper at the current pricing point (PSLAUGHT) than at the futures pricing point some 70 days forward (PFUTURES). In turn, agents transferred or shifted some of the more expensive slaughtered pork demand from the futures point to the current pricing point through hedging agents' futures transactions with CMEGroup's pork futures contract.

Such hedging activity elicited by the drop in the relative slaughter/futures price is clear from Figure 1 that displays a coincidence in a run-up in the CMEGroup contract's daily trading volume (see left vertical axis and smooth blue line) with a trough in the monthly relative slaughter/futures price (see right vertical axis and more discreet red line) during late-1998 – a six month period when the contract's monthly trading volume rose noticeably by 68% from 76,671 trades during June, 1998 to 128,902 trades in December, 1998. In effect, the hedging activity acted to partially offset overall increasing impacts on processor demand at the current slaughter point by a decline in demand from the futures point some 70 days forward.

Although beyond the purview of our model, one can conjecture that hedgings' re-positioning of the overall processor demand for slaughtered pork among the alternative futures to the current pricing points may at least partially explain why the extraordinary late-1998 PSLAUGHT decrease of 76.5% ultimately resulted in a far-milder and delayed drop of only 8.4% during the seven months after July, 1998 in processed pork price or PPROC. PPROC is the price of pork at the next modeled downstream pork product pricing point.

Likewise, these events reversed about as rapidly during early-1999. During the first five months of 1999, pork prices soared from December, 1998 lows by 284% at the slaughter point and by 79% at the futures point. The resulting rise in the relative slaughter/futures price rendered processor demand for slaughtered pork relatively more expensive at the current pricing point (PSLAUGHT) than at the futures pricing point timed at an average forward horizon of 70 days (PFUTURES).

In turn, agents transferred or shifted some slaughtered pork demand from the more expensive current point to the cheaper futures pricing point through hedging agents' position taking with CMEGroup's pork futures contract. Such hedging activity from the rise in relative slaughter/futures price is evident from: (i) Figure 1 whereby contract trading volume (blue line) remained at sustained high levels during early-1999 when relative price rapidly recovered (red line), and (ii) monthly contract trade volume data that reflected that volume was still registering 111,021 trades during May, 1999, a level 45% above the June, 1998 level when this episode of overall market volatility began.

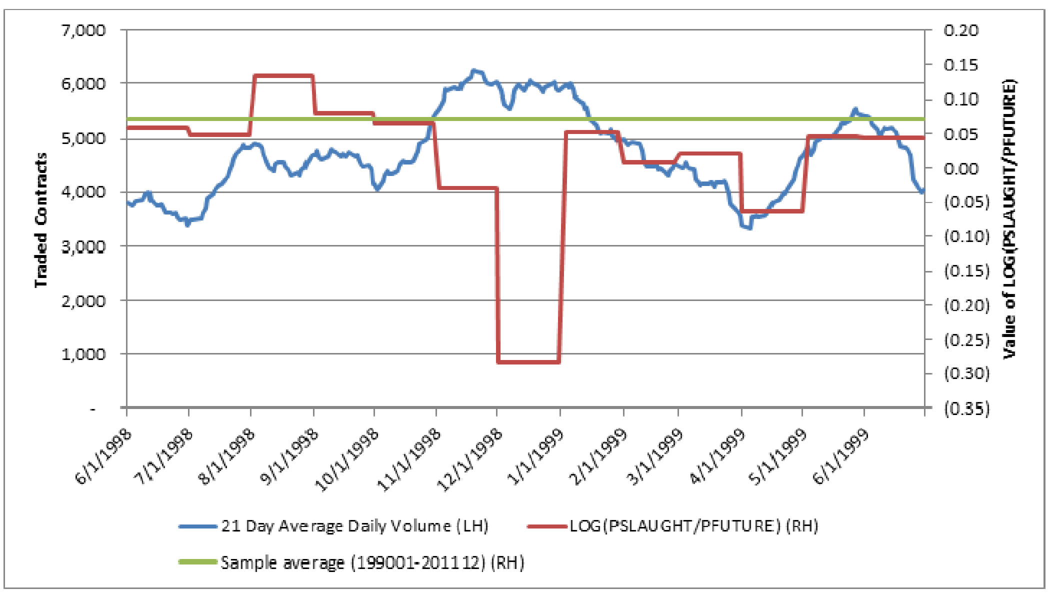

Figure 1. Relationship Between Futures Trading Volumes (LH) and Relative Slaughter/Futures Price (RH).

In effect, the hedging activity apparently acted to cushion PSLAUGHT-induced decreases in processors' demand for slaughtered pork with increases in the relatively less expensive demand at the futures pricing point 70 days forward. Although beyond the purview of our model, this hedging-induced re-positioning of some of the overall processor demand for slaughtered pork from the current to the futures pricing point may explain, at least in part, why an early-1999 PSLAUGHT increase of 284% ultimately resulted in a far-milder and delayed increase of only 12.5% in processor pork price or PPROC during the first 11 months of 1999. Recall that PPROC is the next modeled downstream pork product pricing point.

Therefore, CV2 or equation 7 illuminates an important cointegrated VAR model capability: the ability to devise testable hypotheses that, if accepted, suggest the existing of hedging and risk management activity. Hedging's existence initially emerges as a hypothesis for which one could immediately test after the cointegrated model's re-estimation under an imposed reduced rank (here r=3) in a demand relationship if the commodity's own and futures price in an unidentified CV (normalized on quantity) have estimates that are oppositely signed and of similar absolute values. One then uses equations 4 and 5 to test the restriction as noted above in CV2 with PSLAUGHT and PFUTURES. Rejection of the restriction would suggest little or no evidence of hedging, while acceptance of the restriction would, as in our analysis, suggest evidence of such hedging activity.

Two Price Transmission Mechanisms for Pork-Based Products

Equation 6 (CV1) suggests that futures and slaughter prices closely move in positive long run tandem, a result consistent with the evidence of hedging activity that emerged through processors' demand for slaughtered pork as an input. Further, there appears a close, though less than proportional, relationship among slaughter and processed pork prices, as expected among prices that differ in the levels of congealed value-added. Equation 8 is a reduced form price transmission relationship that suggests that upstream and downstream prices of pork products do tend to move positively, although far less than proportionally. We noted the limitations of these two "irreducible" relationships above due to the limitations of downstream market information sets. Nonethless, such qualitative co-movements of upstream/downstream U.S. pork food costs are of interest to agents focusing on pork food cost pass-through relationships and/or to agribusiness agents who market U.S. pork products.

Policy Implications of Equations (6), (7), and (8)

Among the many sets of policies and market events (policies/events) of relevance to the econometric results in equations 6, 7, and 8, we first and primarily focus on two sets that work through slaughter and futures prices. Thereafter, we demonstrate an important cointegrated VAR model capability that permits researchers to elicit policy effects through interpretation of binary variable coefficient estimates, while focusing on the URUGUAY binary in CV2.

Policies Working through Commodity and Futures Prices

The first is a set of "commodity-focused" policies/events that work primarily through the modeled price linkage at/near the point of pork production. For our model and in light of the data availability issues detailed above, particularly regarding pork product quantity variables, this is the slaughter point of pork price formation. In addition to the above-noted market events that elicited the extraordinary 1998-1999 episodes of U.S. pork product market volatility, commodity-focused policies could include imposing or lifting trade remedies. An example would be an imputed PSLAUGHT effect from the 1985-1999 series of U.S. firm-specific and annually-varying countervailing duties imposed on selected imports of Canadian pork (see Benson et. al.; Duffy; and WTO Panel 1991).

"Financially-focused" policies/events (for lack of a better term) are taken as those that work primarily through the pork futures price, and in turn the relative slaughter/futures price. These may include a range of Federal Reserve monetary policy actions that impact PFUTURES through changes in the CMEGroup pork futures contract's cost of carry and position-taking transaction costs. Other futures price-influencing policies/events conceivably include any of a number of actions potentially taken by the CMEGroup and/or mandated by Dodd-Frank Financial Reform Act that directly or indirectly elicit an imputed futures price effect. Examples include changes in margin requirements, new spot-month speculative position limits and accountability levels, among other changes in the contract's terms and conditions.

The three CVs demonstrate how policies and events with a commodity focus working through slaughter price and with a financial focus working through futures price comparatively influence pork product demand and related pork product prices. And in turn, these CVs, particularly equation 7, demonstrate how such alternative policies/events are potential levers useful in illuminating, addressing, and managing the upstream/downstream patterns of U.S. pork-related food cost inflation when they arise. The results do this in a number of ways.

First, equation (7) clearly suggests that commodity-focused and financially-focused policies/events work through slaughter and futures prices, respectively, and both influence U.S. processor demand for slaughtered pork with similar directness and effectiveness through impacts on the relative slaughter/futures price. In other words, an X% futures price change elicited by a financial event such as a change in the Fed's monetary policy or a Dodd/Frank-mandated change in the CMEGroup's pork contract terms and conditions would as effectively influence U.S. pork product markets as an X% change in slaughter price from a commodity-focused event such as a reduction in the level of U.S. protection from imports arising from a change in the level of assessed foreign dumping or subsidy margins associated with dumping or countervailing duty orders on PSLAUGHT-relevant imports.

Second, food policy-makers, researchers, and agribusiness agents should grasp the importance of equation (7) in managing national pork-based food costs. While precipitous declines/rises in slaughter prices such as those that occurred during 1998-1999 may be taken by some as precursors to pronounced changes in pork-based food costs in the near post-event future, equation (7) suggests that there may be a significant tendency to manage risk along the current/futures price time horizon that can temper the severity of such slaughter price swings on downstream pork-related food costs and hence on U.S. pork-based food inflation.

That our model's finding that U.S. processor demand for slaughtered pork is dependent on relative slaughter/futures price rather than on slaughter price alone comprises a first step in resolving debates over whether commodity- or financially-focused policies/events have real effects on commodity markets. Aulerich et. al. note that debate rages over whether events such as the rise in futures trading volumes, expansion of trading pools to include more speculative profit-seeking traders, and rises in volatility of futures and related commodity prices have real commodity market effects. Our results for the modeled U.S. pork-based markets suggest that if a commodity demand is empirically tied to its relative own/futures price, and if future research can impute that such market events as a rise in futures trading volumes or the number of non-traditional traders affect one price in the relative slaughter/futures price ratio than another, then indeed such events will likely affect the underlying U.S. pork product markets through changes in U.S. processor demand for slaughtered pork. Analogously, if future research on other commodity markets with alternatively specified models can empirically discern that demands (or supplies) are tied to relative own/futures price ratios rather than solely to own prices, than perhaps such research can show real commodity market impacts of such policies/events as discussed above for U.S. pork product markets, be the policies/events financially focused, commodity-focused, or deal with changes in relative price volatility.

Policy Implications of the Cointegrated VAR's Event-Specific Binary Variables: The Case of the Uruguay Round's Effect on Slaughter Pork Demand

Since our model's estimation was in natural logarithms, we followed prior research and applied Halvorsen and Palmquist's well-known convention that permits one to use a binary variable's coefficient estimate to calculate an effect of the binary's defining event as an average percentage effect on the regression relation's dependent variable (Babula; Babula and Lund).[8] Such is another important and useful venue through which the cointegrated VAR may be used to generate policy implications. The Halvorsen-Palmquist (HP) value of -1.9% calculated for the Uruguay Round (UR) binary's coefficient estimate of -0.02 suggests that on average, U.S. processor demand for pork was 1.9% lower after the UR's implementation than prior to it. Such a sign may at first glance appear unexpected because of the well-known industry transformation noted by Morrison (1998). That is, that during the 1990's, the U.S. pork industry transformed from a major pork product importer to a major global pork exporter, as such U.S. pork-importing trading partners as Korea and Japan rendered concessions. Some may therefore have expected a positive β(URUGUAY) estimate. However, a number of interrelated considerations may rationalize the coefficient estimate's low-valued negative sign suggesting a negative effect on QSLAUGHT. First, given that the pseudo-t value of -2.22 has an absolute value that falls below that of the -2.6 critical value, evidence at the 5% significance level is not sufficient to reject the null that the CV2's β(URUGUAY) estimate is zero.[9] Hence the real effect of the UR is likely zero.

Second, the relationship between effects of the trade agreement on U.S. slaughter pork demand is not straightforward, insofar as U.S. processors not only slaughter pork that may be imported, but they also slaughter pork to manufacture processed pork-based food products downstream (ham, sausage, etc) that are exported. So the effect of trade agreements such as the UR on QSLAUGHT, a commercial slaughter volume that compised the only monthly

[8] As noted in Halvorsen and Palmquist for log/log estimations such as those comprising this paper's cointegrated VAR model, one takes e, the base of the natural logarithm; raises it to the power of the binary's β-estimate; subtracts 1.0; and multiplies the result by 100 to render the HP value on the binary's β-estimate. This HP value demonstrates the average percentage that the estimated equation's dependent variable was above/below during the period of occurrence of binary's defining event than during the rest of the sample when the event did not occur.

[9] As noted by Juselius (p. 142), the pseudo-t values in equations 6, 7, and 8 are not Student-t values and do not have the same critical values as defined for Student-t tests of the hypothesis that a regression coefficient is zero. The critical values for tests of the hypothesis that a cointegrated parameter value is statistically zero are +2.6/-2.6 at the 5% significance level.

U.S. pork quantity variable located with enough observations to do this study, is unclear and may account for its statistical insignificance.

Third, it is well-known that binary variable interpretation and the Halvorsen-Palmquist convention are subject to an important limitation: the limitation of imprecision. A binary coefficient, its HP value, and its implied effect on the regression's left-side variable, here the UR effect on U.S. slaughtered pork demand in CV2, cannot be attributed solely to the Uruguay Round's implementation, but as a sum total effect of all events that concurrently occurred during the period since January, 1995. This post-1994 period has certainly been one ridden by numerous trade-influencing events that occurred concurrently with the UR (economic, political, etc), and this array of concurrent events whose effects are captured by URUGUAY is increasing. Such an increasing array of concurrent and trade-influencing events may weaken β(URUGUAY)'s indication of UR-induced impacts on U.S. slaughter demand as time elapses since the January, 1995 UR implementation.[10] So given the insignificance of the URUGUAY β-estimate in CV2 and the ever-increasing set of concurrent events that have been occurring in the "UR" period since 1995, as well as the definition of the QSLAUGHT variable used here, we conclude that the indicated effect for the binary defined for the UR during the period of from 1995 to present may not be straightforwardly signed and is likely statistically zero.[11]

SUMMARY, CONCLUSION, AND POLICY INSIGHTS

We extended prior quarterly research and applied the cointegrated VAR model to a monthly U.S. system of upstream/downstream pork-based product markets; built-in a monthly linkage to a related futures market into this system; and provided three statistically strong and policy-relevant cointegrating relations that illuminate how the monthly markets dynamically interact. Second and in so doing, we provide an important methodological contribution in having demonstrated how cointegrated VAR methods can test for hedging (here at the U.S. pork slaughter point); how hedging activity can be incorporated into the finally estimated model through statistically supported restrictions; and how the final estimates may be interpreted to discern the role and policy implications that hedging has on the workings of the modeled markets and on the dynamic interactions of such markets.

Third, we used the cointegrating parameter estimates, particularly from the U.S. processor demand for slaughtered pork, to show how alternative policies/events with a financial focus and working through futures price are as effective in influencing the modeled U.S. pork product markets as commodity-focused policies and events working through slaughter price. This is because both sets of policies are equally effective in influencing relative slaughter/future price. Importantly, financially-focused and commodity-focused policies and events are equally effective policy levers with which U.S. policymakers can

[10] Had our sample ended earlier after the UR implementation than this study's December, 2011 endpoint, perhaps a more significant URUGUAY β-estimate would have emerged that would also have been more straightforwardly signed.

[11] The discussion of CV2's β(URUGUAY) estimate was provided to demonstrate an important cointegrated VAR capability of eliciting policy implications, and because of the importance of trade patterns since the early 1990's to the U.S. pork industry, as insightfully noted by an anonymous reviewer. Due to the insignificance or marginal significance of the other binary coefficients in equations 6-8 and because of space considerations, we do not discuss the other binary variable estimates.

address U.S. pork food product markets and manage patterns of U.S. pork-related food inflation.

Fourth, we illuminated the possibility of an important role that hedging upstream has in acting as an offset or cushion on effects that upstream policy changes or market events may have on downstream U.S. pork-related markets and related food cost patterns. We demonstrated how the model rationalized/explained hedging's role to resolve and cushion effects on downstream U.S. pork food costs from the extraordinary episodes of slaughter price volatility during late-1998/early-1999.

And finally, we discuss the importance for future research efforts not only for U.S. pork product markets, but for other commodity markets, in empirically tying demand/supply to relative own/futures prices rather to own prices alone, in resolving debates over alleged market effects of policies/events – be they commodity-focused that work through commodity prices or financially-focused that work through relevant futures prices.

REFERENCES

Andersen, L., Babula, R., Hartmann, H., and Rasmussen, M. (2007). A Time series analysis of Danish markets for pork, chicken, and beef. *Food Economics, 4,* 103-118.

Aulerich, N., Hoffman, L., and Plato, G. (2009). Issues and prospects in corn, soybeans, and wheat futures markets: New entrants, price volatility, and market performance implications. U.S. Department of Agriculture publication no. FDS-09G-01, August, pp. 1-44.

Babula, R. (2011). Using cointegrated VAR modeling to comparatively and empirically assess effects of alternatively-focused policies on U.S. soft wheat markets. *Journal of Internatinal Agricultural Trade and Development, 7*, 109-130.

Babula, R., and Lund, M. (2008). The Emergence of a U.S. demand for pork as an input: Applying the cointegrated vector autoregression model to U.S. pork-based markets. *Journal of International Agricultural Trade and Development, 4,*197-215.

Benson, Bruce, Faminow, M., Marquis, M., and Sauer, D. (1994). Intra-national effects of a countervailing duty on the United States/Canadian Hog Market. *Review of Agricultural Economics, 16,* 187-201.

Bessler, D. (1984). An analysis of dynamic relationships: An application to the U.S. hog market.' *Canadian Journal of Agricultural Economics, 32,* 109-124.

Bessler, D., and Akleman, D. (2004). Farm prices, retail prices, and directed graphs: Results for pork and beef. *American Journal of Agricultural Economics, 85,* 1144-1149.

Commodity Futures Trading Commission, Commitments of Traders (CFTC, COT 2011). Historical

Compressed Files found at the following link accessed by the authors on July 23, 2011: http://www.cftc.gov/MarketReports/CommitmentsofTraders/HistoricalCompressed/index .htm.

Dennis, J. (2006). *CATS in RATS: Cointegration Analysis of Time Series, Version 2.* Evanston, IL: Estima.

Duffy, R. (2004). Canadian pork industry faces long and costly litigation battle over U.S. countervailing and dumping challenge. *Better Pork Magazine,* June.

Estima (2007). *RATS: Regression analysis of time series, version 7*. Evanston, IL: Estima.

Goodwin, H., McKenzie, M., and Djunaidi, H. (2003). Which broiler part is the best part? *Journal of Agricultural and Applied Economics, 35,* (3), 483-95.

Granger, C., and Newbold, P. (1986). *Forecasting economic time series*. NY: Academic Press.

Halvorsen, R. and Palmquist, R. (1980). The Interpretation of dummy variables in smilogarithmic equations. *American Economic Review, 9*, 474-475.

Hamilton, J. (1994). *Time series analysis.* Princeton, NJ: Princeton University Press.

Hua, P. (1998). On Primary commodity prices: The Impact of macroeconomic/monetary shocks. *Journal of Policy Model Modeling, 20*, 767-790.

Johansen, S., and Juselius, K. (1990). Maximum likelihood and inference on cointegration: With applications to the demand for money. *Oxford Bulletin of Economics and Statistics, 52,* 169-210.

Juselius, K. (2006). *The cointegrated VAR approach: methodology and applications*. Oxford, UK: Oxford University Press.

Juselius, K., and Franchi, M. (2007). Taking a DSGE model to the data meaningfully. *Economics,1*, 1-38.

Juselius, K., and Toro, J. (2005). Monetary transmission mechanisms in Spain: The effect of monetization, financial deregulation, and the EMS. *Journal of International Money and Finance, 24,* 509-531.

Lambert, D. and Miljkovic, D. (2010). The sources of variability in U.S. food prices. *Journal of Policy Modeling, 32*, 210-222.

Morrison, N. (1996). U.S. pork's appeal has foreign buyers bellying up for more – Includes related article on Russian market. *AgExporter,* June-July.

Sims, C. (1980). Macroeconomics and reality. *Econometrica, 48*, 1-48.

Tiao, G., and Box, G. (1978). Modelling multiple time series: With applications. *Journal of the American Statistical Association, 76,* 802-816.

U.S. Department of Agriculture, Economic Research Service (USDA/ERS 2012). Commercial pork production, federally Inspected and other plants, based on dressed weights in millions of pounds. In *Meat Statistics, Red Meat and Poultry Production.* Accessed by the authors on February 14, 2012 at the following link: www.ers.usda.gov/publications/ldp/LDPTables.htm.

U.S. Department of Labor, Bureau of Labor Statistics (Labor, BLS 2012). Producer price index data bases. Accessed by an author on February 13, 2012 at the following link: www.bls.gov. Accessed December 23.

U.S. Department of Labor, Bureau of Labor Statistics (Labor, BLS 2007). Private email communication from Labor, BLS analysts to an author. The email provided an analysis of reasons for the extraordinary changes in slaughter pork PPI (PSLAUGHT) during 1998-1999.

World Trade Organization Panel (WTO Panel 1991). United States – countervailing duties on fresh, chilled and frozen Pork from Canada. Report DS7/7/R – 38S/30, July 11. Found at the following link accessed by the authors on December 23, 2011: http://www.sice.oas.org/dispute/gatt/89pork.asp.

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TRADE AGREEMENT IMPACT ON TRADE FLOWS, TRADE CREATION, AND TRADE DIVERSION: THE DETERMINANTS OF INTERNATIONAL WHEAT TRADE, 1999-2008

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ABSTRACT

This research identifies and quantifies the determinants of international wheat trade, during the period 1999-2008. Regression results provide evidence that economic factors play a major role in wheat trade, including domestic prices, national income, distance between nations, exchange rates, and inflation. Populations in both importing and exporting nations are also important determinants of the flow of wheat. This analysis extends previous wheat trade research through the study of multilateral relationships and bilateral trade agreements. Results show that World Trade Organization (WTO) membership enhanced wheat trade. This study examined the effects of both trade creation and trade diversion for all regional trade agreements. The model results show that nations which developed agreements with other nations that have diverse economic and resource characteristics are likely to see greater gains from free trade agreements. Nations that engage in agreements with nations with similar locations or income structures, trade diversion can occur.

Keywords: gravity model; trade creation; trade diversion; wheat trade.

JEL Classification codes: F14, Q17

INTRODUCTION

International trade is based on the simple notion of arbitrage: buying a product at a given location at a low price, then transporting it to another location and selling it at a higher price. Real-world international trade, however, is a complex interaction of economic, political, and social factors that encourage or deter the movement of goods from one location to another.

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The interworking of multilateral organizations, trade agreements, environmental controls, food safety policies, and cultural backgrounds creates a multifaceted trading system where numerous barriers and catalysts to trade exist, often in conflict with one another. The purpose of this research is to identify and quantify the determinants of international wheat trade during the time period 1999-2008. Wheat-specific traits, together with economic and policy variables, are found to be important trade determinants because of the evolving nature of trade policies, and the increased need for an efficient global supply chain.

These results are timely, interesting, and important because wheat is one of the most important staples in the world diet, and the highest-volume grain commodity traded in the world, averaging almost 90,000 Million Metric Tons (MMT) per year since 1960 with current annual volumes around 110,000 MMT (USDA/FAS, Production, Supply and Disapperance Online). Since the 1960s, the United States (USA), Canada, Australia, Argentina, and the European Union (EU) have dominated wheat exportation, consistently accounting for over 90 percent of wheat trade. In more recent years, the rise of the former Soviet Union nations of Russia (formally known as the Russian Federation), Ukraine, and Kazakhstan have increased the competition in wheat trade, as they have moved from being net importers of wheat to major exporters. Furthermore, their vast production potential, proximity to importers, and low pricing schemes have pressured established exporters to be more competitive to maintain market share. Better knowledge and understanding of the determinants of wheat trade, and the impact of trade agreements on wheat trade, is important given the large volume and trade value of wheat in an increasingly globalized world. International wheat trade has changed rapidly and dramatically during the period under investigation, requiring a better understanding of the changes and their causes.

Using data provided by the Global Trade Information System (2011), the trade of wheat at the four-digit Harmonized System (HS) code level will be evaluated for 36 exporters during the period 1999-2008, based on data availability. These wheat trade quantity data, combined with statistical data for economic and policy factors form the foundation of the empirical model estimated here. The main driving forces of wheat trade are price differentials across nations that arise from supply and demand shifts in each nation where wheat is produced and/or consumed. The gravity trade model is based on the economic size of two trading partners, and the distance between them. Thus, the model estimated here includes domestic prices, distance between nations, exchange rates, inflation, and population. Additional variables are included to capture the political determinants of wheat trade, including a nation's openness to trade, and trade agreements. The variables are quantified using a conventional gravity model, following Koo and Karamera (1991). The result is an econometric model that quantifies economic and political determinants of wheat trade.

LITERATURE REVIEW

Based on Newton's gravity equation in physics, gravity models were developed to quantify spatial flows (Yeboah et al. 2007). Early gravity models were applied to international trade by Tinbergen (1962), Poyhonen (1963) and Linneman (1966), who theorized that trade was determined by the relative size of each economy, divided by the distance between the two nations. Tinbergen (1962) provided little theoretical justification for the gravity model, but concluded that exporter income was the most important factor in the study of world trade flows, while preferential trade agreements (PTAs) did not have as large of an impact as expected.

Anderson (1979) derived a gravity equation from an expenditure system based on a Constant Elasticity of Supply (CES) utility function, allowing for a non-unity income elasticity and differentiation between traded and non-traded goods. Bergstrand (1985, 1989) identified a general equilibrium equation and expanded the microeconomic foundations of the gravity model.

Bergstrand (1989) also found that luxury goods, being capital intensive, would result in an elasticity of substitution greater than unity ($\sigma > 1$), causing the coefficients for exporter and importer per capita incomes to be theoretically positive. Wheat is likely to be a necessity good, which would have the opposite effect theoretically. Wheat is considered to be a capital-intensive, necessity good. Therefore, as income increases for exporters and importers, the percentage of wheat traded is expected to decrease.

Anderson and van Wincoop (2003) explained that trade deterrence comes from three sources (1) bilateral trade barriers, (2) exporters resistance to trade with all regions, and (3) importers resistance to trade with all regions: "Bilateral trade, after controlling for size, depends on the trade barrier between [nations] i and j, divided by the product of their multilateral resistance" (p 8). Our empirical framework is based on Koo and Karemera (1991), who analyzed factors influencing international wheat trade using a commodity-specific gravity trade model, and evaluated the influence of trade policies on commodity flows. We extend this to reflect current policies and trade agreement in wheat trade.

Grant and Lambert (2005) used a gravity trade model to focus on the effect of regional trade agreements (RTAs) on the flow of agricultural goods, and included variables that capture the possibility of trade creation and trade diversion. Since the work of Koo and Karemera (1991), the use of bilateral and regional trade agreements has become increasingly common in the global trading system.

To accurately reflect the development and implementation of trade agreements, trade creation and diversion measures were defined following Grant and Lambert (2005). These variables are included in the model in this analysis, highlighting the positive or negative influence of each agreement on wheat and more accurately reflect current, up-to-date, policy trends in international trade.

Trade policies that have been examined with a gravity equation include: Baier and Bergstrand (2007), Bayoumi and Eichengree (1995), Rose (2000), and Anderson and van Wincoop (2003). Vollrath, Gehlhar and Hallahan (2009) evaluated bilateral protection for agricultural industry products, including wheat, between 1986 and 2004, incorporating additional variables including the land/labor ratio, exchange rate misalignment, border protection, and colonial heritage. Tinbergen (1962) studied the effects of British Commonwealth preference on trade, and Koo and Karemera (1991) quantified the Economic Community trade agreement effect on international wheat trade.

Conceptual Model

The gravity trade model suggests that the bilateral flow of goods is based on the economic size of the two trading partners and the distance between them. The theoretical gravity model takes the following form (Tinbergen 1962):

$$X_{ij} = \beta_0 Y_i Y_j / D_{ij} \tag{1}$$

where i and j represent the exporter and importer, respectively. The flow of goods is noted by X, here defined to be the quantity of wheat in metric tons (MT), classified under the HS code 1001, which was exported from exporter (i) to importer (j) for the 36 largest wheat exporting nations and 86 importing partners, where an exporter can also be listed as an importer.

When a trade agreement is signed into law, two outcomes are possible: (1) trade creation, and (2) trade diversion. When two nations engage in an agreement, the reduction of tariffs reduces trade costs and can result in "trade creation," which enhances the movement of goods between those two nations, increasing welfare. Grant and Lambert (2005) quantified the effects of trade creation by defining a qualitative variable, equal to one when the year is greater than or equal to the date when the agreement was signed.

When trade costs diminish between nations through a trade agreement, this can also affect the world market, particularly for nations not directly involved in an agreement. The reduction in trade costs brought about by the agreement can make agreement partners more competitive. If the cost is reduced enough, the agreement can cause "trade diversion" from a traditionally more cost effective producer to an agreement partner. Qualitative trade diversion variables were defined and included in a gravity model by Grant and Lambert (2005), equal to one when trade occurs between an agreement and non-agreement member, and the year is greater than or equal to when the agreement was signed.

Empirical Specification

The variable Y is the economic "mass" of each nation, estimated by national GDP. The distance variable (D_{ij}) is defined as the non-weighted distance between the "most important city" in nations i and j. As distance between two trading partners increases, the cost of transportation increases, and the quantity traded is expected to decrease. The variable β_0 is a constant. Koo and Karemera (1991) extended this basic framework in their specification of an empirical gravity model. The authors included the economic variables: tariff (T_j), domestic wheat prices (P_i and P_j), exchange rates (E_{ij}), and inflation variables (I_i and I_j), using a log-log specification, as in equation (2):

$$X_{ij} = \beta_0 Y_i^{\beta 1} Y_j^{\beta 2} D_{ij}^{\beta 3} T_j^{\beta 4} P_i^{\beta 5} P_j^{\beta 6} E_{ij}^{\beta 7} I_i^{\beta 8} I_j^{\beta 9} * e_{ij} \tag{2}$$

where β_n are the parameters to be estimated, and e is an error term.

Following Koo and Karemera (1991), exporter income (Y_i) represents the production capacity of a nation, and importer income (Y_j) represents the nation's purchasing power. Although wheat production is typically a small portion of national income, exporter income is included as a scale factor to test empirically if the size of the economy has an influence on wheat exports. The trade barriers of distance (D_{ij}) and tariffs (T_j) are expected to negatively influence trade. Price variables (P_i and P_j) reflect the domestic price of wheat in each nation. Products are expected to move from a location with a low price to a location with a higher price. Inflation (I_i and I_j) complicates the world market, an exporter with relatively higher inflation is more competitive in the world market because it effectively decreases the price paid for the good by the importing nation. Exchange rates (E_{ij}) also play a strong theoretical role in international trade: as a currency strengthens against another, it will encourage imports and reduce exports of that nation. Exchange rate data (USDA/ERS) are real historical rates for base year 2005 in terms of U.S. dollars. To included variable E_{ij}, is in terms of exporter currency divided by importer currency. As a currency strengthens against its trade partner currency, it increases purchasing power, expected to result in a decrease in wheat traded. Exporters with a weak currency and importers with a strong currency have positive trade expectations. Thus, as the variable decreases, wheat trade is expected to increase.

The model reflects factors specific to the international wheat market. Koo and Karemera (1991) suggested that the tariff variable (T_j) in Equation 2 could be expanded to include factors that further encourage or deter trade. Several variables are included in this study to expand this T_j variable. Population captures aggregate demand for wheat for both importers and exporters. Capacity to produce value added agricultural goods (PAG_i) will also be included. This variable is measured by the production of value added agriculture goods as a percentage of GDP. Because wheat is a commodity good, as the percentage of value added agriculture production increases, the amount of wheat exported is expected to decrease. To quantify national policies toward trade in general, openness to trade variables were developed denoted as PT_i and PT_j. These varibles are defined as the share of trade in GDP, defined as the sum of exports and imports, divided by GDP. As the level of trade as a percentage of GDP increases, the more open a nation is to trade and the greater the amount of wheat flow can be expected. To capture the full effect of national income, ability to produce and willingness to trade, interaction terms among these variables are included in the model (Y_i*PT_i, Y_j*PT_j, and Y_i*PAG_i).

Grant and Lambert (2005) provided justification for the inclusion of colonial relationships, common languages, and contiguity of national borders. Each of these variables measure potential commonalities among nations that aid in international trade. Colonial relationships are measured by variables *COL*, *CCOL* and *COL45*. The variable *COL* signifies any colonial link between nations, past or present. The variable *CCOL* represents a common colonizer after 1945. The variable *COL45* denotes a colonial relationship after 1945. These variables are relevant to international wheat trade, because they represent potential similarities in political structure and culture, and a historical relationship between trading partners, thus increasing trade. Two variables were provided for common language by CEPII (Mayer and Zignago 2006). The variable *CLE* notes a common language spoken by more than nine percent of the population in both nations. The official common language shared between two nations is noted by *CLO*. Adjacency of nations is denoted by *CONTIG*, a qualitative variable if the exporter and importer share a common border (CEPII) (Mayer and Zignago 2006).

Multilateral relations are captured by variables $NONWTO_i$ and $NONWTO_j$. These variables signify nations which do not hold membership in the WTO, to quantify the effects that the WTO has on wheat trade. If the WTO has effectively reduced trade barriers for the wheat market, this should have a negative effect on nations who have not obtained membership in this organization. During the period of analysis, there has been an increase in the quantity of wheat traded. To account for this increase in wheat movement annual qualitative variables were included, with 1999 selected as the default year.

Thirty-seven bilateral trade agreements are included. This is a major contribution to the current wheat trade literature, as the evaluation of trade agreements at the commodity level is not common. Trade agreements have two potential effects (1) trade creation, the value of trade created between partners when a trade agreement is in place and (2) trade diversion which is the value of trade created as a result of diverting trade from non-agreement partners. The variable *CTRADE*, will take the value of one if a trade agreement between the two trading partners is active to denote the possibility of a trade creation impact of a trade agreements. Trade diversion (*DTRADE*) captures the impacts that agreements have on the market. Addition of these variables results in equation (3):

$$X_{ij}=\beta_0 Y_i^{\beta 1} Y_j^{\beta 2} D_{ij}^{\beta 3} P_i^{\beta 4} P_j^{\beta 5} E_{ij}^{\beta 6} I_i^{\beta 7} I_j^{\beta 8} POP_i^{\beta 9} POP_j^{\beta 10} PAG_i^{\beta 11} PT_i^{\beta 12} PT_j^{\beta 13} (Y_i{*}PT_i)^{\beta 14} (Y_j{*}PT_j)^{\beta 15} (Y_i{*}PAG_i)^{\beta 16} \mu^{CONTIG\beta 17} \mu^{LANG\beta 18} \mu^{COLONY\beta 19} \mu^{NONWTO_i \beta 20} \mu^{NONWTO_j \beta 21} \mu^{YEAR\beta 22} \Sigma_{n=1\ldots 37}\alpha_n CTRADE\ \Sigma_{m=1\ldots 35}\ \gamma_m DTRADE\ {*}e_{ij} \quad (3)$$

Following Grant and Lambert (2005), this study uses a semi-logarithmic functional form. Following previous literature, the model is estimated with OLS without fixed or variable effects, since the economic variables for each nation capture the national characteristics better than a country fixed effect variable.

Data

Based on data availability, the data are annual data from 1999-2008, which include 6,351 observations, provided by Global Trade Information Services Inc. (2011). Since the European Union has changed membership, the beginning year of the data set refelcts the first eyar when many EU nations had data available. The 36 included exporting nations accounted for approximately 93 percent of total wheat exported for the 2008/2009 Marketing Year (USDA/FAS Production, Supply and Disappearance Online). Table 1 provides the definition, summary statistics and source of each variable included in the model.

REGRESSION RESULTS

The regression results identify the characteristics of a nation that have the greatest influence on the movement of wheat across international borders. A Breusch-Pagan/Cook-Weisberg test confirmed heteroskedasticitity (Maddala and Lahiri 2009), thus the regressions were estimated with robust standard errors.

The interaction terms ($Y_i{*}PT_i$, $Y_j{*}PT_j$ and $Y_i{*}PAG_i$) included in the estimated model alter the calculation of the elasticities for the variables Y_i, Y_j , PT_i , PT_j and PAG_i (table 1). Marginal effects were evaluated at the mean. Table 2 presents the regression results for all variables other than trade creation (*CTRADE*) and trade diversion (*DTRADE*) variables (tables 3 and 4).

The marginal effect of exporter income (Y_i) is -0.164. The marginal of importer income (Y_j) reflects expected results, and is equal to -0.100. As nations become more affluent, they are more likely to import more processed or value added goods rather than commodity goods such as wheat.

This result is similar to those found by Koo and Karemera (1991), who estimated a coefficient value of -0.1305, significant at the one percent level. Grant and Lambert (2005) estimated the coefficient on Y_j to be equal to -0.63.

Exporter Percent Trade (PT_i) was not significantly different from zero. Because wheat is a staple good, this result can be interpreted to mean that exporters are not reliant on current trading volume relative to other goods. However, importer Percent Trade (PT_j) is significant, with a marginal effect equal to -0.004. The negative sign is logical because as trade increases, the quantity of wheat traded is expected to become smaller.

Table 1. Wheat Trade Gravity Model Variable Definitions

Variable Units	Definition	Source
Xij MT	Wheat traded between i and j in year t	GTIS
Yi,j bil. 2005 $	Gross Domestic Product	USDA/ERS
Dij km	Bilateral distance between economic centers	CEPII
Pi,j $/MT	Domestic price of wheat	UN/FAO
Eij Curri/Currj	Exchange rate	USDA/ERS
Ii,j %	Inflation	USDA/ERS
POPi,j Millions	Total population	World Bank
PAGi,j %	Share of value added agriculture in exporter GDP	World Bank
PTi,j %	Share of GDP from exports and imports	World Bank
Yi*PAGi --	Interaction between GDP and PAG	--
Yi,j*PTi,j --	Interaction between income and percent trade	--
CONTIG 1	Common border	CEPII
CLE 1	Common language	CEPII
CLO 1	Common official language	CEPII
COL 1	Colonial relationship	CEPII
CCOL 1	Common colonizer after 1945	CEPII
COL45 1	Colonial relationship after 1945	CEPII
NONWTOi,j 1	Not a member of the WTO	WTO
YEAR 1	Year of transaction	--
CTRADE 1	Trade agreement	WTO
DTRADE 1	Trade distortion	WTO

Note: CEPII is described in Grant and Lambert (2005) and Head, Mayer and Ries (2010).

Table 2. Wheat Trade Gravity Model Regression Results

Variable	Mean	SD	Min	Max	Coef.	SE	P>|t|	Elast.	% Chg
X_{ij}	135,905	135,905	1	7,282,273	--	--	--	--	--
Y_i	1,653.17	3,057.90	5.59	13,228.90	0.362	0.114	0.001	-0.164	--
Y_j	676.72	1531.38	0.72	13,228.90	-0.306	0.068	0.000	-0.100	--
D_{ij}	3,535.89	3,908.74	59.62	19,263.88	-0.140	0.065	0.030	--	--
P_i	148.82	64.51	55.9	655.1	-3.666	0.219	0.000	--	--
P_j	210.74	144.31	48.4	1,583.80	1.903	0.154	0.000	--	--
E_{ij}	9.29	39.43	2.58E-5	764.39	0.184	0.021	0.000	--	--
I_i	81.08	18.47	15.08	105.74	-0.379	0.213	0.076	--	--
I_j	80.9	17.71	4.2	127.52	0.366	0.218	0.094	--	--
POP_i	58.9	74.2	0.396	304	0.158	0.120	0.188	--	--
POP_j	51.3	141	0.388	1,320.00	0.717	0.070	0.000	--	--
PAG_i	3.97	3.28	0.36	15.86	0.369	0.053	0.000	-0.185	--
PT_i	78.44	41.67	18.97	326.76	-0.002	0.003	0.474	-0.016	--
PT_j	82.6	40.18	18.97	326.76	-0.017	0.003	0.000	-0.004	--
(Y_i*PAG_i)	21.47	14.65	1.32	66	-0.090	0.011	0.000	--	--
(Y_i*PT_i)	442.12	185.18	108.06	1,197.12	-0.002	0.001	0.000	--	--
(Y_j*PT_j)	386.71	205.23	-23.11	1,197.12	0.002	0.001	0.000	--	--
CONTIG	0.163	--	0	1	1.666	0.131	0.000	--	429.161
CLE	0.145	--	0	1	1.041	0.145	0.000	--	183.211
CLO	0.128	--	0	1	0.523	0.145	0.000	--	68.660
COL	0.082	--	0	1	-0.751	0.187	0.000	--	52.819

Variable	Mean	SD	Min	Max	Coef.	SE	P>|t|	Elast.	% Chg
CCOL	0.024	--	0	1	1.247	0.287	0.000	--	247.882
COL45	0.029	--	0	1	-0.122	0.321	0.704	--	11.471
$NONWTO_i$	0.110	--	0	1	1.489	0.163	0.000	--	343.255
$NONWTO_j$	0.146	--	0	1	0.173	0.204	0.396	--	18.903
2000	0.089	--	0	1	0.226	0.181	0.212	--	25.394
2001	0.092	--	0	1	0.072	0.180	0.689	--	7.439
2002	0.100	--	0	1	0.098	0.181	0.588	--	10.283
2003	0.097	--	0	1	0.591	0.188	0.002	--	80.527
2004	0.095	--	0	1	0.431	0.205	0.035	--	53.867
2005	0.103	--	0	1	-0.041	0.205	0.843	--	-3.990
2006	0.108	--	0	1	0.525	0.219	0.017	--	69.106
2007	0.105	--	0	1	1.556	0.298	0.000	--	373.963
2008	0.120	--	0	1	1.959	0.324	0.000	--	609.112
Constant	--	--	--	--	2.505	2.895	0.387	--	--
R-squared	0.4042								
Root MSE	2.8594								

Table 3. CTRADE Variable Regression Results

	Mean	Coeff	SE	P>|t|	% Chg
Armenia - Russia	0.002	0.852	0.863	0.323	134.547
Australia - NZ	0.002	5.092	0.412	0.000	16163.95
Canada - Chile	0.002	13.326	1.777	0.000	6.13E6
Chile - Mexico	0.002	3.355	0.395	0.000	2765.228
Common Econ. Zone	0.004	0.210	0.708	0.767	23.331
CIS	0.017	1.642	0.466	0.000	416.797
EFTA - Turkey	0.003	-3.677	0.598	0.000	-97.469
EU	0.305	1.592	0.162	0.000	391.422
EU - Albania	0.003	1.996	0.542	0.000	635.971
EU - Algeria	0.008	0.626	0.359	0.081	87.030
EU – Bosnia/ Herz.	0.003	0.637	0.435	0.143	89.066
EU - Croatia	0.006	-0.424	0.491	0.388	-34.567
EU - Egypt	0.003	0.347	0.462	0.453	41.455
EU - Israel	0.008	1.679	0.603	0.005	435.961
EU - Jordan	0.003	1.938	0.864	0.025	594.390
EU - Lebanon	0.003	-0.343	0.729	0.638	-29.062
EU - Macedonia	0.006	2.114	0.359	0.000	728.248
EU - Morocco	0.012	0.641	0.290	0.027	89.821
EU - Norway	0.014	1.199	0.362	0.001	231.757
EU – Switz/Liecht.	0.017	-1.672	0.348	0.000	-81.207
EU - Tunisia	0.011	1.114	0.273	0.000	204.728
EU - Turkey	0.011	-0.587	0.401	0.143	-44.380
Eurasian Econ. Comm.	0.007	-2.463	0.703	0.000	-91.478
Georgia - Russia	0.002	-0.253	0.937	0.787	-22.386
GSTP	0.006	-4.194	0.964	0.000	-98.492
NAFTA	0.008	0.056	0.368	0.878	5.805
PICTA	0.001	-0.050	0.390	0.899	-4.831
Protocol on Trade Neg.	0.010	1.823	0.505	0.000	519.216
MERCOSUR	0.004	0.844	0.556	0.129	132.512
Turkey - Georgia	0.000	3.585	0.473	0.000	3505.750
Ukraine - Azerbaijan	0.000	-2.775	1.686	0.100	-93.764
Ukraine - Belarus	0.000	0.727	1.462	0.619	106.923
Ukraine - Moldova	0.001	-2.119	0.814	0.009	-87.990
Ukraine - Russia	0.001	-3.036	0.960	0.002	-95.196
United States - Chile	0.001	0.685	0.583	0.240	98.301
United States - Israel	0.002	1.412	0.655	0.031	310.288
United States - Morocco	0.000	0.983	0.619	0.112	167.186

Since wheat is a raw commodity, as a nation increases international trade, it is likely to trade more processed goods. Exporter Percent Agriculture (PAG_i) is significant at the one percent level, with a marginal elasticity of -0.185. This indicates that value added agricultural

products are more likely to be traded than a commodity (staple) such as wheat, following trade theory. As nations become more affluent, they are likely to substitute out of trade in higher value added goods, and into commodity goods.

The influence of distance as an approximation for transportation costs (D_{ij}) is highly significant and equal to -0.140. This implies that trade agreements may have the capacity to reduce tariffs in a manner that compensates for the costs of shipping. When tariffs are lower, nations can afford to import products from greater distances.

The domestic prices of a good, P_i and P_j, are both statistically significant at the one percent level, with the largest estimated elasticities for economic variables. Endogeneity between domestic prices and trade flows is a possibility, leading to the potential for biased estimates. However, the bilateral trade flows reported in the data set are small relative to the domestic wheat market in each nation. Current prices are used, rather than exogenous lagged prices to more accurately reflect actual market conditions. Instrumental variables for current prices are not readily available.

This confirms classical trade theory, as it highlights the movement of goods from an area with a low domestic price to an area with a high domestic price. For an exporting nation P_i, a one percent increase in domestic price would decrease the quantity exported by 3.666 percent. A one percent increase in the domestic price of wheat for an importing nation (P_j) would cause a 1.903 percent increase in the quantity of wheat traded.

The exchange rate (E_{ij}) is positive and statistically significant at the one percent level: for a one percent increase in the importer currency relative to the exporter currency results in quantity of wheat traded increasing by 0.184 percent. Inflation for exporting nations (I_i) is significant at the ten percent level with an elasticity of -0.379. A one percent increase in the inflation rate (I_j) results in a 0.366 percent increase in wheat imports. As the value of wheat increases in importing nations due to inflation, consumers demand a relatively lower priced product from the international market. Exporter population (POP_i) is insignificant. Importer population (POP_j) is an indication of the demand for wheat. This coefficient is significant at the one percent, with a coefficient of 0.717.

The variable *CONTIG* is significant at the one percent level, with a positive coefficient of 1.666, which results in a 429.161 percent change in wheat flow when nations share a common border. This indicates that nations are more likely to trade wheat when they share a common border, as expected.

A common language is significant at the one percent level. The variable *CLE* has a positive coefficient of 1.041. This coefficient exhibits a 183.211 percent increase in wheat traded between nations in which nine percent of the population shares a common language. This could imply that a common language is important in wheat trade, with an increased ability to communicate with other nations resulting in more trade.

Common Official Language, *CLO*, is significant at the one percent level with a coefficient of 0.523. This suggests that a common official language is important in wheat trade, increasing trade 68.660 percent. Similar to *CLE*, *CLO* might indicate that an increased ability to effectively communicate with other nations results in more wheat trade between nations who share a common language.

Table 4. DTRADE Variable Regression Results

DTRADE Variable	Mean	Coeff	SE	P>|t|	% Chg
Armenia - Russia	0.003	1.159	0.592	0.050	218.641
Australia - New Zealand	0.004	-0.369	1.283	0.774	-30.829
Canada - Chile	0.005	6.778	1.327	0.000	87722.515
Chile - Mexico	0.006	-6.398	1.257	0.000	-99.834
CEZ	0.013	-1.890	0.512	0.000	-84.897
CIS	0.047	-1.231	0.451	0.006	-70.810
EU	0.102	1.074	0.202	0.000	192.673
EU - Albania	0.001	1.818	1.072	0.090	516.152
EU - Algeria	0.004	1.715	0.651	0.009	455.488
EU - Croatia	0.001	2.139	0.552	0.000	748.900
EU - Egypt	0.005	1.109	0.385	0.004	203.149
EU - Israel	0.008	1.258	0.645	0.051	251.875
EU - Jordan	0.004	1.908	0.519	0.000	573.937
EU - Lebanon	0.004	1.875	0.407	0.000	551.948
EU - Macedonia	0.001	2.097	0.576	0.000	714.527
EU - Morocco	0.008	0.746	0.251	0.003	110.934
EU - Norway	0.007	-0.426	0.488	0.383	-34.707
EU – Switz./Liecht.	0.009	-0.389	0.431	0.367	-32.204
EU - Tunisia	0.007	1.338	0.319	0.000	281.307
EU - Turkey	0.006	-0.568	0.628	0.366	-43.329
Eurasian Econ. Comm.	0.026	-0.040	0.509	0.937	-3.921
Georgia - Russia	0.006	0.075	0.546	0.891	7.769
GSTP	0.162	1.442	0.169	0.000	322.824
NAFTA	0.016	-3.835	0.339	0.000	-97.839
PICTA	0.003	-1.575	1.515	0.299	-79.303
Protocol on Trade Negotiations	0.117	0.704	0.189	0.000	102.206
MERCOSUR	0.010	-2.880	0.530	0.000	-94.384
Turkey - Georgia	0.001	2.369	0.686	0.001	969.193
Ukraine - Azerbaijan	0.004	-0.352	0.559	0.529	-29.694
Ukraine - Belarus	0.009	2.028	0.526	0.000	659.552
Ukraine - Moldova	0.006	0.473	0.589	0.422	60.514
Ukraine - Russia	0.024	-0.861	0.447	0.054	-57.728
United States - Chile	0.005	-3.628	0.794	0.000	-97.344
United States - Israel	0.016	-0.108	0.567	0.849	-10.217
United States - Morocco	0.007	0.188	0.327	0.565	20.681

Colonial relationships were represented by three variables. The variable *COL*, indicating a past or present colonial tie is significant but negative at the one percent level with a coefficient of -0.751. This implies that past colonial relations had a negative effect, decreasing wheat trade 52.819 percent among importers and colonially linked exporters. The variable *CCOL*, noting common colonies, exhibits a positive coefficient of 1.247 significant at the one percent level. Colony origins often result in similar cultures and political structures, leading to increased trade by 247.882 percent. The variable *COL45* is statistically insignificant. This implies that even though there are several nations with colonial relationships after 1945, they have no positive or negative influence on the movement of wheat.

The variable $NONWTO_i$ is significant at the one percent level with a coefficient of 1.489, indicating a 343.255 percent increase in wheat export for these nations. Kazakhstan, Russia and Ukraine are the only exporters classified under this variable. This implies that for these three nations, not being a member of the WTO had a positive effect on their ability to export wheat. One explanation for this could be the lack of liberalization achieved by the WTO for agricultural goods. Because little has been done to reduce barriers for agriculture, there is not a negative implication for these nations relative to other exporters. However, it is more likely that is variable is capturing the export capabilities and growth in wheat production of these three nations.

For importing nations, $NONWTO_j$ is statistically insignificant, signifying that even with a larger number of nations in this category, membership in the WTO has no measurable effect on the trade of wheat during the time period analyzed. The year variables indicate that a significantly higher volume of wheat was traded in 2003, 2004, 2006, 2007 and 2008, compared to the base year 1999.

The inclusion of up-to-date trade agreement variables is one of the main contributions of this analysis to wheat trade literature. Table 3 presents the coefficients, standard errors, P-values, and percentage changes for the trade creation (*CTRADE*) qualitative variables. Sixteen of the 37 agreements resulted in positive and significant *CTRADE* coefficients, demonstrating an increase in wheat trade after trade agreement were initiated. Seven agreements had negative, statistically significant coefficients, indicating that some trade agreements resulted in lower levels of wheat trade. These agreements were concentrated between the EU and the former Soviet states, reflecting an increase in domestic wheat production in the new nations, and a declining demand for imported wheat from Europe.

The bilateral agreement between Canada and Chile has the highest magnitude effect on the trade of wheat, with a coefficient of 13.326. Therefore, when Canada exports to Chile, the quantity is almost 61,299,887 percent higher than the quantity Canada would export to any other nation because of this agreement. The large magnitude results from the unlikely probability that the trade agreement between Canada and Chile had an impact on wheat trade to any other nation. Restated, the agreement only affected Chilean wheat trade. Similar interpretations can be developed for the other agreements. Many of the agreements that exhibit trade creation are between nations in different regions of the world, political structures and economic status. This implies that although commonalities between nations, such as common languages, adjacency, and colonial heritage, encourage trade between nations, benefits from trade are gained from the diversity of nations involved in an agreement.

When the coefficient of *CTRADE* is negative, the trade of wheat is suppressed under the given agreement. An interesting observation in this category is that several of the nations are located in the Baltic region. This suppression of trade could be a result of homogeneity among nations in terms of location, or perhaps factors such as political structure and ability to produce wheat. If these nations are more homogenous, this will decrease their incentive to trade wheat relative to other goods.

Table 4 displays the coefficients, standard errors, P-values, and calculated percentage change for each of the trade diversion variables (*DTRADE*). Eleven of the 35 *DTRADE* variables were not statistically significant different from zero, indicating that there was no measureable trade diversion as a result of these agreements. Sixteen *DTRADE* variables had positive and significant coefficients. These coefficients indicate that despite decreased tariffs brought about by agreements, wheat trade actually increased between agreement members and non-members. Using the agreement between the Canada and Chile as an example, the regression resulted in a coefficient of 6.778. This is used to calcualte that Chile imports 87722.515 percent more from non-agreement members than from the Canada, thus the tariff reduction provided for wheat by this agreement was not enough to warrant imports of wheat from Canada.

Inversely, a negative *DTRADE* coefficient indicates that because of the trade agreement, wheat is more likely to be purchased from an agreement member than a non-agreement member. Seven agreements in this study displayed this form of trade diversion. One explanation of this is that the reductions in tariffs imposed by the trade agreement make the product from agreement members more competitive. For example, the agreement between Russian and the Ukraine has a coefficient of -0.861. After transforming this coefficient, it is clear the as a result of this trade agreement 57.728 percent of the wheat trade between Russian and the Ukraine is a result of trade diversion.

Although some of the trade agreements had an overall negative effect on wheat trade, it is important to remember that when developing trade agreements, wheat is not the only consideration. In trade, there are always benefits and costs. Trade agreements are only successful if the total benefits to each nations involved is greater than the costs, even if it means that some producers are harmed due to an agreement. Wheat industry participants should evaluate the effects of each proposed trade agreement carefully to ensure that they are benefiting wheat producers accordingly.

IMPLICATIONS AND CONCLUSION

International wheat trade remains a complex interaction of economic, political and social factors that encourage or deter the movement of wheat from one location to another. A gravity model was developed to identify and quantify the determinants of international wheat trade for the time period 1999-2008, and the possibility of trade creation and trade diversion for relevant trade agreements was tested.

The regression results provide evidence that economic factors continue to play a major role in wheat trade. Domestic price plays a key role in the movement of wheat. The regression results highlight that national income, transportation costs, exchange rate, inflation, and importer population are important and statistically significant in determining the flow of wheat. Quantitative measures of openness to trade and agriculture production were statistically significant, and highlight the commodity nature of wheat trade. It was shown that qualitative factors also play a key role in determining wheat trade. Sharing a common border and language are also positive factors in wheat trade. Colonial heritage does have a small but statistically significant impact, negative for direct colonies and positive for nations sharing a common colonizer. Although these factors are important in determining the flow of wheat, there is little that commodity organization and policy making bodies could do to modify these attributes.

This analysis extends previous international wheat trade research through the inclusion of multilateral relationships and bilateral trade agreements into the gravity model. To highlight multilateral relationships, membership in the WTO was considered. This study showed that not being a member of the WTO is a positive factor in wheat trade. As stated earlier, this could be a result of a lack of liberalization accomplished through the current multilateral system for agricultural goods or highlight the attribution of nations not involved in the WTO. For WTO membership to have a positive effect on the trade of wheat, this organization will need to further deplete trade barriers for agricultural goods among its members to make a significant impact.

The review of bilateral trade agreements added depth to this study by examining both trade creation and trade diversion for each applicable agreement. The estimated models show that nations which develop agreements with contrasting qualities from themselves are likely to see higher gains from free trade agreements. However, when nations engage in agreements with nations in a similar location or income structure, trade diversion occurs. This is an important consideration for policy making bodies considering engaging in trade agreements. The analysis also shows that trade agreements can also overcome factors that may have a negative impact on the trade of wheat such as distance or colonial relationships. Wheat is not the only product that should be considered when developing a trade agreement, therefore trade agreements may have an adverse impact on wheat trade in some circumstances. However, national wheat promotion boards need to seriously consider both the potential for positive and negative impacts of each proposed agreement.

Additional areas of research that could contribute to further our knowledge of wheat trade include: evaluating specific policy mechanisms, performing a more in-depth evaluation of the Baltic region exporters, and evaluation of the influence of environmental mandates. As the face of international trade continues to evolve, there will be a need to continue to analyze the determinants of international wheat trade. Consumers, producers, wheat promotion boards, governments and international policy making bodies all need to be aware of the impacts that their actions are having on the flow of this staple commodity. By understanding the determinants of wheat trade, players in the wheat market can create a more transparent and fluid market.

REFERENCES

Anderson, J. E. 1979."A Theoretical Foundation for the Gravity Equation." *American Economic Review* 69(1):106-116.

Anderson, J. E. and E. van Wincoop. 2003. "Gravity with Gravitas: A Solution to the Border Puzzle." *American Economic Review* 93:170-192.

Anjaria, S. J. 1987. "Balance of Payments and Related Issues in the Uruguay Round of Trade Negotiations." *World Bank Economic Review* 1(4): 669–88.

Baier, S. L., Bergstrand, J. H. 2007."Do Free Trade Agreements Actually Increase Members' International Trade." *Journal of International Economics* 71(1): 72–95.

Bayoumi, T. and B. Eichengreen. 1995. "Is Regionalism Simply a Diversion? Evidence from the Evolution of the EC and EFTA." CEPR Discussion Papers, No. 1294, London: Centre for Economic Policy Research.

Bergstrand, J. H. 1985. "The Gravity Equation in International Trade: Some Microeconomic Foundations and Empirical Evidence." *Review of Economics and Statistics*, 67(3):474-81.

Bergstrand, J. H. 1989. "The Generalized Gravity Equation, Monopolistic Competition, and the Factor-Proportions Theory in International Trade." *Review of Economics and Statistics*, 71(1):143-53.

Centre d'Études Prospective et d'Informations Internationales (CEPII). http://www.cepii.fr/

Global Trade Information System(GTIS). 2011. http://www.gtis.com/english/ (accessed March 11, 2011).

Grant , H. J. and D. M. Lambert. 2005. "Regionalism in World Agricultural Trade: Lessons From Gravity Model Estimation". Paper presented at the American Agricultural Economics Association Annual Meeting, Providence, RI, July 24-27, 2005. http://ageconsearch.umn.edu/bitstream/19269/1/sp05gr01.pdf (accessed September 10, 2011).

The Harmonized System. 2011. http://export.gov/logistics.eg_main_018119.asp (accessed September 9, 2011).

Head, K., T. Mayer, and J. Ries. 2010. "The erosion of colonial trade linkages after independence." *Journal of International Economics* 81 (1):1–14.

Koo, W.W. and D. Karemera. 1991. "Determinants of World Wheat Trade Flows and Policy Analysis." *Canadian Journal of Agricultural Economics* 39 (3):439-55.

Linneman, H. 1966. *An Econometric Study of International Trade Flows*. Amsterdam: North Holland.

Mayer, T. and S. Zignago. *CEPII Distances Database.* May 3, 2006. http://www.cepii.fr/distance/noticedist_en.pdf (accessed September 11, 2011).

Maddala, G.S., and K. Lahiri. 2009. *Introduction to Econometrics.* 4th Edition. Chichester, West Sussex: John Wiley and Sons Ltd.

McCallum, J. 1995. "National Borders Matter: Canada-U.S. Regional Trade Patterns." *American Economic Review* 85(3): 615-23.

Poyhonen, P. 1963. "A Tentative Model for the Volume of Trade between Countries." *Weltwirtschaftliches Archiv* 90:93-99.

Rose, A. K. 2000. "One Money, One Market: Estimating the Effect of Common Currencies on Trade" *Economic Policy* 30, 9-45.

Tinbergen, J. 1962. "Shaping the World Economy: Suggestions for an International Economic Policy." New York: Twentieth Century Fund.

United Nations Food and Agriculture Organization. August 9, 2011. http://www.fao.org/ (accessed September 10, 2011).

United Nations Statistics Division. 2011. http://unstats.un.org/unsd/default.htm (accessed September 10, 2011).

United States Department of Agriculture. Economic Research Service. (USDA/ERS) http://www.ers.usda.gov/ (accessed September 10, 2011).

United States Department of Agriculture. Foreign Agricultural Service. (USDA/FAS) http://www.fas.usda.gov/ (accessed September 10, 2011).

United States Department of Agriculture. Foreign Agricultural Service. (USDA/FAS) *Production, Supply and Disapperance Online.* http://www.fas.usda.gov/psdonline (accessed September 9, 2011).

Vollrath, T. L., M. J. Gehlhar, and C. B. Hallahan. 2009. "Bilateral Import Protection, Free Trade Agreements, and Other Factors Influencing Trade Flows in Agriculture and Clothing." *Journal of Agricultural Economics*, 60(2): 298-317.

World Data Bank. July 28, 2011. http://databank.worldbank.org/ddp/home.do (accessed September 10, 2011).

World Trade Organization. 2011. http://www.wto.org/index.htm (accessed 9/9, 2011).

Yeboah, O., S. Shaik, A. Allen, and V. Ofori-Boadu. "Trade Effects of the Central American Free Trade Agreement." Presented at the American Agricultural Economics Association Annual Meeting. Portland, Oregon, 2007. http://ageconsearch.umn.edu/bitstream/9815/1/sp07ye01.pdf (accessed 9/10/11)

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MARKET INTEGRATION AND RELATIONSHIP BETWEEN FARM-LEVEL PRICES: EVIDENCE FROM CHERRY MARKETS IN BC, WASHINGTON AND CALIFORNIA

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ABSTRACT

An understanding of the behavior of farm-level cherry fruit prices has implications for tree fruit investment decisions and market development strategies.

This paper investigates market integration, price leadership, and price formation in British Columbia (BC), Washington and California cherry markets over the period 1971-2011. Employing annual farm-level prices and Engle-Granger and Johansen co-integration/error correction techniques, this paper examines whether cherry prices have converged in BC, Washington and California during a period characterized by greater trade integration between Canada and the United States.

Results show that BC and Washington markets are strongly integrated with a high degree of price transmission, while BC and California markets are partially integrated based on trends in total bilateral trade.

The price formation analyses reveal BC fresh cherry prices are influenced significantly by their own per capita farm-level quantities and Washington per capita farm-level quantities.

In contrast, Washington and California cherry prices are mainly explained by their own per capita farm-level quantities.

Keywords: Canada; cherry markets; error correction modeling; market integration; the United States
JEL Classification codes: Q17, Q13

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INTRODUCTION

For many decades economists and historians have been preoccupied with the subject of commodity market integration and the analytical techniques and data for measuring it. Market integration or convergence in commodity prices have been studied in geographic regions of the world, such as Europe, where the integration of national economies have been fostered by the reduction of trade barriers (Goldberg and Verboven, 2005; Federico, 2010). Price differences for identical or homogenous products between spatial markets have been shown to be influenced by transport costs, product quality differences and trade barriers. Since 1988, the Canada-U.S. Free Trade Agreement (CUSFTA) has resulted in the lowering of tariff and non-tariff barriers resulting in large volumes of trade in fruits and vegetables between both countries. In 2009, the U.S. accounted for U.S $4.1billlion or 59% of the value of fruits, vegetables and their products imported into Canada (Agriculture and Agri-Food Canada, 2011). High value fresh sweet cherries are an important stone fruit traded between both countries. Between 1988 and 2011, Canada fresh cherry imports from the U.S. increase 13-fold in value from $ 10.3 million in 1988 to $137.4 million in 2011 (Statistics Canada, 2012a).

Sweet cherries have increased in popularity over the last decade in part due to reported health benefits on the demand side, while their relatively high price premiums have contributed to increased supplies (Kahlke et al., 2009). The top three producing countries in the world are Turkey, U.S. and Iran followed by Italy and Spain (Perez and Plattner, 2012). The U.S. is the world's second largest cherry producer accounting, for over 10% of world production and 19% of world exports, with Canada the largest importer of U.S. sweet cherries averaging 12,894 metric tonnes (mt) at a value of $60 million between 2004 and 2006 (USDA, 2007). While British Columbia (BC) is the major sweet cherry producer in Canada, Washington and California are the principal fresh sweet cherry producing states in the United States (Boriss, Brunke, and Kreith, 2011). Approximately 90% of the U.S. sweet cherry crop is produced in Washington, California and Oregon (Perez and Plattner, 2012). In 2011, the value of Washington and California fresh sweet cherry farm production totaled $515 million and $252million, respectively, while BC sweet cherry sales in British Columbia amounted to $31 million (USDA, 2012a; Statistics Canada, 2012b).The bulk of sweet cherries produced in BC, Washington and California are sold at the fresh market.

While the few Washington cherry studies undertaken so far have focused on price formation relationships (McCracken, Casavant and Miller, 1989; Schotzko, Wilson and Swanson, 1989; Flaming, Marsh and Wahl, 2007), there has been little empirical research employing time series data to analyze market integration and the spatial relationships of cherry prices between Canada and the United States. Market integration study results have varied depending on the structure or segment of the market (farm, wholesale, retail), commodity in question, frequency of the data, and whether spatial or temporal prices are employed (Barrett, 2001). There have been a few studies in the literature (Vollrath and Hallahan, 2006; Ghoshray, 2007; Miljkovic, 2009) that have looked at whether markets are integrated between Canada and the U.S. as a whole. Ghoshray (2007) employed monthly trade price data to study the relationship between Canadian and U.S. durum markets. The author found that changes in Canadian domestic policy in the mid-1990s resulted in greater integration of durum markets between Canada and the U.S. It was shown from co-integration tests that livestock and meat product markets are integrated between both countries despite different responses of Canadian and U.S. livestock prices to exogenous trade shocks (Miljkovic, 2009). Given the degree of market integration between the two economies, one would have expected that, for exchange rate adjusted prices, the prices Canadians would pay for similar products should equate U.S. prices, according to the law of one price. However, a recent Canadian study for several product groups showed that prices in Canada and the U.S. generally do not match or compare, since substantial variability in relative prices is closely associated with changes in the market exchange rate (Gellatly and Yan, 2012).These results are similar to an earlier study by Engel and Rogers (1996) who found that market segmentation or the variation in prices of similar goods between Canada and the U.S. was attributed to a combination of distance between markets and the price-setting process.

The objective of this article is to investigate whether sweet cherry markets in BC, Washington and California are integrated or share a long-run price relationship by employing co-integration and error-correction models. A secondary objective is to quantify the factors that impact price formation in cherry markets in the Pacific Northwest (e.g., BC, Washington) and California. This is important since it will identify the supply and demand factors that explain the geographic variation in cherry prices.

The contribution of this study to the literature is twofold: first, we apply cointegration techniques and vector error correction models to farm-level data from BC, Washington and California; second, we focus on a differentiated and high quality perishable stone fruit, per capita consumption of which in North America has increased over the years, and that is harvested, shipped, and traded, primarily in the months of May-August. The aim is to investigate the degree of price transmission between BC, Washington and California cherry markets and the nature of price leadership in cherry markets in the Pacific Northwest and California; both markets differ in climate, geography, varieties grown, production scale, and marketing arrangements. The investigation involves using co-integration techniques and error-correction models to examine long-run price relationships in cherry markets. Although sweet cherry demand studies have been undertaken in the U.S., little market integration research has been undertaken in BC; a region that imports a significant amount of fresh sweet cherries from the U.S., especially from Washington State. According to O'Rourke and Casavant (1974), the sweet cherry industry in the Pacific Northwest is so interrelated that it is not statistically or economically sound to regard Washington sweet cherry market as independent of cherry demand in other regions of the Pacific Northwest.

The structure of the article is as follows: In the first section a description of the production structure and trade in fresh cherry trade is provided. This is followed by a discussion of the methodology, estimation strategy, description of the data and empirical results. Discussion and implications are provided in the final section.

BC, WASHINGTON AND CALIFORNIA CHERRY PRODUCTION

Productivity growth has been the principal driver responsible for the expansion of sweet cherry production in BC, Washington and California. Since the early 1990s, there have been significant increases in fresh sweet cherry production, especially for British Columbia (Figure 1).

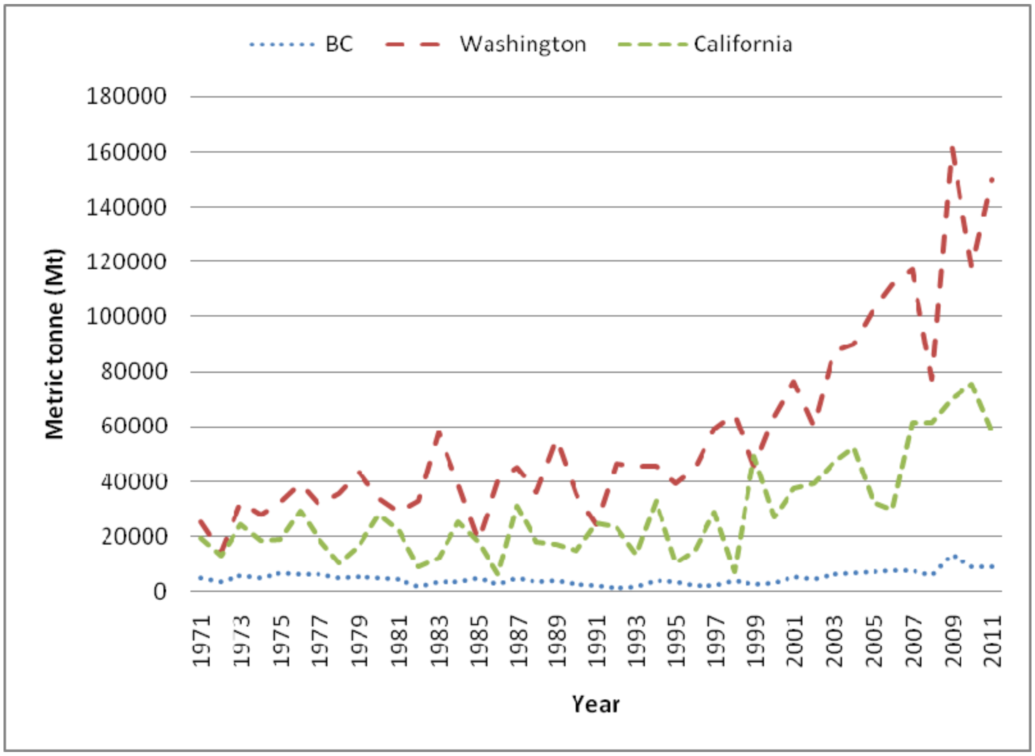

Figure 1. Farm-level fresh sweet cherry production in BC, Washington and California.

Washington and California fresh sweet cherry production increased, respectively, from an average of about 38,585mt and 20,260mt in the period 1991-93 to about 143,033mt and 67,736mt in the period 2009-11, while the BC fresh sweet cherry production increased by 593% from 1,528mt in years 1991-93 to 10,590mt in the period 2009-2011 (Figure 1). Production expansion in BC has been driven by a combination of demand increases in both domestic and global markets, adoption of late season cultivars, and technology developments in crop management and packaging (Kappel, 2006).

Washington is the largest producer in the U.S. with its production share increasing slightly from 64% in 1991-93 to 65% in 2009-11, while BC's production share increased from 3% to 5% over the same time period. California's production share decreased from 34% to 31% over the same time period.

Sweet cherry production expansion in Washington and California has been due to area expansion and yield increases. Yields are likely to vary from year to year and are attributed to differences in climate conditions and characteristics of the varieties grown. For example, the lower Washington's production in 2008 (Figure 1) was because of poor pollination caused by cool temperatures during the bloom period and extensive crop damage from the mid-April frost, both of which affected mostly the earlier maturing varieties (Perez and Pollack, 2008).

In California, most of the varieties mature early in the season. The 'Bing' variety accounts for 60-70% of the sweet cherry varieties packed by domestic shippers with large shipments exported to Japan, its most important lucrative overseas market (Grant, 2009). Besides 'Bing', other important varieties packed by California shippers included 'Tulare', 'Brooks', 'Rainier', and 'Sequoia'. Unlike their northern counterparts, California producers are not interested in extending their cherry season with late season varieties because varieties ripening after 'Bing' would increase the harvest season overlap with the Northwest cherry producing states (Hansen, 2010). Overlapping in harvest seasons as a result of adverse weather conditions is likely to exert downward pressure on prices. In 2003, cool and wet weather conditions delayed the start of California's cherry season, and resulted in increased supplies which lowered California and Washington cherry prices (Pollack and Perez, 2004).

Sweet cherries have a relatively short marketing season, with California's cherry shipments beginning in late April and ending in the late June, while Washington's marketing season is from June through mid-July (Perez and Pollack, 2007). Over the past decade, California has diversified its variety mix towards early maturing varieties, while Washington has shifted towards the late maturing varieties (Grant, 2012). The top five Washington sweet cherry varieties grown in 2011 are 'Bing' followed by' Sweetheart', 'Rainier', 'Skeena' and 'Chelan' (USDA, 2011).

For late maturing varieties, the harvest season continues to around mid- to- late August. BC growers, in general, tend to grow late maturing varieties to minimize the overlap between Washington and BC's harvest seasons.'Lapins', 'Sweetheart', 'Staccato', and 'Santina' were the principal late maturing varieties accounting for 85% of new variety re-plantings in BC (Okanagan Tree Fruit Authority, 2010).

'Sweetheart', developed by Agriculture Canada's breeding program and released in 1994, has become popular in BC and Washington because it is a late maturing cultivar that allows growers to improve productivity and extend the marketing season. According to World Sweet Cherry Review (2011), 'Sweetheart' has gained rapid acceptance as a mid-to-late season variety in Washington while 'Bing', the historically established variety there, is losing its notoriety.

British Columbia-U.S. Cherry Trade

Historical trade statistics between Canada and the U.S. have been available over many decades. However, Canadian provincial/U.S. state trade statistics under the Harmonized Commodity Description were not available until 1988 (Statistics Canada, 2012a). British Columbia is a net importer of cherries from the U.S. with imports increasing from \$1.8 million in 1988 to \$31 million in 2011 (Statistics Canada, 2012a).

Under North American Free Trade Agreement (NAFTA) preferential tariff, cherry trade between both countries is duty free. While BC's cherry production has been increasing over the last two decades, it has expanded its volume of sweet cherry imports from the U.S. The increase equaled about 300%, from an average of 1,421mt in 1988-90 to 5,676mt in 2009-11. Following the implementation of NAFTA on January 1st ,1994, there has been a steady upward trend in BC cherry imports from the U.S.(Figure 2). The bulk of BC cherry imports from the U.S. occur in the months of May-August. Since 2000, BC cherry exports to the world and the U.S. have increased significantly (Figure 3). In the early 2000s BC started to diversify its export markets to Asian (Taiwan, Hong Kong) and European countries (United Kingdom, The Netherlands).

Therefore, it is evident from the graphical import and export analysis that BC-Washington cherry markets are integrated since there is a fair amount of trade taking place between both border regions.

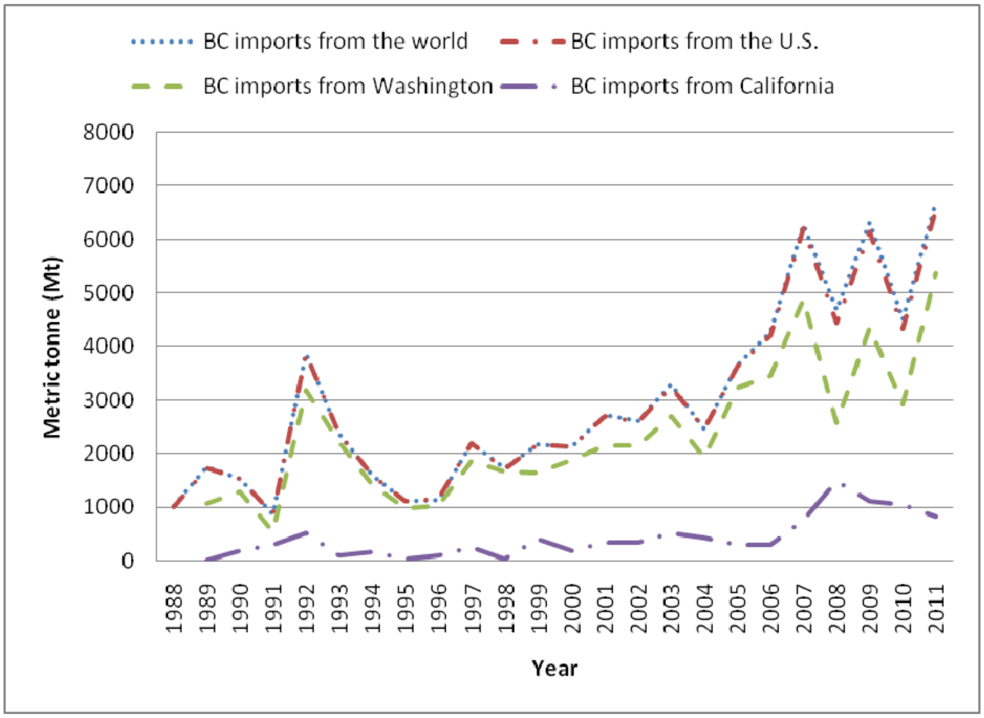

Figure 2. BC fresh cherry imports from the world, U.S. Washington and California.

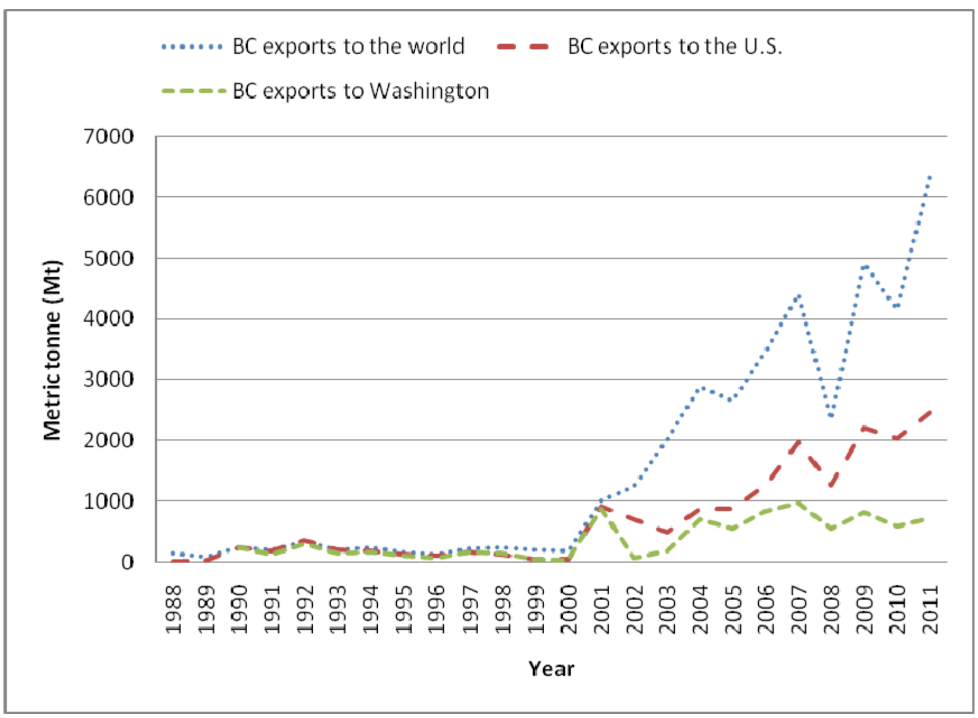

Figure 3. BC fresh cherry exports to the world, U.S. and Washington.

METHODOLOGY

Co-integration techniques including the Engle and Granger (1987) and Johansen's (1995) maximum likelihood procedures are employed to determine whether cherry markets are integrated in BC, Washington and California. The error-correction model, which involves estimating a long-run co-integrating relationship has been applied to study the integration of commodity agricultural markets and energy industries throughout the world (Barrett,2001;Wårell, 2006; Petersen and Muldoon, 2007;Mela and Canali, 2012). The Engle and Granger (1987) co-integration test involves two steps. The first step involves estimating equation (1) by OLS in order to determine whether the two price series exhibit a stable long-run relationship described as:

$$P_t^{bc} = \alpha + \beta P_t^{us} + \varepsilon_t \quad (1)$$

where P_t^{bc} and P_t^{us} are BC and U.S. (Washington, California) cherry prices, respectively, that are nonstationary I(1), α=constant term that accounts for factors such as transportation costs or product quality differences that may explain price variations, β denotes the long term relationship between BC and U.S. prices, and ε_t= error term. If β =1, relative prices are constant. Even though the two price series (P_t^{bc} , P_t^{us}) are nonstationary, a linear combination of them can be stationary, I(0).

The second step of the Engle Granger procedure requires testing the stationarity of the residuals of equation (1) by means of the standard Augmented Dickey Fuller (ADF) test. When the null hypothesis of no co-integration is rejected, the error correction model can be estimated by OLS employing the residuals from the long-run equation depicted as equation (1).

Once it has been established that a linear combination of the price series (P_t^{bc} , P_t^{us}) is stationary, the price series are said to be co-integrated (Engle and Granger, 1987). The existence of a co-integration relationship between two or more variables implies the existence of an error correction model which describes the long-run relationship of prices that is consistent with the short run price association(Verbeek, 2012). The error correction model is described as:

$$\Delta P_t^{bc} = \alpha + \delta \sum_{i=0}^{T} \Delta P_{t-i}^{us} + \beta \sum_{i=1}^{T} \Delta P_{t-i}^{bc} + \gamma \epsilon_{t-1} + \mu_t \quad (2)$$

where δ and β = the short run adjustment and γ =error-correction or speed of adjustment term.

Since price leadership was suggested by Flaming, Marsh, and Wahl (2007) as a topic for future research, Granger (1969) causality tests were employed between cherry prices in BC, Washington and California. Granger causality test measures whether prices in BC can be explained by past prices in U.S. markets, in addition to previous price information from the BC market. This is described as:

$$P_t^{bc} = \alpha + \sum_{i=1}^{p} \beta_i P_{t-1}^{bc} + \sum_{i=1}^{p} \delta_i P_{t-1}^{us} + \varepsilon_t \quad (3)$$

where the null hypothesis that P_t^{us} does not cause P_t^{bc} can be tested. If the U.S. price (P_t^{us}) causes the BC price (P_t^{bc}) with no evidence of causality in the other direction, this indicates evidence of price leadership exhibited by U.S. cherry producers and marketers.

ESTIMATION STRATEGY

(i) Unit Root Tests/Cointegration Analysis

In the analysis, cherry prices are expressed in natural logarithmic form. The estimation strategy involved first employing the Augmented Dickey-Fuller (Dickey and Fuller, 1979) test on prices to determine whether cherry prices are stationary or not. The standard ADF test was undertaken following the procedure specified by Mahadeva and Robinson (2004) as:

$$\Delta Y_t = \mu_0 + \alpha Y_{t-1} + \beta T + \sum_{i=1}^{i=T} \mu_i \, \Delta Y_{t-i} + \epsilon_t \quad (4)$$

where Δ is the difference operator, Y_t is the price of cherries at time t, T is a time trend variable that captures the deterministic growth in the time series over time. The Akaike Information Criterion (AIC) was used to select the optimal number of lags[two lags for California; five lags for BC and Washington] for the ADF test.

Once it was determined that the three price series were I(1), the second step of the analysis involved the Engle and Granger (1987) and Johansen (1991, 1995) cointegration analysis tests. For the cointegration approach, two price series form a long-run relationship and a linear combination of the data series will produce a residual series that is stationary (Engle and Granger, 1987). The Engle-Granger (1987) tests use a standard OLS estimation approach to analyze movements in prices in different geographic regions to ascertain whether there is a long run relationship. Unlike the Engle-Granger (1987) tests that fails to identify the number of co-integrating vectors in a multivariate framework, the Johansen test models the price relationships in a vector autoregressive (VAR) system. Consider a VAR of order k described as:

$$y_t = \Pi_1 y_{t-1} + \cdots + \Pi_k y_{t-k} + \gamma + \varepsilon_t \quad (5)$$

where y_t is a vector containing three non-stationary I(1) price variables, Π is a N x N matrix of parameters, and γ is a constant term. The VAR system of equations in (5) can be written in error correction form as

$$\Delta y_t = \sum_{i=1}^{k-1} \Gamma_i y_{t-1} + \Gamma_k y_{t-k} + \gamma + \varepsilon_t \quad (6)$$

where Γ_k is the long-run solution to equation (6). If Δy_t is a vector of I(1) variables, the left hand side and the first k-1 variables on the right hand side of equation (6) are stationary, and as a result the error term by assumption is stationary or I(0). Either y_t contains a number of co-integrating vectors or Γ_k must be a matrix of zeros. The rank of Γ_k , denoted by r, determines how many linear combinations of y_t are stationary. If r=N, then the variables are stationary in logs; if r=0, there exists no linear combinations of the variables that are stationary; and if $0<r<N$ there are r stationary linear combinations of y_t. The Johansen method tests for the number of co-integrating ranks with two asymptotically equivalent tests for co-integration: the trace test and the maximum eigenvalue test. The computed statistic for both tests is used to test the null hypothesis that there is no co-integrating vector (r=0). The Johansen method also allows testing of linear restrictions on the co-integrating vector by employing the likelihood ratio test (Johansen and Juselius, 1990).

DATA DESCRIPTION AND SOURCES

Annual farm-level price and quantities of BC, Washington and California fresh cherries are employed for the period 1971-2011 (Shinder, 2012; USDA, 2012a, 2012b). With the exception of a few studies (Toivonen, Toppinen, Tilli, 2002) the majority of market integration studies have employed monthly trade data. BC annual farm-level cherry prices are calculated by dividing the farm-gate value by the production of reported survey respondents for the crop year, May-November (Shinder, 2012). In contrast, Washington and California cherry farm prices are based on USDA surveys of warehouses which pack and ship cherries, and who report the volume of cherries packed during the season and the overall average price received (Ross, 2012).

Farm-level cherry prices in BC are converted to U.S. dollars using the Bank of Canada average annual exchange rates (Smith, 2010; Bank of Canada, 2012). While BC cherry prices are deflated by the BC producer price index for farm products (Statistics Canada, 2012c), Washington and California prices in U.S. dollars are deflated by the U.S. producer price index for farm products (U.S. Department of Labor, 2012). Figures 4 and 5 sketch the real cherry price levels and differentials.

Figure 4 shows the trend behavior of real cherry prices over the last four decades. BC prices exceeded Washington and California prices in the 1970s and in the late 2000s (Figure 5); a period characterized by an appreciated Canadian currency relative to the U.S. dollar. In contrast, California prices were higher than Washington prices primarily in the 1970s. Several reasons may explain the variation in cherry prices between the three markets.

Apart from product characteristics differences such as variety mix, quality, grade, and fruit size, production costs variations are likely to explain some of the variations in farm-level prices. Because Washington orchards, in general, have higher yields than California, their relative unit costs are likely to be lower than California and, therefore growers can market their cherries at lower prices than California (Grant, 2012).

The data were compiled from several sources. Per capita personal disposable income and population for Washington and California were obtained from the U.S. Department of Commerce (2012a), while BC personal disposable income, producer price index, consumer price index, and population were obtained from Statistics Canada (2012c). U.S. per capita personal disposable income was deflated by the U.S. Implicit Price Deflator (2005=100) (U.S. Department of Commerce, 2012b), whereas BC personal disposable income was deflated by the Consumer Price Index (all items) (Statistics Canada, 2012c). Table 1shows the variable definitions and Table 2 provides summary statistics of the variables.

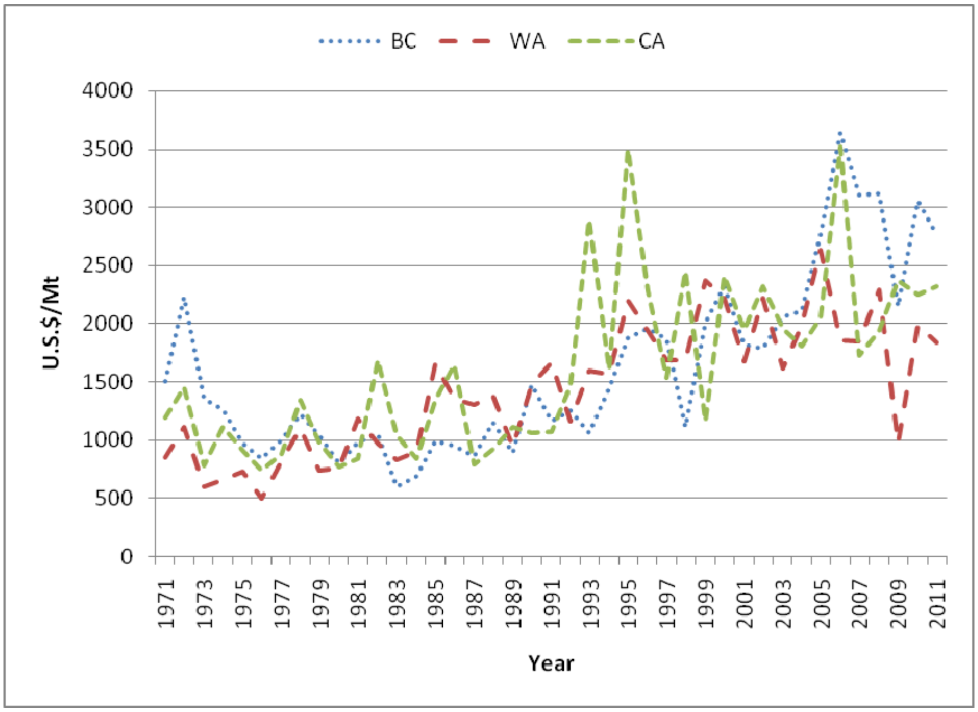

Figure 4. British Columbia, Washington and California real fresh sweet cherry prices.

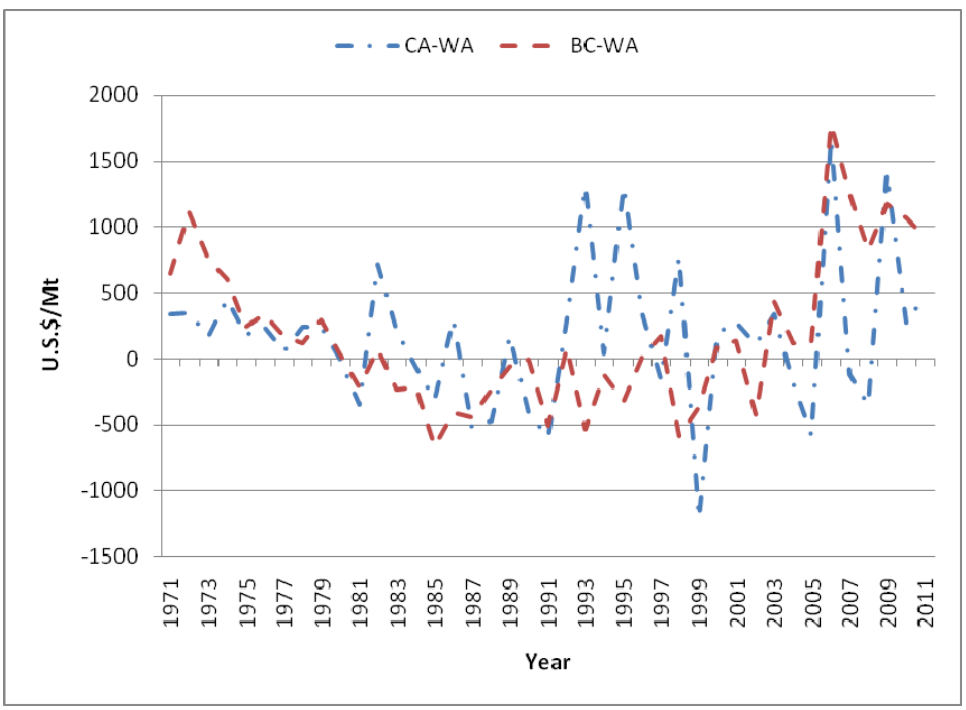

Figure 5. Price differential of California-Washington and BC-Washington real fresh sweet cherry prices.

Table 1. Variable names, definitions and units of measurement

Variable	Variable definition	Unit
P_{bc}	Real farm-level price of BC fresh sweet cherry	\$/mt
P_{wa}	Real farm-level price of Washington fresh sweet cherry	\$/mt
P_{ca}	Real farm-level price of California fresh sweet cherry	\$/mt
CQ_{bc}	Per capita BC fresh sweet cherry production	Kg
CQ_{wa}	Per capita Washington fresh sweet cherry production	Kg
CQ_{ca}	Per capita California fresh sweet cherry production	Kg
SQ_{bc}	Per capita BC fresh strawberry production	Kg
SQ_{wa}	Per capita Washington fresh strawberry production.	Kg
SQ_{ca}	Per capita California fresh strawberry production	Kg
IN_{bc}	Real per capita personal disposable income in BC	\$
IN_{wa}	Real per capita personal disposable income in Washington	\$
IN_{ca}	Real per capita personal disposable income in California	\$
POP_{bc}	Population in British Columbia	Mil.
POP_{wa}	Population in Washington	Mil.
POP_{ca}	Population in California	Mil.
PPI_{us}	U.S. producer price index for farm products (base, 1982=100)	
PPI_{bc}	BC producer price index for farm products (base 1997=100)	
IPD	U.S. implicit price deflator (base, 2005=100)	
CPI	BC consumer price index (base, 2002=100)	
CD/US	Canadian/U.S. exchange rate	
DU	Dummy variable (equal to one since 1988, otherwise zero)	

Note: All prices are in U.S. dollars.

Table 2. Descriptive statistics of real farm-level fresh sweet cherry prices, per capita farm-level fresh cherry quantity, per capita farm-level fresh strawberry quantity, real per capita disposable income, and population in British Columba, Washington and California, 1971-2011

Variable	British Columbia (BC)		Washington (WA)		California (CA)	
	Mean	SD	Mean	SD	Mean	SD
Real price of fresh cherries (U.S.$/mt)[1]	1616	770	1438	560	1613	728
Differential: BC-WA (U.S.$/mt	178	570	–	–	–	–
Differential: CA-WA (U.S.$/mt)	–	–	–	–	175	566
Per capita fresh cherry quantity (kg)	1.41	0.67	10.63	4.43	0.93	0.45
Per capita fresh strawberry quantity (kg)	0.47	0.18	0.26	0.13	12.41	5.97
Real per capita disposable income (U.S.$)	17,417	3,458	25,478	6,576	26,412	5,526
Population (mil.)	3.41	0.7	5.05	1.1	29.41	5.6

Note: [1]Coefficient of variation (measure of price dispersion) of real cherry prices varied over cherry markets from 0.48 for BC to 0.39 for WA and 0.45 for CA.

EMPIRICAL RESULTS

(i) Cointegration Analysis

Employing the ADF test for cherry price series in logarithmic form indicates the null hypothesis of non-stationarity could not be rejected at the 5% level and therefore BC, Washington and California cherry prices are non-stationary and are integrated of order one (I(1).The ADF test on the residual of the Engle-Granger (1987) regression (for example: regression of BC on Washington) shows that the null hypothesis of no cointegration is rejected between BC and Washington prices and between Washington and California prices. These results indicate that BC, Washington and California cherry prices share a long-run relationship and belong to the same market. Regressing BC on Washington prices was decided based on the results of the Granger (1969) causality tests. For example, if causality runs from Washington to BC, then BC prices are regressed on Washington prices. The residuals of the following regressions had no unit root and hence were stationary.

The results of the Johansen ML trace test statistics and maximum eigenvalue statistics confirm the results of the Engle-Granger (1987) cointegration tests that BC, Washington and

California prices are cointegrated at the 5% level (Table 3). Both the trace and the maximum eigenvalue test indicate one cointegrating vector at the 5% level. As a result, cherry prices share a common stochastic trend and, consequently, the BC, Washington and California prices are cointegrated and hence spatial markets are integrated.

Table 3. Johansen and Engle and Granger Co-integration Test

BC-WA (Johansen)	Max Rank	Log Likelihood	Max.eigenvalue	Trace Statistics	5% Critical Value
	0	16.995413	.	20.3170*	15.41
	1	24.803099	0.33697	4.7017**	3.76
	2	27.153925	0.11638		
BC-WA (Engle/Gr)	ADF-test statistic				
	-3.489				-2.958
BC-CA (Johansen)	Max Rank	Log Likelihood	Max.eigenvalue	Trace Statistics	5% Critical Value
	0	-26.732689	.	34.4821*	15.41
	1	-10.829017	0.54850	2.6748	3.76
	2	-9.4916321	0.06468		
BC-CA (Engle/Gr)	ADF-test statistic				
	-6.713				-2.958
WA-CA (Johansen)	Max Rank	Log Likelihood	Max.eigenvalue	Trace Statistics	5% Critical Value
	0	-33.154513	.	36.1912*	15.41
	1	-17.788446	0.53620	5.4591**	3.76
	2	-15.058909	0.12757		
WA-CA (Engle/Gr)	ADF-test statistic				
	-6.057				-2.958

Note: BC =British Columbia; WA= Washington, and CA=California.
*Indicates significance at the 1% level; **Indicates significance at the 5% level.

Table 4. Result of Granger Causality test

	Direction of Causality	Chi^2	p-value
1971-2011	BC→ WA	7.06	0.070
	WA→ BC	8.98*	0.030
	BC→ CA	19.22*	0.000
	CA→ BC	0.388	0.534
	WA→ CA	14.37*	0.000
	CA→ WA	1.27	0.259

Notes: BC=British Columbia, WA=Washington, and CA=California.
* Indicate significance at 5%.

The Granger (1969) causality tests conducted to determine whether there was a causality relationship between current BC prices and Washington or California prices indicate the direction of price causality for some markets is counterintuitive (Table 4). For example, BC prices are said to Granger-cause California prices if they lead California prices. Whether these counterintuitive results may be due to the structure of the cherry market is difficult to conclude from the limited statistical analysis. It may be argued that since BC is a major importer of cherries from the U.S., the BC prices may have a major influence on farm-level cherry prices in California. Because Washington has a relatively larger production share of the cherry market they may be exerting some 'price leadership' and are said to Granger-cause BC prices as well as California prices (Table 4). O'Rourke (2012) argues institutional factors may have caused these counterintuitive results, and may be a combination of expectations about the size of the Pacific Northwest crop that might have affected California prices coupled with the number of generic promotions that retailers would have planned with California shippers.

(ii) Error-Correction Model (ECM) Analysis

ECM equations were estimated for three pairs of prices (BC-Washington, BC-California, and California-Washington). The ECM controls for a long run and a short-run relationship in cherry prices. The residuals of the Engle-Granger co-integration relationship were used as the error correction term in the ECM equation. Up to three lags were used in the ECM analysis. The number of lags for the ECM equations was determined by a vector autoregressive for each pair using the Bayesian Information Criterion (BIC).The error correction model and the speed of price transmission results are shown in Table 5.

BC-Washington (3 lags)

The causality runs from Washington (W) to BC (B).

$$\Delta B_t = D + \alpha_1 e_{t-1} + \alpha_2 \Delta B_{t-1} + \alpha_3 \Delta B_{t-2} + \alpha_4 \Delta W_t + \alpha_5 \Delta W_{t-1} + \alpha_6 \Delta W_{t-2} \quad (7)$$

$$\Delta B_t = D + \beta_1 e_{t-1}^{-} + \beta_2 e_{t-1}^{+} + \beta_3 \Delta B_{t-1} + \beta_4 \Delta B_{t-2} + \beta_5 \Delta W_t + \beta_6 \Delta W_{t-1} + \beta_7 \Delta W_{t-2} \quad (8)$$

The ECM coefficients indicate the speed of adjustment from Washington (W) to BC (B) is not significant even at the 10% level. The coefficient is insignificant even when controlled for asymmetry. However, when $\alpha_4 \Delta W_t$ is dropped from equation (7) (i.e., control for the contemporaneous effect) the coefficient (speed of adjustment) becomes significant (-0.31) at 10% level. This implies the convergence of the value to its long term shared path with the other variable.

The short run adjustment coefficients (i.e.,α_4, $\alpha_5$and$\alpha_6$in equation7) are all significant (0.73, 0.63, and 0.85, respectively), indicating that changes in Washington cherry prices from one period to the next have a significant impact on BC cherry prices from one period to the next for up to three periods. Results indicate that a 1% change in Washington prices from the previous period leads to a 0.73% (the value of α_4) change in BC prices. The result indicates that BC prices are not insulated from Washington prices and, therefore, adjust to long run disequilibrium changes in the prices of Washington cherries.

Table 5. The Error correction model and speed of transmission

		Co-integrated?			Error Correction Model						Asymmetry of adjustment		
	Unit Root? (5%) level	Johansen Test (5%)			Speed of Adjustment			Short-run adjustment			+/-	+/-	+/-
	ADF	BC	WA	CA	BC	WA	CA	BC	WA	CA	BC	WA	CA
BC	Yes	-	Yes	Yes	-	N/A	-1.070*	-	N/A	d(0.47)*	-	N/A	-1.23/-0.82*
WA	Yes	Yes	-	Yes	-0.057	-	0.92*	d(0.73) ld(0.63) l2d(0.85)	-	d(0.36)*	0.11/-0.23	-	-1.05/-0.80*
CA	Yes	Yes	Yes	-	N/A	N/A	-	d(0.12) ld(0.00)	-	-	N/A	N/A	-

Note: BC=British Columbia, WA=Washington, and CA=California.
* Indicate significance at 5% level.
Transmission runs from row to column.

BC-California (1 lag)

The causality runs from BC (B) to California (C)

$$\Delta C_t = D + \alpha_1 e_{t-1} + \alpha_2 \Delta B_t \tag{9}$$

$$\Delta C_t = D + \beta_1 e_{t-1}^{-} + \beta_2 e_{t-1}^{+} + \beta_3 \Delta B_t \tag{10}$$

Price transmission is from BC to California and is significant even at 1%. The speed of transmission is high (at >-1), which may indicate that it is instantaneous[1]. The asymmetry test indicates that price increases in BC are transmitted more slowly than price decreases (-0.82 vs -1.23). However, statistically the two figures are the same and, therefore, overall we conclude that the transmission is symmetric. The short run adjustment (i.e., coefficient on ΔB_t in equation 9) is 0.47 and significant, which indicates that a 1% change in the BC price from one period to the next leads to 0.47% change in the price of California cherry.

California-Washington (1 lag)

The causality runs from Washington (W) to California (C)

$$\Delta C_t = D + \alpha_1 e_{t-1} + \alpha_2 \Delta W_t \tag{11}$$

$$\Delta C_t = D + \beta_1 e_{t-1}^{-} + \beta_2 e_{t-1}^{+} + \beta_3 \Delta W_t. \tag{12}$$

The coefficient for the speed of transmission from Washington to California is significant and indicates that convergence to the long term path is fast (at α_1 = -0.92 in equation 11). The asymmetry test indicates that price increases are transmitted slower (-0.80 vs -1.05 for β_1 *and* β_2 in equation (12). However, as stated earlier, the two figures are statistically the same indicating no asymmetry. The short run adjustment is 0.36 and significant (0.36 is the value of α_2 in equation 11).

In summary, the three price series of BC, Washington and California cherry markets are cointegrated. The coefficient on the error correction model of the BC-Washington analysis indicates that it is insignificant, illustrating that there is no convergence to a long run common path. However, the short run transmission is strong and significant, as demonstrated by the short run adjustment coefficients, but it's difficult to say if the series share a long term common path.

(iii) Price Formation Analysis

Employing the neo-classical demand theory, an analysis of price-quantity relationships is developed to show how cherry quantities in one market and cherry quantities from adjacent markets affect cherry prices. A double logarithmic inverse demand model with spatial separate markets is described below:

$$\Delta lnP_i = \alpha_0 + \Delta\alpha_1 \ln(Q_i) + \Delta\alpha_2 \ln(Q_i) + \Delta\alpha_3 \ln(Q_k) + \Delta\alpha_4 \ln(M_i) + e_t \tag{13}$$

[1] Coefficient larger than -1 does not make interpretive sense, so it should be treated as close to -1.

Where lnP_i=natural logarithm of real farm-level cherry prices in market i, lnQ_i=natural logarithm of per capita farm-level cherry quantity in market i, lnQ_j=natural logarithm of per capita farm-level cherry quantity from market j, lnQ_j=natural logarithm of per capita farm-level quantity of substitute fruit (strawberry) in market i, and lnM_i=natural logarithm of real per capita disposable income in market i.

The variables above are tested for unit roots using the Augmented Dickey Fuller (ADF) test. The ADF test on the converted series indicate that none of the variables are stationary but have a unit root and are integrated of order one I(1). In order to convert the variables into a stationary format, they were differenced. The differenced variables are stationary and confirmed to be I(0).

The models were tested for heteroskedasticity and multicollinearity**.** The variance inflation factor (VIF) was used to test for multicollinearity. The highest VIF is 5.85 and most are 3 or less indicating a very low probability of multicollinearity in the model specifications. The Breusch-Pagan test for heteroskedasticity shows that the residual of the models has constant variance. The OLS regression results for the five models are shown in Table 6.

Table 6. Price formation regression results for BC, Washington and California, 1971-2011

Variables	BC[1] (Model 1)	BC[2] (Model 2)	Washington[2] (Model 3)	California[2] (Model 4)
Constant	-0.0361 (-0.54)	0.0117 (0.20)	0.0450 (0.92)	0.0563 (1.18)
$Log(CQ_{bc})$	-0.2400* (-2.86)	-0.2520* (-2.86)	-	-
$Log\ (CQ_{wa})$	-0.3491* (-3.59)	-0.3778* (-3.67)	-0.7066* (-8.03)	-0.1339 (-1.39)
$Log(CQ_{ca})$	0.0014 (0.03)	-0.0009 (-0.02)	-0.0889*** (-1.87)	-0.5375* (-10.55)
$Log(SQ_{bc})$	-0.2077 (-1.56)	-0.2362*** (-1.68)	-	-
$Log(SQ_{wa})$	-		-0.1162 (-1.02)	-
$Log(SQ_{ca})$	-			0.4518 (1.14)
$Log(IN_{bc})$	1.1694 (0.83)	-0.2310 (-0.37)	-	-
$Log(IN_{wa})$	-	-	-0.0648 (-0.04)	-
$Log(IN_{ca})$	-	-		-2.7907 (-1.52)
Log(CD/US)	0.6118 (0.80)	-		
DU	0.0744 (0.96)	0.0372 (0.50)		

Notes: [1] Variables (prices, income) are denoted in Canadian dollars. [2] Variables (e.g., prices, income) are denoted in U.S. dollars. Numbers in parentheses are t-values. *Significant at the 1% level; ***Significant at the 10% level.

The dummy variable controlling for the pre- and post-free trade era is not significant in any of the BC models. The variable is not significant even at the 10% level. Therefore, the price of cherries at the farm level in BC is not affected by the implementation of CUSFTA in 1989. The results of estimating the second version of the BC model (Table 6) indicate that per capita farm-level quantity of cherry produced in BC and Washington are significant variables in explaining the price of cherries in BC. The per capita production quantity of strawberries is not significant at 5%, however it is significant at 10%. Real income per capita and Canadian/US exchange rate were included as explanatory variables for different versions of the BC cherry model. However, real income per capita and the exchange rate are not significant variables. For the Washington cherry model, the price of Washington farm-level cherries is explained primarily by own-production cherries in Washington. The price of California cherries is explained also only by farm-level per capita production quantity in California and none of the other variables are significant. Alternative model specifications (e.g., non-differenced income variable) were explored in the Washington and California models resulting in real per capita disposable income having a significant effect on cherry prices. In summary, a 1% increase in difference in logs of quantity of BC cherries leads to decline of 0.22% to 0.25% in difference in logs of cherry prices. The numbers are between 0.35% and 0.38% for production quantities in Washington.

Discussion and Implications

Fresh sweet cherry production and trade has expanded in the U.S. and Canada over the last two decades. The steady demand growth has been driven by the popularity of sweet cherries, productivity growth and the CUSFTA (later NAFTA), which liberalized trade between the U.S. and Canada. The fresh sweet cherry trade is concentrated between the states of California and Washington and the province of British Columbia in the western North America. Canada's cherry import value from the U.S. increased 13-fold between 1988 and 2011 and about 17-fold to BC. The changes in the sweet cherry industry justified the examination of producer price relationship between cherry markets of each region and investigation of factors that impact price formation. The study updates the previous U.S. studies on fresh sweet cherry market demand and, more generally, contributes to the market integration literature by considering a semi-perishable commodity that is produced primarily in the Pacific Northwest. The study uses farm-level prices, production quantities and provincial trade statistics, unavailable prior to 1988 (the introduction of CUSFTA).

Overall, the plots indicate an acceleration of trade in fresh cherries between BC and Washington since the adoption of NAFTA, while the price plots in BC have been often higher than in other U.S. states. Price differences between cherry markets are attributed to the variety mix, among others, which affects the sequencing of the harvests, with early maturing varieties dominating California industry and late-maturing varieties being important in BC.

Cointegration techniques and vector error correction models are employed in the analysis of three cherry price series over the period 1971-2011. Our results indicate the three producer prices are cointegrated and therefore cherry markets are integrated. The ECM results show that Washington cherry prices have a significant and instantaneous effect on BC cherry prices.

There is also some indication that prices of both regions share the long term path. The effect of BC cherry prices on California cherry prices indicate the stronger dampening effect of BC prices, which suggests that California growers are less likely to expand production by planting the late-maturing varieties.

The change in the BC prices appears to have an effect in the following harvest season in California (0.47% change) and could reflect the specificity of the cherry market where the demand remains strong over time and consistent with the perceived health benefits associated with cherry consumption. California prices also adjust instantaneously in response to changes in Washington cherry prices. We cannot decisively conclude that the prices across the three regions converge to a long run equilibrium.

The investigation of factors influential for price formation shows that cherry production in each production region negatively influences the region's cherry prices and the larger the region's production share, the larger the decrease in cherry prices (ranging from 0.71% in Washington to 0.24 % in BC).

A 1% increase in the quantity supplied (produced because of the perishable nature of cherries) in Washington lowered prices in BC by 0.35-0.38% depending on the currency in which values are expressed, while Washington cherry prices are lowered by about 0.09% in response to a 1% California production increase. Interestingly, an increase of 1% in BC strawberry production tends to lower the BC cherry prices expressed in U.S. currency by about 0.24%. Such effect was not confirmed in the case of strawberry supply in Washington or, especially, California, a relatively large strawberry producer.

Cherry production in the three regions will continue to expand in response to prices. while the production in BC and the region's cherry imports will also continue to increase in response to the growing price differentials. BC growers will likely continue the dual marketing strategy to supply consumers within the province and North America, but also in importing European and Asian countries to exploit greater earning opportunities arising from segmenting the market and managing the increasing supply of cherries. In years of crop shortage, prices are likely to increase, but the ultimate price increases in each region will depend on the geographical scope of supply disturbances. Localized supply shortages may be alleviated by the supply from the neighboring area although the timing of potential deliveries will depend on the harvest timing resulting from the variety maturation.

Despite the increased trade in fruits and vegetables since the adoption of CUSFTA/NAFTA, it has not significantly influenced cherry prices or, at least, the applied specification has not supported such effect. It is possible that data of higher frequency than annual data could help in identifying a different pattern. Also, the per capita income has not influenced the prices of cherries, but cherries represent a rather small share of the consumed fruit or food overall.

Higher frequency data could establish a different effect. However, given the seasonal nature of cherry production and consumption, high frequency data can become available in the future if cherries imported from the Southern Hemisphere assure their availability to consumers. The increasing popularity of sweet cherries and expanding availability in regions that do not produce them will eventually create new and adequate data applicable in generating a more in-depth analysis than the current study.

References

Agriculture and Agri-Food Canada. 2011. An Overview of the Canadian Agriculture and Agri-Food System.Ottawa, Ontario,p. 27.

Bank of Canada. 2012. Annual Exchange Rate. Available online at:http://www.bankofcanada.ca/rates/exchange/exchange-rates-in-pdf/ (Accessed July 18, 2012)

Barrett, C. B. 2001. Measuring integration and efficiency in International Agricultural Markets. *Review of Agricultural Economics*, 23(1): 19-32.

Boriss, H., H. Brunke, , and M. Kreith, 2011. Cherry Profile. Agricultural Marketing and Resource Centre, Agricultural Issues Centre, University of California (updated by Geisler, M; AgMRC, Iowa State University). Available online at: http://www.agmrc.org/commodities_products/fruits/cherry profile.cfm. (Accessed July 19, 2012.

Dickey, D. A. and W. A. Fuller, 1979. Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, 74: 427-31.

Engel, C. and J. H. Rogers. 1996. How wide is the border? *American Economic Review*, 86: 1112-1125.

Engle, R. F. and C. W. J. Granger 1987. Co-integration and error correction: representation, estimation and testing. *Econometrica*, 55: 251-76.

Federico, G. 2010. When did European markets integrate? European Review of Economic History, 15: 93-126.

Flaming, S., T. L. Marsh, and T. Wahl, . 2007. Farm-level price formation for fresh sweet cherries. *Journal of Food Distribution Research*. Available online at http://ageconsearch.umn.edu/bitstream/43495/2/3802fr43.pdf, 39-49. Accessed April 21, 2009).

Gellatly, G. and B. Yan. 2012. New Evidence on exchange-rate adjusted prices in Canada. Economic Analysis Division, Statistics Canada. Catalogue No. 11-626-X-No.002.

Ghoshray, A. 2007. An examination of the relationship between U.S. and Canadian durum wheat prices. *Canadian Journal of Agricultural Economics*, 55: 49-62.

Goldberg, P. K. and F. Verboven. 2005. Market integration and convergence to the Law of One Price: evidence from the European car market. *Journal of International Economics*, 65: 49-73.

Granger, C. W.J. 1969. Investigating causal relations by econometric models and cross-spectral methods. *Econometrica* 37: 424-438.

Grant, J. 2009. Sweet cherry production and marketing. 22nd California Small Farm Conference,Sacramento, California, March 1.

Grant, J. 2012. UC. Cooperative Extension. Stockton, California. Personal communication. July 17, 2012.

Hansen, M. 2010. Cherry industries collaborate. *Good Fruit Grower*. Available online at http://www.goodfruit.com/core/pagetools.php?pageid=13513andurl=%2FGood-Fruit-Grower%2FAugust. Accessed October 15, 2010.

Johansen, S., and K. Juselius (1990). Maximum likelihood estimation and inference on co-integration with applications to the demand for money. *Oxford Bulletin of Economics and Statistics,* 52: 169-210.

Johansen, S. 1991. Estimation and hypothesis testing of co-integration vectors in Gaussian autoregressive models. *Econometrica* 59:1551-1580.

Johansen, S. 1995. Likelihood Based Inference in Cointegrated Vector Autoregressive Models. Clarendon Press. Oxford University.

Kahlke, C. J., O. I. Padilla-Zakour, H. J. Cooley, T. L. Robinson. 2009. Shelf-life and marketing window extension in sweet cherries by the use of modified atmosphere packaging. *New York Fruit Quarterly*, 17(2): 21-24.

Kappel, F. 2006. Sweet cherry revolution in British Columbia, Canada. *ChronicaHorticulturae* 46(4): 17-20.

Mahadeva, L., and P. Robinson. 2004. Unit Root Testing to Help Model Building, Handbook in Central Banking, Centre for Central Banking Studies, Bank of England, July, No.22.

McCracken, V. A., K. L. Casavant, D. C. Miller. 1989. Export demand for high-value products - A case study of sweet cherries in Japan. *Agribusiness*, 5(1): 1-8.

Mela, G., and G. Canali. 2012. EU and World Agricultural Markets: Are They more Integrated after the Fischler Reform? Paper prepared for the 123rd EAAE Seminar "Price Volatality and Farm Income Stabilisation:Modelling Outcomes and Assessing Market and Policy Based Responses. Dublin, February 23-24, 2012.

Miljkovic, D. 2009. U.S. and Canadian livestock prices: market integration and trade dependence. *Applied Economics*, 41: 183-193

O'Rourke, A. D.,and K. L. Casavant. 1974. Marketing Pacific Northwest sweet cherries. Washington State University, College of Agriculture Research Centre, Bulletin 800.

O'Rourke, D. 2012. Belrose Inc. Personal communication. July 16, 2012.

Okanagan Tree Fruit Authority. 2010. Special tabulation request of new sweet cherry variety plantings in British Columbia. Summerland, British Columbia.

Perez, A., and S. Pollack. 2007. Bigger crops forecast for most stone fruit and almonds in 2007 than a year ago. Fruit and Tree Nuts Outlook. United States Department of Agriculture. FTS-328.

Perez, A., and S. Pollack. 2008. Lower U.S. production forecast for most stone fruit in 2008. Fruit and Tree Nuts Outlook, United States Department of Agriculture. FTS-333.

Perez, A., and K. Plattner. 2012. Lower supplies aid early summer fruit prices. Fruit and Tree Nuts Outlook, United States Department of Agriculture. FTS-352.

Petersen, E. H. and G. Muldoon. 2007. "Wholesale and retail price integration in the live reef-fish food trade". In B. Johnston (ed), Economics and Market Analysis of the Live Reef-Fish Trade in the Asia-Pacific Region, pp. 61-75. ACIAR Working Paper No. 63, Canberra, Australia.

Pollack, S., and A. Perez. 2004. Smaller supply of citrus this summer, but more sweet cherries and blueberries. Fruit and Tree Nuts Outlook, United States Department of Agriculture. FTS-311.

Ross, J. 2012. U.S. Department of Agriculture, National Agricultural Statistics Office. Personal communication. July 17, 2012.

Schotzko, R. T., W. W. Wilson, and D. Swanson. 1989. Demand for Washington fresh sweet cherries. Washington State University, Research Bulletin No. XB1007, Pullman, Washington.

Shinder, L. 2012. Statistics Canada. Tabulation Data Request (BC Fresh Berry Statistics). Personal communication. June 18, 2012.

Smith, N. 2010. Historical Canada/U.S. Exchange Rate. Bank of Canada. Personal communication. May 5, 2012.

Statistics Canada. 2012a. International Merchandise Trade Database. http://www5.statcan.gc.ca/cimt-cicm/topNCountryCommodities-marchandises?countryId=999andtradeType=1andusaState=andtopNDefault=10andfreq=6andcommodityName=Cherries%2C+freshandlang=engandrefYr=2010andsectionId=andchapterId=8andarrayId=9800000andprovId=1andrefMonth=5andcommodityId=80920 (Accessed July 18, 2012).

Statistics Canada. 2012b. Fruit and Vegetable Production. Catalogue No. 22-003-X.

Statistics Canada. 2012c. Personal Disposable Income (Table 384-0013), Population (Table 051-0001), Consumer Price Index, All Items (Table 326-0021), and BC producer price index (Table 002-0022). Available online at:http://www5.statcan.gc.ca/cansim/home-accueil?lang=eng (Accessed July 18, 2012.

Toivonen, R., A. Toppinen and T. Tilli. 2002. Integration of roundwood markets in Austria, Finland and Sweden. *Forest Policy and Economics*, 4: 33-42.

U.S. Department of Agriculture. (2007). Foreign Agricultural Service. World Markets and Trade, pp 1-3. Available online at:http://www.fas.usda.gov/htp/2007_Cherries.pdf. Accessed April 22, 2009.

U.S. Department of Agriculture. 2011. National Agricultural Statistics Service. Washington Tree Fruit Acreage Report 2011. Olympia, Washington.,

U.S. Department of Agriculture. 2012a. National Agricultural Statistics ServiceNoncitrus Fruits and Nuts 2011 Preliminary and Annual Summary. Available online at:http://usda.mannlib.cornell.edu/MannUsda/viewDocumentInfo.do?documentID=1113 (Accessed July 18, 2012).

U.S. Department of Agriculture. 2012b. National Agricultural Statistics Service. Data and Statistics. Quick Stats 1.0. Historical data for strawberry by U.S. state (Washington and California). Available online at:http://www.nass.usda.gov/Quick_Stats/. (Accessed July 18, 2012).

U.S. Department of Commerce. 2012a. Bureau of Economic Analysis. Regional economic accounts: income and population data by state. Available online at:http://www.bea.gov/iTable/iTable.cfm?reqid=70andstep=1andisuri=1andacrdn=4(Accessed July 18,2012).

U.S. Department of Commerce. 2012b. Bureau of Economic Analysis. National Economic Accounts. Table 1.1.9. Implicit Price Deflator for Gross Domestic Product. Available online at:http://www.bea.gov/iTable/iTable.cfm?ReqID=9andstep=1 (Accessed July 18, 2012).

United States Department of Labor. 2012. Bureau of Labor Statistics. U.S. Producer Price Index for farm products. Available online at:http://data.bls.gov/cgi-bin/srgate. Accessed July 9 2012.

Verbeek, M. 2012. A Guide to Modern Econometrics, 4th Edition, Chichester, West Sussex: John Wiley and Sons.

Vollrath, T and C. Hallahan. 2006. Testing the integration of U.S.-Canadian meat and livestock markets. *Canadian Journal of Agricultural Economics* 54: 55-79.

Wårell, L. 2006. Market integration in the International Coal Industry: a cointegration approach. *The Energy Journal* 27(1): 99-118.

World Sweet Cherry Review. 2011. Washington: Belrose, Inc. Pullman.

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IMPACT OF TRADE LIBERALIZATION IN RICE: ASSESSING ALTERNATIVE PROPOSALS

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ABSTRACT

The objective of this paper is to analyze impacts of alternative trade reform proposals on the global rice economy. The impacts of these proposals are analyzed with respect to the three pillars of the Doha Development Agenda: improved market access, reduced domestic support, and termination of export subsidies. The Arkansas Global Rice Model, an econometric model of the global rice market is used to simulate the proposed reform scenarios.

Estimates of world reference prices and changes in trade volumes of major rice importing and exporting countries are reported. The results suggest that an increase in global rice trade is largely attributable to market access reforms rather than changes in domestic support. Elimination of export subsidies has no impact on global rice trade. The US reduction in domestic support is only partially compensated by increases in world prices over time. There is significant tariff reduction and tariff rate quota expansion only in the US proposal.

Keywords: Rice, Doha Development Agenda, Market Access, Domestic Support

JEL codes: F13, Q17, Q18

INTRODUCTION

Rice plays a key role in the food security in most Asian countries. About ninety percent of global rice production is in Asian countries and their rice sectors are subject to much government intervention through support prices and trade measures to protect domestic rice markets. As rice is a basic staple crop, most Asian countries justify their rice policies on the basis of food security and multi-functionality; concepts that have gained traction in the Doha Development Agenda negotiations (FAO, 2004).

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Agreements and reforms at national and international levels (Uruguay Round on Agriculture Agreement (URAA), and regional trade agreements) have led to increased rice trade (Wailes, 2004).

Assuming no change from current policies the USDA baseline in 2012 projected a 2.9 percent annual increase in global rice trade from 2012 to 2021, exceeding 45 million metric tons in 2021 (USDA, 2012). Based on Wailes' (2004) assessment since the mid 1990s rice trade has doubled in both volume and as a share of global consumption.

The global rice markets are distorted by importing countries using tariffs and tariff rate quotas (TRQ) to protect domestic producers, while exporting countries distort the global market using price and income supports to assist domestic producers (Wailes, 2004). Although the URAA achieved some reductions in protection, overall the tariff rates remained relatively high for rice. The rate of import tariffs for rice was among the highest of all at 73.34 percent as compared to wheat at 68.18 percent, sugar at 59.14 percent and other grains at 11.02 percent (Diao, Somwaru and Roe, 2001). The global unweighted average tariff bound rate for agricultural commodities was 62 percent, which is relatively high compared to other product classes (ERS, 2001). Likewise, based on Wailes' assessment in 2000, medium and long grain markets had average tariff rates of 217 percent and 21 percent respectively (Wailes, 2004).

Hence we conclude that a combination of high levels of geographic concentration, domestic protection, erratic weather effects, and inelastic price response in production and end use markets with a relatively thin volume of rice traded results in volatile prices and trade (Wailes, 2002).

Problem Statement

Negotiations in the Doha Development Agenda (DDA) have been long-running, similar to the URAA. Negotiators failed to meet the deadline of March 2003 for developing an explicit framework for reforms. Not until August 2004, was there a concrete framework agreed upon for the DDA (WTO, 2004). Since August 2004 negotiations have resulted in the submission of proposals from key nations and groups of nations. The present paper examines four proposals from individual nations and groups of nations that have been offered to extend the reform process within the Doha Development round.

The specific objectives of the study are to assess the proposed reforms for global rice trade as contained in alternative Doha round submissions for the global rice economy. The Arkansas Global Rice Model (AGRM) an existing global rice econometric simulation model is used to estimate supply and demand to simulate the US, EU, G20[1], and G10[2] proposals and to examine how world imports and exports are impacted by the alternative reform proposals. The goal of this study is to help policy makers and people involved in the international rice market to better understand and improve decision-making with regard to DDA negotiations.

[1]G20 – is a group of developing countries with China, India, Brazil, South Africa, and other developing countries. G20 group consists of countries that signed the alternative proposal to the EU and the US for Cancun Ministerial meeting on 20 August 2003.

[2] G10 – a group consist of Switzerland, Japan, South Korea, Taiwan, Norway, Iceland, Israel, Liechtenstein, Bulgaria, and Mauritius - major importers of agricultural commodities.

Characterization of the Global Rice Economy

A brief overview of global rice economy is discussed in this paragraph. Rice accounts for about 20 percent of global calories consumed (FAOSTAT). As a source of complex carbohydrates it is the dominant staple food crop in many Asian countries. Global rice production has increased faster than population growth over last three decades mainly due to technological progress (Wailes, 2002). As a staple food, the demand for rice is inelastic with respect to own price and income changes. In many Asian countries the traditional rice diet is being replaced by fruits, vegetables, and meat, making rice an inferior good with respect to income (Ito et al.).

The major types of rice traded worldwide are long, medium, and short grain rice. Long grain rice accounts for more than 75 percent of the global rice trade and is grown in tropical and subtropical regions (USDA, 2005). Aromatic or fragrant rice especially jasmine rice from Thailand and basmati from India and Pakistan, account for about 10 percent of the global trade and is sold at premium in world markets (USDA, 2005). In addition to classification of global rice trade by rice type (long and medium), it is also characterized by degree of processing (milled, brown, and paddy) and quality or percentage of broken kernels (Wailes, 2004). In 2005, the global rice trade accounted for 6.4 percent of the world's production (USDA, PSD 2006). Even though this level of rice trade relative to production has increased over time it is low compared to wheat trade at 18.5 percent, corn at 11.8 percent, and soybean at 29.9 percent (USDA, PSD 2006). The major long grain importing countries are Indonesia, Bangladesh, the Philippines and Malaysia. Imports in these countries are subject to production shocks as a result of variable weather conditions and natural calamities (monsoons and typhoons). Other major long grain, importing countries are Iran, Iraq, Saudi Arabia, Nigeria, Cote D'Ivoire, Senegal, South Africa, Brazil, China, and the EU. Similarly, major medium grain importing countries are Japan, South Korea, Taiwan and Turkey. The major long grain rice exporting countries are Thailand, Vietnam, India, Pakistan, the United States, Uruguay and Argentina. The major medium grain rice exporting countries are Egypt, China, Australia, Italy, and the United States.

Previous Literature

In 2000 Sumner and Lee studied the impact of the URAA on the world rice market by using a partial equilibrium model and summarized the overall change in international rice markets. It was concluded that the URAA made relatively small changes in rice policy world-wide; however, these small changes had a negative effect on the rice production subsidy in South Korea. The model projected an increase in the price of high quality japonica rice by 7 percent, due to imports from Japan and South Korea. Likewise, Wailes (2005) conducted a comprehensive analysis on global rice trade and studied protectionist policies and the impact of liberalization. The study stated that the trade distorting policies; such as imports tariffs and TRQ's adopted by Japan and South Korea has led to more distortion of medium grain markets as compared to long grain rice markets. Similarly, Wailes (2005) also assessed the direct effects of domestic price support policies on world rice trade and prices using the Arkansas Global Rice Model (AGRM). He estimated that there was no impact on long term basis with elimination of domestic support in developed countries (the US, the EU, and Japan). But, global rice trade increases by 10 to 15 percent with the elimination of import tariffs, TRQ's and export subsidy.

Trade policy reforms will be expected to increase export prices for exporters by 25 to 35 percent and decrease import prices for importers by 10 to 40 percent (Wailes, 2005). To conclude, the lack of policy reforms at the national level in developing countries for past decade has led to increased price volatility, high economic costs to poor consumer and excessive cost in undertaking food distribution programs for governments (Wailes, 2005).

In 2005, using the AGRM model Wailes studied trade liberalization in rice. It was concluded that domestic support had a negligible impact on long and medium grain prices, whereas full trade liberalization including elimination of tariffs and subsidies increased long grain and medium grain price by 22 and 80 percent respectively. Babcock et al. (2003) were the first to carry out an analysis of the proposed DDA modalities. It was found that expansion of TRQs under the proposed DDA would provide the maximum increase in trade volume. The greatest growth in quota level would take place in China, Indonesia and the Philippines. They found that the commitment to reduce the export subsidy to zero would have a negligible impact on rice trade.

The only study that focused on a specific proposal in the ongoing WTO negotiations was carried out by the Food and Agricultural Policy Research Institute (FAPRI). As part of that study, the Center for Agricultural and Rural Development (CARD) in 2005 investigated the impact of the US proposal on the US and world agriculture. It was found that the US proposal would increase the world reference long grain and medium grain rice price by 9 percent and 25 percent, respectively.

Arkansas Global Rice Model (AGRM)

As developed and modeled by Fuller, Wailes, and Djunaidi (2003) the Arkansas Global Rice Model is a multi-country, partial equilibrium, dynamic econometric simulation framework.

The AGRM is a mathematical representation of the world rice economy. The model consists of six sub regions and 32 specific countries. Each country model has a supply sector, demand sector, a trade and price linkage equations. Estimates are based on the exogenous macroeconomic factors such as income, population, inflation rate, technology development and especially government determined policy variables. The Thai FOB (5% broken, Bangkok) price is used to clear the international rice market for the long grain and the California ex mill FOB price is used to clear the market for the medium grain. The projections in the model are based on national levels of production (area harvested and yields), consumption, net trade, stocks and prices. The model provides projections of the rice economy for a 10-year period.

The baseline projection and policy analysis are estimated by simulation. Trade distortion in the international rice market is captured by explicitly incorporating government policies in the estimated equations. Most government induced policy distortions are clearly stated in the model's structure.

The policies are incorporated in the model through supply, demand, export (or import), ending stocks and price transmission equations and are thus embedded in the model solution (Fuller, Wailes, and Djunaidi, 2003).

Data Sources

Macroeconomic economic data used in the AGRM model are based on Wharton Economics Forecasting Associates and project LINK (FARPI, 2006). Similarly, the US Department of Agriculture (USDA) online database was used for the Production, Supply and Distribution (PSD) of agricultural commodities in the AGRM model. The WTO database outlining the modalities on the alternative proposals are used for scenario simulations. USDA, Foreign Agricultural Services (FAS) Attaché Reports, and other sources were used to develop the 2006 AGRM baseline.

Policy Assumption for Analyzing Alternative DDA Proposals

The US, EU, G20, and G10 proposals are analyzed based on the policy changes over a period of eight years (2007-2014). The analysis of the proposals is based on the AGRM 2006 baseline and the proposal modalities (Table 1). This baseline was developed in early 2006 and reflects baseline projections developed by the FAPRI consortium. The computation of tariffs is based on Ad Valorem Equivalents (AVE) with changes expressed in percentage of AVE equivalent (see WTO committee on Agriculture for methodology on computing AVE equivalents). The results focus on the final year of the projection period (2014) to describe the differences among proposals on prices and rice trade.

Impacts on World Trade

The results for the different proposals for changes in net world rice trade are increases of 4.95, 1.76, 1.55, and 1.55 percents under the US, the EU, the G20, and the G10 proposals, respectively, over the baseline by 2014 (Table 2). The US proposal increases global rice trade by 1,474,000 mt (4.95 percent) above the baseline. Increased trade under the US proposal is largely due to estimated tariff cuts of 75 percent. The G10, and the G20 proposals have the least increase in net global trade by 460,000 mt (1.55 percent) above baseline. The average estimated tariff cut under the US proposal is 25-30 percent, while the G20 proposal has an estimated tariff cut of 54 percent (CRS Report, 2005).

The long grain rice trade has the highest increase under the US proposal followed by the EU and the G20 proposals by 2014. The US proposal increases the global net long grain trade by 1,066,000 mt or 3.70 percent. The G20 proposal has the second highest estimated tariff cut of 54 percent, which increases net long grain trade by 501,000 mt or 1.74 percent by 2014 (Table 2).

Net medium grain trade is largely influenced by market access reforms in major medium grain rice importing countries like Japan and South Korea. The US proposal has the highest increase in medium grain followed by the G10, the EU and the G20 proposals, respectively.

The US proposal would increase net medium grain trade by 598,000 mt or 20.12 percent by 2014. The G20 proposal would increase the net medium grain trade by only 17,000 mt or 0.56 percent by 2014. The G20 proposal does not have much market access reforms in major rice importing countries like Japan and South Korea.

Table 1. Policy Assumption under the US, the EU, the G20 and G10 Proposals

Market Access	US	EU	G20	G10
Developed Countries				
Estimated Average Tariff Cuts (%)	25-30	46(39)[a]	54	75
Developing Countries				
Tariff Cap %	100	150	150	No Cap
Domestic Support				
Target Price	-7	-7	-8	-7
Loan Rate	-11	-11	-13	-11
Export Competition				
Elimination of Export subsidies	2010	2013	2010	2013
Source: CRS Report, 2005				
[a] USTR estimate				

Table 2. Impacts of Alternative WTO Proposals on Rice Trade, 2007-2014

Year	2007	2008	2009	2010	2011	2012	2013	2014	Avg
Net Trade					(Thousand Metric Tons)				
Baseline	25,686	26,734	27,392	27,825	28,319	28,817	29,268	29,737	27,972
Proposals	% Change								
US	0.68	1.29	1.64	2.20	2.97	3.85	4.54	4.95	2.77
EU	0.19	0.31	0.40	0.49	0.65	1.12	1.52	1.76	0.81
G20	0.17	0.28	0.31	0.35	0.58	0.98	1.37	1.55	0.70
G10	0.23	0.33	0.40	0.42	0.64	1.01	1.28	1.55	0.73
Net Long Grain Trade					(Thousand Metric Tons)				
Baseline	24,580	25,669	26,318	26,759	27,286	27,825	28,324	28,813	26,947
Proposals	% Change								
US	0.50	0.89	1.16	1.46	1.94	2.61	3.25	3.70	1.94
EU	0.21	0.36	0.46	0.56	0.77	1.15	1.50	1.75	0.84
G20	0.20	0.37	0.47	0.56	0.78	1.17	1.52	1.74	0.85
G10	0.15	0.23	0.23	0.24	0.39	0.63	0.89	1.08	0.48
Net Medium Grain Trade					(Thousand Metric Tons)				
Baseline	2,910	2,887	2,938	2,954	2,967	2,948	2,955	2,972	2,941
Proposals	% Change								
US	3.04	5.42	10.25	13.56	18.16	19.50	20.21	20.12	13.78
EU	0.33	0.52	0.54	0.64	2.11	1.56	2.19	2.31	1.28
G20	0.03	-0.13	-0.48	-0.79	-0.68	0.02	0.60	0.56	-0.11
G10	0.88	1.67	2.31	4.17	5.48	4.73	6.21	5.29	3.84

Impacts on Prices

In general, long grain rice prices increase under the US, the EU, the G20 and the G10 proposals. An increase in world prices is largely attributed to lower import tariffs in major long grain importing countries like Nigeria, Indonesia and the Philippines. By 2014, world prices would increase by 8.59, 3.49, 3.73 and 2.06 percents under the US, the EU, the G20 and the G10 proposals respectively (Table 3).

Table 3. Impacts of Alternative WTO Proposals on Rice Prices, 2007-2014

Year	2007	2008	2009	2010	2011	2012	2013	2014	Avg
World Price, Thai 100% FOB				(US Dollars per Metric Ton)					
Baseline	253	248	263	276	286	298	311	323	282
Proposal	% Change								
US	1.16	2.72	3.96	5.19	6.36	6.81	7.46	8.59	5.28
EU	0.55	1.35	1.99	2.43	2.87	3.16	3.13	3.49	2.37
G20	0.50	1.29	1.97	2.44	2.82	3.10	3.26	3.73	2.39
G10	0.36	0.97	1.43	1.71	2.02	2.03	1.82	2.06	1.55
US No.2 Medium Grain Rice fob CA				(US Dollars per Metric Ton)					
Baseline	571	541	551	569	564	556	541	542	554
Proposal	% Change								
US	7.48	18.94	20.75	33.50	40.70	37.20	34.13	37.71	28.80
EU	1.06	2.75	4.88	7.02	3.81	6.08	3.36	3.20	4.02
G20	-0.06	0.58	1.86	3.02	2.63	0.02	-2.00	-1.44	0.58
G10	3.17	6.85	10.84	10.59	12.86	16.37	8.84	12.80	10.29

The US proposal would increase the reference price by USD 28 per mt or 8.59 percent, while the EU and the G20 proposal would increase world reference by USD 11 per mt or 3.49 percent and USD 12 per mt or 3.73 percent respectively. A larger increase in world price for the US, the EU and the G20 proposals are due to deeper tariff cuts in major long grain importing countries like the Philippines and Indonesia. These tariff cuts increase trade, which directly increases the world reference price. The G10 proposal results in a minor increase of USD 7 per mt or 2.06 percent in world price. A minor increase in the world reference price under the G10 proposal is due to little market access reform in major long grain importing countries such as Indonesia and the Philippines. The U.S. proposal has the highest increase in world price due to significant market access reforms in Indonesia, Philippines and the EU contributing to increased import demand. The medium grain world reference price (US No.2 MG Rice fob CA) increases by 37.71, 3.20, and 12.80 percents under the US, the EU, and the G10 proposals, respectively.

Higher medium grain reference prices are largely attributed to increased minimum market access (MMA) in South Korea and TRQ expansion with reductions in tariffs in Japan under the US proposal. The TRQ in Japan is almost doubled under the US proposal, while under the G10 proposal there is an expansion of its TRQ by 83 percent. There is no expansion of the TRQ and MMA in Japan and South Korea under the G20 proposal, resulting in the decrease of medium grain price by USD 7 per mt or 1.44 percent by 2014. The US proposal has the largest increase in medium grain price by USD 205 per mt or 37.71 percent. These results are quite similar to estimates by Wailes (2005) on complete trade liberalization using the RICEFLOW and AGRM models.

Prospects for Long Grain Rice Exporters and Importers

In this section the impacts of various trade proposals with respect to major trading countries are discussed. Trade liberalization under all four proposals would increase exports for major long grain exporting countries (Table 4).

Initial reduced exports from the US are largely attributed to substantial cuts in its amber box subsidies. Amber box subsidies would be reduced in the U.S. by 60 percent under the US, the EU and the G10 proposals, while the G20 proposal would reduce the US amber box payments by 70 percent. There would be an increase in US long grain exports under the US, the G20, the EU and the G10 proposals over the longer run. The largest increase would be under the US proposal resulting in an increase of 3.30 percent or 98,000 mt by 2014, whereas the effects of the other proposals on US long grain exports are slightly smaller (Table 4). However on average, only the US proposal would increase US exports, by 9,000 mt or 0.25 percent. Thailand, India, Vietnam and Pakistan increase their exports by more than twice as much on average by 2014 under the US proposal when compared to the EU, the G20, and the G10 proposals.

Thailand, which is the world's largest exporter of long grain rice, would have a steady increase in its exports by 2014 with 3.92 percent more or 396,000 mt over the baseline under the US proposal. Other major long grain exporters would be India, Pakistan, and Vietnam. India and Vietnam would increase their exports due to opening of markets in Indonesia and the Philippines. Indian rice export supply increases primarily because of higher yields. While Vietnam increases in export supply is a result of both higher yields and a larger area harvested. Exports from India and Vietnam under the US proposal by 2014 are higher compared to the baseline by 3.67 percent or 181,000 mt and 2.52 percent or 154,000 mt, respectively. Likewise, Pakistan has the highest percent increase in exports under the US proposal and the highest among all major long grain exporting countries, with an increase of 6.93 percent or 158,000 mt over the baseline. Higher yields and increases in area contribute to larger exportable rice supplies in Pakistan.

Indonesia, the Philippines and the EU would be the major long grain importers under all four proposals. There would be large increases in imports by 2014 for Indonesia under the tariff reductions of the EU and the G20 proposals by 43.54 percent (791,000 mt) and 42.81 percent (777,000 mt) respectively. The G10 proposal provides the smallest increase in imports by Indonesia because it only requires a 37 percent reduction in tariffs.

In our analysis the Philippines is assumed to declare rice as a sensitive product[3] (ICTSD, 2006). Imports increase in the Philippines mainly due to the lower in-quota tariffs. The present in-quota tariff of 50 percent in the Philippines would be reduced to 5 percent over a period of 10 years under the US proposal.

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[3] Sensitive products under the US proposal is 1% of tariff lines while, TRQ expansion would be based on 7 percent of the domestic consumption to base year 1999-2001. MMA is assumed to be 7.5 percent of the average annual domestic consumption of year 1999-2001. Similarly the EU proposal would expand TRQ based on percentage of imports with lower tariff cuts and MMA is assumed to be 5 percent of the average annual domestic consumption of year 1999-2001. Likewise, under G 20 proposal TRQ expansion would be 6 percent of the domestic consumption to base year 1999-2001. Finally, under G 10 proposal, if the current TRQ for sensitive products is equal or less than 5 percent of domestic consumption then it should be doubled. If the TRQ is greater than 5 percent of the domestic consumption, it would be increased in progressively lower amounts.

Table 4. Impacts on Major Long Grain Exporting and Importing Countries

Net Exporters	Baseline	Proposals			
		US	EU	G20	G10
	(Thousand Metric Tons)	% Change 2014			
US	2,996	3.30	2.69	2.80	1.73
Thailand	10,073	3.92	1.72	1.77	1.12
India	4,930	3.67	1.53	1.63	0.92
Vietnam	6,134	2.52	1.06	1.13	0.66
Pakistan	2,286	6.93	2.98	3.10	1.81
Indonesia	1,817	38.29	43.59	42.81	18.00
Bangladesh	1,266	-25.71	-11.97	-12.03	-7.88
Philippines	1,875	43.54	42.82	5.50	4.73
European Union	1,076	9.81	0.36	-2.81	3.11
US	2,996	0.25	-0.32	-0.25	-0.84
Thailand	10,073	1.97	0.91	0.90	0.61
India	4,930	2.12	0.97	0.97	0.64
Vietnam	6,134	1.97	0.95	0.94	0.65
Pakistan	2,286	3.83	1.77	1.77	1.18
Indonesia	1,817	23.66	25.69	25.69	9.93
Bangladesh	1,266	-15.08	-7.47	-7.35	-5.25
Philippines	1,875	23.56	4.08	2.27	1.84
European Union	1,076	7.68	0.73	-0.91	3.20

This results in a reduction in area under rice production and a large increase in imports for the Philippines occurs with the US proposal, where by 2014, there is a 43.54 percent increase or 817,000 mt over the baseline. Other proposals do not have that impact on increases in imports as there is no requirement to decrease in-quota tariffs. Bangladesh, a major long grain importer, is been classified as a least developed country under all four proposals, so there are no policy changes in Bangladesh that would increase trade. Bangladesh would decrease imports under all four proposals by about 10 percent because of higher world prices. For example under the US proposal, the much higher world prices would decrease Bangladesh imports by 25.71 percent or 371,000 mt.

The EU increases its imports under the US, the EU and the G10 proposals. Under the US proposal the general tariff in the EU on milled rice is reduced from 416 EUR/mt to 87.76 EUR/mt, a deep cut in its tariff over a period of five years. The other proposals do not have tariff cuts that substantially increase the EU imports, and under the G20 proposal imports are estimated to decline relative to the baseline.

Prospects for Medium Grain Rice Exporters and Importers

In every proposal, major medium grain exporting countries are the US, the EU, China, Egypt and Australia; whereas major importing countries are Japan, South Korea, and Taiwan (Table 5). The US is a major exporter of medium grain rice to Northeast Asia and its exports increase the most under the US proposal.

Table 5. Impacts on Major Medium Grain Exporting and Importing Countries

Net Exporters	Baseline	Proposals			
		US	EU	G20	G10
	(Thousand Metric Tons)	% Change by 2014			
US	710	36.11	7.36	2.28	14.99
China	388	21.35	0.55	0.13	1.37
European Union	145	104.03	0.98	-1.97	7.12
Australia	672	14.12	1.29	-0.39	4.95
Egypt	857	1.51	0.50	0.63	0.23
Japan	682	97.14	13.79	0	20.00
South Korea	409	25.52	0	0	0
Taiwan	127	71.93	0	0	100.00
Turkey	386	-19.44	-2.73	0	-8.23
US	710	19.88	1.46	-1.27	6.11
China	388	12.06	0.46	0.14	0.93
European Union	145	90.33	7.19	-0.23	24.71
Australia	672	11.20	1.59	0.25	4.05
Egypt	857	1.41	0.61	0.64	0.38
Japan	682	72.86	1034	0	15.0
South Korea	409	11.83	0	0	0
Taiwan	127	53.95	0	0	75.00
Turkey	386	-12.69	-2.23	-0.50	-5.32

In fact all exporters benefit the most from the US proposal. This expansion is primarily a result of the increase in the TRQ in Japan from 682,000 mt to 1,344,500 mt as required under the US proposal.

China, the European Union and Australia, increase their exports under all four proposals; with larger increases in exports under the US proposal due to the TRQ expansion in Japan and an increase in the MMA (Minimum Market Access) in South Korea. The European Union increases exportable supplies in response to higher world prices. Egypt, one of the largest exporters of medium grain rice, would have a small increase of 1.51 percent or 13,000 mt under the US proposal. There would be an initial increase in exports over a period of five years but exports would start to decline in later years of the implementation period as a result of no available land to increase area harvested. Japan and South Korea are assumed to declare rice as a sensitive product. Japan increases imports by 97 percent or 663,000 mt and by 20 percent or 136,000 mt under the US, and G10 proposals respectively. This increase in imports is based on the expansion of the TRQ as stated earlier. Japanese imports would increase by 13.7 percent or 94,000 mt over the baseline under the EU proposal as the TRQ increases from 682,000 mt to 776,000 mt. The G20 proposal has no required adjustment on the current Japanese TRQ and therefore there is no change in imports. Likewise South Korea, a major importer of medium grain rice, has agreed to increase its MMA by 2005 to 205,000 mt, so there would be no impact of the EU, the G20 and G10 proposals. However under the US proposal, South Korea would almost double its imports by 104,000 mt.

In the case of Taiwan there imports increase by 71.9 percent or 92,000 mt and 100 percent or 128,000 mt under the US, and the G10 proposals respectively. Imports in Taiwan under the G10 proposal increase as a result of a doubling of its TRQ, required if the quota is less than 5 percent of domestic consumption. In the case of the US proposal there would be an increase in imports by 92,000 mt or 71.93 percent as the TRQ expansion is based on 7.5 percent of annual domestic consumption.

As a result of market access reforms in Japan, South Korea and Taiwan, their reduction in medium grain rice acreage is compensated by acreage expansion in Australia, the US and the EU. Turkey, one of the major medium grain importers, would decrease its imports under all four proposals. With relatively low import protection, higher world medium grain prices that result from all four proposals reduces import demand by Turkey.

Impacts of Alternative Proposals

The analyses of the alternative proposals on global rice trade presented in this study fulfill the underlying objectives for the proposal as offered by each country or group of countries. The US proposal, the most ambitious of all proposals fulfills the basic objective of greater market access by reducing tariff and expanding TRQs globally. The US proposal results in greater market access to long grain importing countries and the protected markets of northeast Asia (Table 4).

On the other hand the G20 proposal focuses on reduction of domestic support in developed countries and a reduction of tariffs globally. The G20 proposal has substantial reduction in tariffs, 80 percent cut in amber box subsidies for developed countries (CRS Report, 2004), which would reduce medium grain exports of developed countries as shown in Table 5. However, the G20 proposal has no impact on protected medium grain rice markets of northeast Asia. The EU and G10 proposals have the moderate market access reforms in medium grain importing countries. The G10 proposal expands TRQs in Japan and Taiwan based on a flexibility formula (ICTSD, 2006). The EU proposal has major market access reforms for the long grain importing countries by reducing tariffs, but moderate market access reforms in medium grain importing countries as compared to other proposals.

CONCLUSION

The key results of the analysis are that the US proposal is the only proposal that results in non-trivial expansion in trade and higher prices. This is a result of the greater reforms required in market access, compared to the other proposals. The US, the EU, the G20, and the G10 proposals have positive impacts on global rice trade and prices. The results for the US proposal are similar to the FAPRI study on impacts of the US proposal on US and world agriculture. However, there are no previous studies that have examined and compared the relative effects of the US proposal against the EU, the G20 and the G10 proposals. The key finding is that the US proposal generates trade and price projections that are much larger than the other three proposals. There would be little expansion in rice trade under the reforms proposed by the EU, the G20, and the G10. The results of this study suggest that trade liberalization would be largely influenced by policy changes with respect to market access rather than changes in domestic support and export competition. Regarding long grain rice trade, Thailand, India, Vietnam, Pakistan, and the US would be the major exporters with increased exports. Indonesia, the Philippines and the European Union would be major long grain importers with increased rice imports. Likewise, medium grain rice markets in Japan and South Korea would be further liberalized under the US and the G10 proposals. Due to the further opening of these protected markets, there would be a non-trivial increases in the world reference price of medium grain rice. The expansion of trade and price impacts in the medium grain market under the US and the G10 proposals would continue trade and price effect found in previous studies on the Uruguay Round Agreement on Agriculture.

APPENDIX: AGRM BEHAVIORAL EQUATIONS

Supply Sector

$HA_t = f_1 (HA_{t-1}, P^e_t, W^e_t, e_{1t})$
HA_t = Harvested Acreage
HA_{t-1} = Harvested Acreage (for previous year)
P^e_t= Expected price received by the producers
W^e_t = Expected input price
e_{1t} = Error term
$Y_t = f_2 (P^e_t , W^e_t, T_t , e_{2t})$
Y_t = Expected yield
P^e_t = Actual input price
W^e_t = Expected input price
T_t = Research expenditure (if available)
e_{2t} = Error term

Demand Sector

$DPC_t= f_3 (M_t, RP_t, WP_t, e_{t3})$
D_t= Total domestic demand (=DPC_t *Population)
M_t = Per capita income in real terms
RP_t = Rice retail price (Weighted Average of free market price/Government ration price)
WP_t= Wheat price
e_{t3}= Error term
$EXP_t = f_5 (RESD_t, FOB_t, e_{t5})$
Demand for exports is the function of difference between domestic consumption and export prices (FOB)
EXP_t = Exports
$RESD_t$ = Residual of total production net of total consumption
FOB_t = Free on board export price measured in local currency
e_{t5} = Error term

Price Linkage

Farm Price
$P_t = f_6 (RP_t, e_{t6})$
P_t = Farm price
RP_t = Retail price
e_{t6} = Error term
Retail Price
$RP_t = f_7 (FOB_t, e_{t7})$
FOB_t = Export price
e_{t7} = Error term
Export Price
$FOB_t = f_8 (THAI\ FOB_t, e_{t8})$
$THAIFOB_t$ = Thai Price (100%B)
e_{t8} = Error term

Market Clearance

$S_t = PROD_t + S_{t-1} - D_t - EXP_t$
S_t = Ending Stocks
$PROD_t$ = Total production (total area harvested *yield)
S_{t-1} = Beginning stock
D_t = Total domestic demand
EXP_t = Exports

REFERENCES

Babcock, B.A., Fabiosa, J.F., Matthey, H., Isik, M., Tokgoz, S., Elobeid, A., Hart, C.E. Fuller, F.H. and Meyer, S. 2003. "*An Analysis of the Proposed Doha Modalities*." Staff Report 03-SR 98. Center for Agricultural and Rural Development, Iowa State University, Ames, Iowa.

Burfisher, M.E., ed.2001. *Agricultural Policy Reform in the WTO—The Road Ahead*. Market and Trade Economic Division, U.S. Department of Agriculture, Economic Research Service, Agricultural Economics Report No.802, May. http://www.ers.usda.gov/publications/aer802/aer802.pdf (Last accessed on 30 May, 2006)

Diao, X., A. Somwaru, and T. Roe. 2001. "A Global Analysis of Agricultural Reform in WTO Member Countries," *Background for Agricultural Policy Reform in the WTO: The Road Ahead*.U.S.Department of Agriculture, Economic Research Service, ERS-E01-001.

FAPRI. 2005a. "U.S. Proposal for WTO Agriculture Negotiations: Its Impact on U.S. and World Agriculture." Working Paper 05-WP 417, Center for Agricultural and Rural Development, Iowa State University, Ames.

FAPRI. 2005b. "Potential Impacts on US Agriculture of the U.S. October 2005 WTO Proposal." Report # 16-05: 1-27. University of Missouri, Columbia.

FAPRI. 2006. U.S. and World Agriculture Outlook, Iowa State University and University of Missouri Ames, Iowa, USA, FAPRI Staff Report 06 -FSR1 January. http://www.fapri.iastate.edu/outlook2006/text/OutlookPub2006.pdf (Last accessed on 3 September 2006)

Fuller, F., E. J. Wailes, and H. Djunaidi. 2003. "Revised Arkansas Global Model Rice Model" Department of Agricultural Economics and Agribusiness, University of Arkansas, Fayetteville, AR 72701, Staff Paper 01 June.

Hanrahan, C. and R. Schnepf. 2005. "WTO Doha Round: Agriculture Negotiation Proposals." Congressional Research Service (CRS) Report for Congress. RL 33144. http://www.fas.usda.gov/psd/complete_tables/GF-table7-85.htm (Last accessed on 30 May 2006)

ICTSD. 2006. "WTO Agriculture Weeks Sees G10 Proposals, but a Little Progress" *Bridges weekly* main page Vol.10 No.3. http://www.ictsd.org/weekly/06-02-01/story2.htm (Last accessed on 30 May 2006)

Ito, S., E. Wesley, F. Peterson and W.Grant. 1991. "Rice in Asia: Is It Becoming an Inferior Good? *American Journal of Agricultural Economics,* 73(2):528-532.

OECD. 1998. "Agriculture in the Uruguay Round." OECD Workshop on Emerging Trade Issues in Agriculture, Directorate for Food, Agriculture and Fisheries Trade Directorate. http://are.berkeley.edu/courses/ARE242/spring05/classReadings/papers/OECD_AgInUruguay_98.pdf (Last accessed on 13 March 2006)

Sumner, D.A. and H. Lee. 2000. "Assessing the Effect of the WTO Agreement on Rice Markets: What Can We Learn from the First Five Years?" *American Journal of Agricultural Economics* 82(3):709-717.

United Nations, Food and Agriculture Organization (FAO). 2004. "Economics and International Year of Rice" Rice is Life – International Year of Rice 2004, Pg 2 of 2. http://www.fao.org/rice2004/en/f-sheet/factsheet4.pdf (Last accessed on 30 Sept 2006)

United Nations, Food and Agriculture Organization (FAO). 2005. Commodity Market Review 2005-2006, Rome. ISBN 92-5-105437-1. ISSN 1020-492X http://www.fao.org/docrep/008/a0334e/a0334e00.htm (Last accessed on 28 November 2006)

U.S. Department of Agriculture. 2005. Briefing Room, Rice Background, Economic Research Service. Washington, DC.

U.S. Department of Agriculture. 2006. Production, Supply and Distribution Online. USDA, Foreign Agriculture Service. Washington, DC.

U.S. Department of Agriculture. 2012. USDA Agricultural Baseline Projections to 2021, Long-term Projections Report OCE-2012-1. Washington, DC.

Wailes, E. J. 2002. "Trade liberalization in Rice." in *Agricultural Trade Policies in the New Millenium*, ed. P. Lynn Kennedy. Haworth Press.

Wailes, E. J. 2004. "Implication of the WTO Doha round for the Rice Sector" selected paper at FAO Rice Conference, Rome Italy, 11-13 February.

Wailes, E. J. 2005. "Rice: Global Trade, Protectionist Policies and the Impact of Trade Liberalization." In *Global Agricultural Trade and Developing Countries*. Eds. M. Ataman Aksoy and John C. Beghin. World Bank.

Wailes, E. J. 2006. Arkansas Global Rice Model (AGRM) Baseline Projections 2006 - 2016, University of Arkansas, Fayetteville, AR 72701.

World Trade Organization. 2004. "The Issues and Where We are Now". WTO Agriculture Negotiations. 1st Update December 2004. http://www.wto.org/English/tratop_e/agric_e/agnegs_bkgrnd_e.pdf (Last accessed on 2 February 2006).

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BORDER EFFECTS ON SPATIAL PRICE TRANSMISSION BETWEEN FRESH TOMATO MARKETS IN GHANA AND BURKINA-FASO: ANY CASE FOR PROMOTING TRANS-BORDER TRADE IN WEST AFRICA?

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ABSTRACT

Trans-border, agricultural commodity trade within sub-regional economic blocks in Sub-Sahara Africa (SSA) is believed to be faster, cheaper, more convenient and welfare-enhancing than trade between SSA countries and the USA, EU or China. Trade flow difficulties across international borders in SSA are however a fundamental disincentive to cross-border price transmission and market integration. This study examines the effect of the border between Ghana and Burkina-Faso on price transmission between tomato markets in these countries. Applying a regime-switching Vector Error Correction Model (VECM) to semi-weekly prices of tomato in major Ghanaian tomato markets trading with a production centre in Burkina-Faso, we discover that price transmission between the markets contains evidence of border effects. High transaction costs and tariffs, the perishable nature of tomato, and poor quality roads and transport facilities between the markets make trade very costly and risky to arbitrageurs. The findings have theoretical and practical relevance for cross-border trade in West Africa.

Keywords: Price Transmission, Trade, Fresh Tomato, Ghana, Burkina-Faso

JEL Codes: C32, Q11, Q13, Q17, Q18

1. INTRODUCTION

Agricultural trade in the emerging economies of Brazil, Russia, India and China (BRIC), as in many other developing countries, is undergoing a tremendous, directional shift. The old trading order is characterised by the conventional exports of raw material and primary food

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products by developing countries overseas to countries with which they have historical, colonial links and conversely imports of finished products from the latter to the developing countries. The emerging order of trade regimes tends to be within regional and sub-regional economic blocks and aims at improving sub-regional supply chains and meeting regional development partnership agreements. The ongoing process of globalisation and domestic market reforms under WTO establishments even appears to have assumed diminished importance under this new trading pattern.

The new trend comes not as a surprise to many economic analysts. With deadlock on WTO Doha Round of negotiations, the agricultural sectors of many low income WTO member countries have been left opened to the vagaries of global economic shocks. The agricultural sectors of Sub-Sahara African (SSA) countries have particularly become extremely susceptible to shocks from the world economic and food price crises. The crises-related effects on agricultural markets in SSA countries have been compounded by the increasing need for high quality standards in products, sanitary and phytosanitary norms and changing demand in the markets of SSA's traditional trading partners.

While the changing scenario of agricultural trading systems is unabated, it is unclear how successful this new system of trade will be. The question is does the new system offer great opportunities for the agricultural sectors of developing countries or does it represent a likely case in which the exports of developing countries have become more vulnerable to the risks and uncertainties of regionalism? Whatever the case may be, what is clear is that the future of the new system will largely depend on the incentives of cross-border trade, including efficient price transmission and market integration between sub-regional trading partners.

The transmission of price signals between spatially separated markets plays an important role in explaining market performance, their degree of integration or isolation, and the speed at which price signals are transmitted between surplus, producer markets and deficit, consumer markets for a given commodity. Knowledge on the extent of price transmission is useful in guiding production and consumption decisions, and in stimulating inter- and intra-regional trade flows needed to buffer the price and welfare effects of local supply and demand shocks (VON CRAMON-TAUBADEL and IHLE, 2009). If no barriers to trade exist, prices of a given commodity at geographically separated locations should be so strongly linked that price shocks in individual markets within a given country evoke responses in the corresponding markets of its trading partners.

The seasonal arbitrage of fresh tomato between Ghana and her northern neighbour Burkina-Faso represents one of the largest and strongest cross-border trades in agricultural commodities in West Africa. Between January and June yearly, about 35,000 tonnes of fresh tomato are imported from Burkina-Faso into Ghana via the arbitrage activities of itinerary female traders called *market queens* (IRIN AFRICA, 2009)[1]. While there exists a free flow of tomato across the border of the two countries because of the absence of formal trade restrictions, the presence of the international border may nonetheless impose important costs in the form of tariffs, arbitrage delays, corruption and harassment of traders by border officials, the cost of changing exchange rates and communication barriers between Ghanaian traders and Burkinabe tomato producers. This means, tomato price formation in this period (hereafter called the Burkina-Faso regime) is affected not only by demand-supply shocks in

[1] This is only the recorded volume of trade; the non-recorded component may be as high.

Ghana and the above listed costs to trade, but also by price dynamics in Burkinabe tomato markets.

Most empirical research on vertical, spatial or temporal price relationships in agricultural markets tends to examine the underlying factors likely to drive arbitrage, price transmission and integration between markets. For example, Von Cramon-Taubadel (1998), Abdulai (2000), Meyer and Von Cramon-Taubadel (2004) test the implications of market power on asymmetric price transmission in selected markets in Germany and Ghana. Moser et al (2006) examine the role of crime rate, remoteness and lack of information in the sub-regional integration of rice markets in Madagascar.

Recently, Jensen (2007) and Aker (2008) shed light on the importance of information flow on stochastic price processes in spatially separated markets in India and Niger respectively. Stephens et al. (2008) and Ihle et al (2010) also examine the possibility of mechanisms, other than physical trade flow, causing stochastic price adjustments in periods during which no direct trade takes place between spatially separated markets.

Whereas the above studies, among others, immensely contribute to our understanding of the performance of agricultural produce markets, their common limitation include failure to consider the existence of political impediments such as regional or national borders to arbitrage and spatial price transmission.

This study intends to add to our understanding of price transmission by extending the analysis to examine whether or not price transmission between fresh tomato markets in Ghana and Burkina-Faso displays evidence of distance and border effects. We fundamentally seek to determine the nature of price dynamics in Ghanaian tomato markets in the season of supply from Burkina-Faso and in the season without tomato imports from Burkina-Faso.

Specifically, we intend to address the following objectives:

1. To determine the speed of price transmission between four major fresh tomato consumer markets in Ghana and the producer market, when Burkina-Faso is the major source of tomato in the Ghanaian fresh tomato market by estimating cross-country price transmission parameters.
2. To determine the speed of price transmission between the four consumer tomato markets in Ghana and the producer market, when Ghana’s fresh tomato supply source is local by estimating within-country price transmission parameters, and
3. To check whether distance and the international border between markets in the two countries matter for price transmission by comparing the estimated cross – and within-country price transmission parameters.

With respect to the above objectives, the analysis is performed under two major, seasonally depended, tomato production regimes – a Burkina-Faso regime of fresh tomato supply to Ghanaian markets from January to May and a non-Burkina-Faso (Techiman) regime of fresh tomato supply from June – December, yearly.

We therefore use a regime-dependent vector error correction model (VECM) and a semi-weekly, wholesale level dataset of prices from Burkina-Faso and Techiman as tomato-producing markets and Tamale, Kumasi and Accra as tomato consuming markets, for the analysis.

Our task is to quantify the degree of market integration, subject to regime disparities in fresh tomato supply levels, inter-market trade flows, prices and transaction costs between the selected Ghanaian tomato markets connected by trade to production areas in Burkina-Faso. In this way, our analysis is spatiotemporal in nature since it examines price transmission across space (market locations) and over time (season-to-season). The estimated parameters are expected to serve primarily as indicators of border and distance, and secondarily as seasonal effects on the integration of the markets under study. Generally, they indicate the potential of cross-border trade in agricultural commodities within the Economic Community of West African States (ECOWAS).

In the next section, the study area and data used for the analysis are described, while two variants of the VECM - standard VECM and Regime-dependent VECM are specified and their suitability for applying them to the dataset are stated in section 3. In section 4 we present and discuss the findings, and finally draw the conclusions to the study and suggest its implications for policy and further research in section 5.

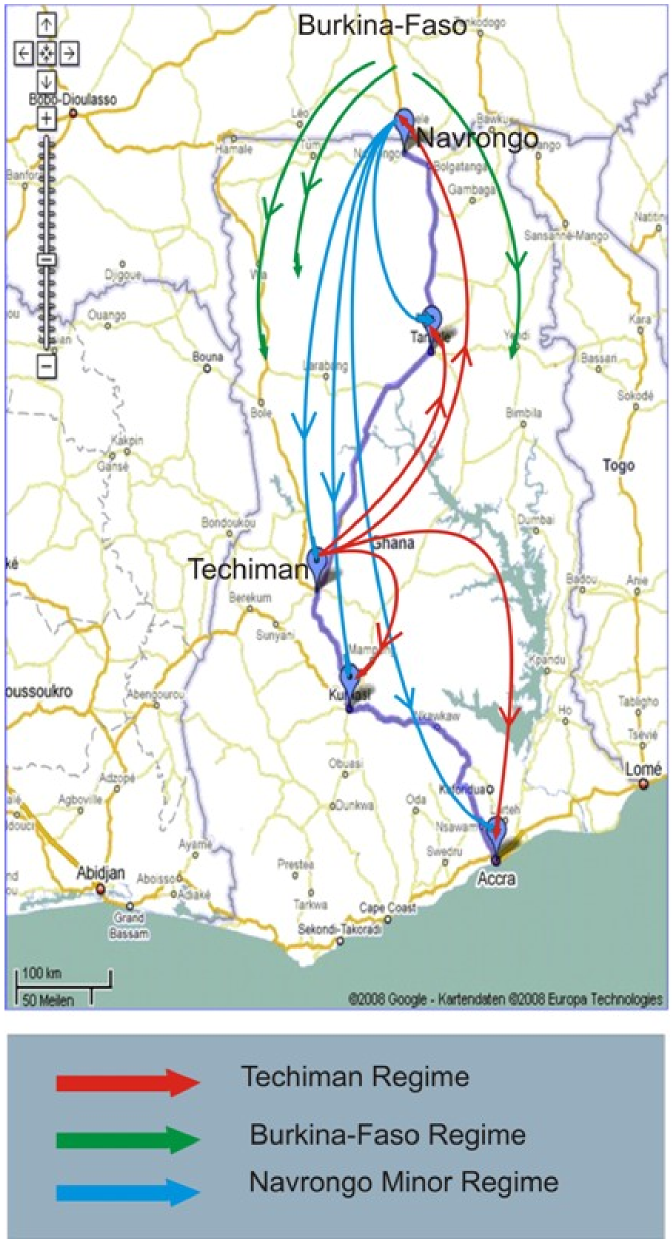
2. STUDY SETTING AND DATASET

Four major Ghanaian tomato markets – Accra, Kumasi, Techiman and Tamale, and a key production centre - Po in the Central Province of the Republic of Burkina-Faso are considered under the analysis. The Ghanaian tomato markets include one net producer market - Techiman which supplies a substantial share of Ghana's fresh tomato in the rainy season – June to December, and three net consumer markets namely Tamale, Kumasi and Accra located in the three largest cities of Ghana. The markets under study and the pattern of trade flows between them are depicted in Figure 1.

Due primarily to differences in the weather conditions between Ghana's major producer market – Techiman and that of Burkina-Faso – Po, fresh tomato supply to the Ghanaian tomato market is seasonally-switching. The producer market in Po is located in the Sudan Savannah climatic zone and is dependent on irrigated production. It is a prominent source of tomato supply, contributing about 60% of the total demand of fresh tomato to the Ghanaian market in the dry season (December - May). Techiman (and surrounding areas) located in the southern, forest region of Ghana, using a rain-fed production system, supplies the marketing system with tomato in the rainy season (June- December). Alongside the Po supply season is another major supply and trade flow regime – Navrongo, which supplies an estimated 40% of fresh tomato within the dry season.

From the above description, three trading regimes can be identified in Ghana's tomato marketing system in terms of supply and trade flows, namely the Navrongo and Techiman regimes with domestic sources of supply, and the Po regime representing fresh tomato imports from Burkina-Faso. Table 1 presents the mean values and standard deviations of each of the price series under the analysis for the Po and Techiman regimes.

Examining the mean and the standard deviation values of the price series under each regime in Table 1, it can be observed that the mean values ranging from about *GH¢*59.00 to *GH¢*99.00, and the standard deviations ranging from *GH¢*24.00 to *GH¢*73.00 for the price series under the Po regime are higher than the corresponding values from *GH¢*37.00 to *GH¢*73.00 and *GH¢*18.00 to *GH¢*40.00 respectively under the Techiman regime. The regime-disparities in the descriptive statistics computed above may influence dynamic price adjustment processes between the markets under study.



Source: Google Earth with Own Illustrations.

Figure 1. Map of Ghana showing the markets and pattern of seasonal trade flow.

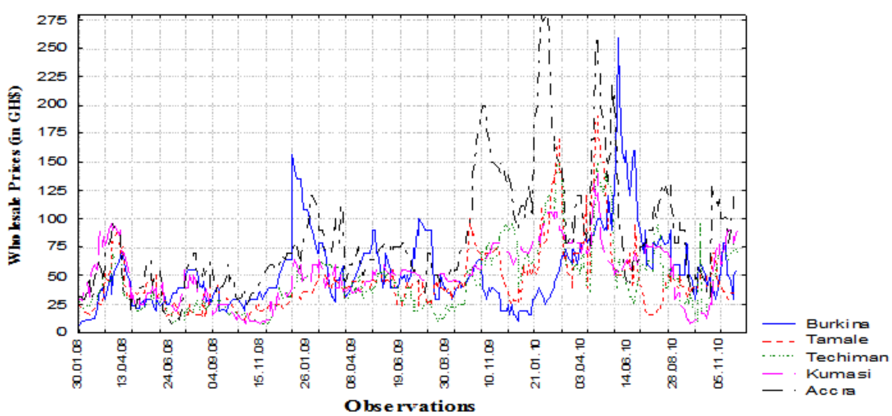
From the above description, we hypothesized that border and distance effects influence the integration of the markets under study. We expect these effects to delineate the inter-market arbitrage processes into two principal regimes, defined by the sources of tomato supply and trade flow. These, as defined above, are the domestic, Techiman regime and the cross-border Po regime. That is, changing supply levels from the two sources act as shocks to market equilibrium and solicit different responses, depending on the patterns of trade flow and switching in the source of supply as illustrated in Figure 1.

Table 1. Regime-dependent average and standard deviation of the price series (GH¢/crate)[2]

Market	Po Regime		Techiman Regime	
	Mean	Stand. Dev.	Mean	Stand. Dev.
Po	61.21	39.92	50.35	28.37
Tamale	64.36	72.96	37.27	17.67
Techiman	58.65	32.29	37.14	23.08
Kumasi	62.37	24.25	44.82	22.44
Accra	98.95	62.25	73.42	40.06

Source: Authors' Own Computation.

The complete data used for the analysis comprises a two-year long, semi-weekly, wholesale level price series from 1/2008 and to 12/2010 with 348 observations. The series is generated through self-conducted market surveys administered continuously for the four tomato production seasons in the five tomato markets – Po, Tamale, Techiman, Kumasi and Accra. The prices are those observed for the best quality of tomato available at each point of the survey in the given markets. Figure 2 depicts a graphical plot of the five price series.



Source: Own Plot from Market Data.

Figure 2. Wholesale Level, Semi-Wekkly Prices of Fresh Tomato in Ghana and Burkina-Faso (1.2008 - 9.2010).

The graphical analysis of fresh tomato prices in the five markets show the normal pattern of variability pertaining to prices of perishable commodities. It appears that the markets of Kumasi, Techiman, Tamale and Po represent price series that are very much related in terms of co-movement. Whereas Accra follows a similar variability pattern for much of the time in the period of the analysis, fresh tomato prices in Accra tend to largely lead the rests of the prices over the entire period of the analysis. This is expected, since Accra is the largest consumer market for fresh tomato in Ghana, while its farthest location (longest distance) from the domestic and cross-border production centres implies

[2] Since the Navrongo regime supplies just about 40% of usually a lower quality fresh tomato, and occurs simultaneously with and is geographically proximate to the Po regime; and since our interest is to examine within-country and cross-border price transmission processes in distinct supply regimes, we drop Navrongo and use only the Techiman and Po regimes in the analysis.

higher transaction costs and arbitrage risks, major components of the price for perishable commodities in Sub-Sahara Africa. These features theoretically represent trade and price transmission impediments between Accra and the producer markets.

3. METHODOLOGY

The main model applied to analyse our high frequency data is the vector error correction model (VECM) which focuses on prices and trade regimes instead of price margins and transaction costs like the threshold autoregressive (TAR) models. Two variants of this model are used – the standard VECM and the regime-dependent VECM. We first estimate the standard VECM which does not account for regime disparities in estimating the price dynamics. Then we estimate the regime-dependent VECM which decomposes price adjustment parameters into dynamic reactions of prices to deviations from long-run equilibrium for periods with and without tomato imports from Burkina-Faso. In this way, we obtained evidence on the nature of price transmission in the absence of direct, physical trade flows of fresh tomato from Burkina-Faso to Ghana and in the presence of same. In this section, we present the theoretical framework of the VECM and specify the variants of the VECM applied to our analysis.

If the prices on the net producer market "s" and the net consumer market "c" have a long-run relationship between them i.e. they are cointegrated, we may denote the equilibrium relationship between the net consumer prices series P_t^c and net producer price series P_t^s as: $P_t^c - \beta_1 P_t^s - \beta_0 = v_t$. If v_t, the error term, is assumed to follow an autoregressive (AR) process, then $v_t = \alpha v_{t-1} + \varepsilon_t$. This means the equilibrium relationship between P_t^c and P_t^s can be expressed as:

$$P_t^c - \beta_1 P_t^s - \beta_0 = \alpha v_{t-1} + \varepsilon_t \tag{1}$$

3.1. The Standard VECM

Equation (1) implies that the long run cointegration relationship between P_t^c and P_t^s is a function of the autoregressive process v_{t-1}. In the above linear representation, v_{t-1} represents deviations from long run equilibrium, and is called the error correction term (ECT), while α measures the response of P_t^c and P_t^s to deviation from equilibrium following random shocks to the markets. Formally, α is the loading coefficient[3] or speed of price adjustment (Lütkepohl and Krätzig, 2004). We derive the standard VECM from equation (1) by specifying changes in each of the contemporaneous prices, $\mathbf{\Delta P_t^c}$ and $\mathbf{\Delta P_t^s}$, as a function of the

[3] This is because α contains some weights attached to the cointegration relationship in the individual equations of the VECM and determines if the cointegration relationship enters the equations significantly. For instance, a loading coefficient with a t-value ≥ 2 implies the cointegration relationship is an important term in the VECM equations.

lagged short term reactions of both prices, $\mathbf{\Delta P^c_{t\text{-}k}}$ and $\mathbf{\Delta P^s_{t\text{-}k}}$, and their deviations from equilibrium at period $t-1$ (i.e. ECT_{t-1}) as follows:

$$\mathbf{\Delta P^c_t} = \delta_1 + \boldsymbol{\alpha}^{\mathbf{c}}\left[ECT_{t-1}\right] + \sum \boldsymbol{\beta}^{\mathbf{c}}_{\mathbf{k}}\mathbf{\Delta P^c_{t\text{-}k}} + \sum \boldsymbol{\beta}^{\mathbf{cs}}_{\mathbf{k}}\mathbf{\Delta P^s_{t\text{-}l}} + \varepsilon^c_t$$
$$\mathbf{\Delta P^s_t} = \delta_2 + \boldsymbol{\alpha}^{\mathbf{s}}\left[ECT_{t-1}\right] + \sum \boldsymbol{\beta}^{\mathbf{sc}}_{\mathbf{k}}\mathbf{\Delta P^c_{t\text{-}k}} + \sum \boldsymbol{\beta}^{\mathbf{s}}_{\mathbf{k}}\mathbf{\Delta P^s_{t\text{-}l}} + \varepsilon^s_t \quad (2)$$

where $\mathbf{\Delta P_t = (\Delta P^c_t \ \Delta P^s_t)'}$ is a vector of first differences of prices in the markets *c* and *s*; the $\boldsymbol{\beta}_{\mathbf{k}}$, *k = 1,…, n,* are *(2 x 2)* matrix of coefficients quantifying the intensity of the response of the contemporaneous price differences to their lagged values (i.e., they express the short-run reactions of the matrix of prices $\mathbf{P_t}$ to random shocks), and ε_t is assumed to be a white noise error term. The two equations in (2) can be generally reformulated as:

$$\Delta\mathbf{P}_t = \boldsymbol{\alpha}_{\mathbf{0}} + \boldsymbol{\alpha}_1\mathbf{EC}T_{t-1} + \sum_{i=1}^{k}\mathbf{\Gamma}_i\Delta\mathbf{P}_{t-1} + \varepsilon_t \quad (3)$$

where $\mathbf{\Gamma}_i$, i = 1… k, is a k x k matrix of short run coefficients ($\boldsymbol{\beta}_{\mathbf{k}}$) with k = 2 in our pair-wise analysis. The error correction term, ECT_t , so named because it depicts deviations from the long run relationship or 'errors' that are 'corrected' by the price transmission process,?? is a continuous and linear function of the deviation of $\mathbf{P}_t$ from the long-run equilibrium relationship following a shock on P^s_t or P^c_t; the $\boldsymbol{\alpha}_{\mathbf{0}}$ denotes long-run inter-market price margins. The loading coefficient $\boldsymbol{\alpha}_1 = (\alpha^s, \alpha^c)$ are the elasticity of price transmission or the speeds of price adjustment by the net producer and net consumer markets respectively, to deviations from long-run equilibrium. The closer a value of $\boldsymbol{\alpha}$ approaches one in absolute terms; the faster the deviations from equilibrium become corrected.

3.2. The Regime-Dependent VECM

The regime-dependent VECM, unlike its linear form above, is specified to distinguish between unique price adjustment behaviour in periods with or without fresh tomato imports from Burkina-Faso to Ghana. As stated in section 3, switching between the two principal sources of fresh tomato supply – Po and Techiman, to Ghanaian markets occurs seasonally. The regime-dependent VECM is therefore used to establish whether the degree of price transmission between Po and Ghana's tomato markets under study differs across the specified regimes. We do this by estimating pair-wise the speed of price transmission between each of the producer markets and the consumer markets.

Using information on trade flows contained in our dataset, we specify, following Ihle et al (2010) an indicator variable q_t^{cs} for tomato imports via the indicator function $I_t^{cs} = \mathbf{I}(q_t^{cs})$. The q_t^{cs} which denotes the quantity of imports equals 1 if imports of tomato from Po to the markets in Ghana occurs i.e. $q_t^{cs} > 0$, and 0 if no imports between Po and the Ghanaian tomato markets occurs (i.e. $q_t^{cs} \leq 0$).[4] This means we assume a stable long run equilibrium relationship that distinguishes between imports (the Po regime) and no-imports (the Techiman regime), and specify the following empirical model:

$$\mathbf{\Delta P}_t = \boldsymbol{\alpha}_1 ECT_{t-1}^{Po} I_t^{cs} + \boldsymbol{\alpha}_2 ECT_{t-1}^{Techiman}(1 - I_t^{cs}) + \sum_{i=1}^{k} \mathbf{\Gamma}_i \mathbf{\Delta P}_{\mathbf{t-1}} + \varepsilon_t \tag{4}$$

where ECT^i ($i \approx$ imports and no-imports or the Po and Techiman regimes) again denotes the error correction term, α_1 measures the speed of price adjustment under the Po regime and α_2 measures the speed of price adjustment under the Techiman regime. All other notations are as defined under the standard VECM. It can be seen that model (4) is an extension of (3), with the error correction term in (4) specified to behave differently under periods of imports (direct trade between Burkina-Faso and Ghana) and no-imports (local supply within Ghana/from Techiman). To emphasise, this varying behaviour of price adjustment under the two supply patterns constitutes two regimes – the Po and Techiman regimes.

Having pointed out earlier that transaction costs and trade flow reversal are the two most important determinants of price dynamics in the markets under study, models like the RVECM that are capable of accounting for both would be the most ideal for analyzing price transmission in the markets.

4. Results and Discussion

4.1. Unit Root and Cointegration Tests

Following the traditional approach of time series analysis, we first test for a random walk or stationarity in the individual price series by hypothesising unit roots in the levels and first differences of each price series using the KPSS test. We estimate the random walk with only a drift but without a trend because visually examining the graphical plot of the series in Figure 2 reveals the unlikelihood of a non-zero expected mean in the levels of the series. The plots however show no obvious persistent trending behaviour in the data. Therefore, we omit a deterministic trend but include a drift in both the KPSS test for unit roots and in the

[4] A reversal in direction of trade with imports flowing from Ghana to Burkina-Faso instead of the current pattern is theoretically possible.

Johansen's test for cointegration. The chosen lag lengths in both tests are based on the Akaike Information Criterion (AIC). The results of the unit root test are presented in Table 2.

From the above unit root results, we strongly reject the null hypothesis of no unit roots (i.e. the series is stationary) in the level of the prices series at the 1% and 5% significant levels, but cannot reject the null hypothesis of no unit roots at the first difference of the price series. Therefore, the series under study are (first) difference stationary processes [i.e. they have unit root or are I (1)]. Put differently, fresh tomato prices in the markets under study can be said to be non-stationary in their levels but stationary in their first differences. With the proof from the univariate analysis that the price series are non-stationary in their levels, we proceed to test for cointegration between the net producer/net consumer market pairs using the Johansen's maximum likelihood VAR approach. The results of the cointegration test between the market pairs are presented in Table 3.

Table 2. Results of KPSS unit root tests on the price series

Series	Test Statistics (Levels)		Test Statistics (First Diff.)	
	Statistic	No. of Lags	Statistic	No. of Lags
Po	1.412**	4	0.102	4
Tamale	1.969**	4	0.025	4
Techiman	2.134**	4	0.024	4
Kumasi	1.341**	4	0.035	4
Accra	2.890**	4	0.017	4

Notes: The asterisks ** and * denote rejection of the null hypothesis at the 1% and 5% significance levels. The respective critical values at the 1% and 5% significance levels are 0.347 and 0.463 for both the test at the levels and first difference of the series. The lag value of 4 is a suitable choice.

Source: Author's Own Computation

Table 3. Johansen's test of cointegration

Market Pair	Test Statistic (Trace Test)	
	Ho: r = 0	Ho: r = 1 No. of Lags
Po - Accra	26.40**	11.14* 2
Po - Kumasi	22.49**	6.46 2
Po – Techiman	27.76*	8.84 1
Po - Tamale	36.50**	13.58* 1
Techiman -Accra	56.34**	7.74 1
Techiman -Kumasi	34.75**	6.01 1
Techiman -Tamale	32.63**	8.74 1

Notes: The asterisks ** and * denote rejection of the null hypothesis of no cointegration vector at the 1% and 5% levels respectively. The critical values for r = 0 and r = 1 at the 1% and 5% significance levels are 24.69 and 12.53 and 20.16 and 9.14 respectively.

Source: Author's Own Computation.

The results provide evidence in favour of cointegration between the tomato market pairs under study. The null hypothesis of r = 0, implying an absence of a cointegration relationship between the producer and consumer markets is rejected for all the market pairs at both the 1%

and 5% significance levels. We cannot however reject the null hypothesis of one cointegrating relation (i.e. r = 1 between pairs of net producer/net consumer markets, especially at the 5% significance level). This means, there exists at least one stationary cointegration relation (r = 1) between the pairs of net producer and net consumer price series measured semi-weekly, and by implication in the tomato marketing system under study.

The findings imply that the series have a particularly strong link that is of interest from the economic point of view. It may be because similar stochastic processes, possibly induced by efficient information flow or seasonal effects, drive the behaviour of prices in the system of markets (MOTAMED et al, 2008). Therefore tomato prices in the producer and consumer markets do not drift apart in the long run. The proof of cointegration is also evidence for a common interstate tomato market between Burkina-Faso and Ghana, where inter-market prices adjust to achieve long-run, market equilibrium. Perhaps the seasonal nature of tomato production, with either the Po/Navrongo or Techiman market being a major source of supply to the other markets in the system per season, the effective network of traders between producer and consumer markets and recent improvements in roads, means of transportation and information flow via mobile phones, explain this outcome. Whatever the case may be, the evidence of at least one cointegrating relation between the market pairs provides an ideal setting for us to use the VECM to explore border and distance effects on price transmission and market integration between the selected markets.

4.2. Results of the Vector Error Correction Models

Having significant cointegrating vectors between the net producer and net consumer tomato market pairs is a necessary condition for using the VECM to determine the effects of price shocks on price adjustment. In this section, the results of the estimated standard VECM and the regime-dependent VECM are presented and their implications discussed. Both models are estimated by means of the reduced rank estimation procedure using the JMulti software.

4.2.1 Results of the Standard VECM

The results of the econometric estimation of the standard VECM which does not account for regime disparities are presented in Table 4.

The Table shows estimated speeds of price transmission and their corresponding half-lives for market pairs involving Po/Techiman and Accra, Kumasi and Tamale. Two forms of the speed of price transmission (adjustment) parameter are estimated in each VECM equation - α^s which measures the response to price shocks by the producer market Po/Techiman to correct disequilibrium and α^c measuring the price adjustment by the consumer markets to correct disequilibrium following shocks. The two parameters represent the dynamic interactions between Po/Techiman and Accra, Kumasi and Tamale pair-wise for the entire period of the study. The price adjustment half-lives λ^s and λ^c, for the producer and consumer markets respectively, measure the time required by the producer market (in the case of λ^s) or the consumer for (λ^c) to correct half of the deviations from equilibrium.

Table 4. Estimated speeds of price transmission and half-lives in the standard VECM

Market Pair	Price Adjustment Parameters and Half-lives			
	α^s	λ^s	α^c	λ^c
Po – Accra	-0.070***	9.55	-0.022 [-1.068]	-
Po – Kumasi	-0.080***	8.31	-0.009 [-0.474]	-
Po – Tamale	-0.070***	9.55	0.035 [1.700]	-
Po – Techiman	-0.071***	9.41	0.025 [1.247]	-
Aver. adjustment with Po[5]	-0.073	9.93	0.035	
Techiman - Accra	-0.097*** [-3.361]	6.79	0.094*** [3.213]	7.02
Techiman - Kumasi	-0.141*** [-5.284]	4.56	0.029 [1.166]	
Techiman - Tamale	-0.069*** [-3.658]	9.69	0.066*** [3.190]	10.15
Techiman - Po[6]	-0.011 [-1.247]	-	0.030 *** [3.682]	22.76
Aver. adjustment with Tech	-0.102	7.01	0.063	13.93

Notes: The asterisks *, ** and *** denote significance at the 10%, 5% and 1% level respectively. Half-lives, measured in semi-weeks, are only computed for significant speeds of price transmission (adjustment parameters).

Source: Authors' Own Computation.

From the estimates involving Po as the producer market, α^s, the significant adjustment made by Po to correct deviations from equilibrium range from -0.070 (7%) to -0.080 (8%), averaging 0.073 (7.3%). The corresponding adjustment, α^c, made by the consumer markets including Techiman[7] vary from -0.009 to 0.035. None of these are however significant. In addition, only the estimates for Po-Tamale (0.035) and Po-Techiman (0.025) have the correct sign.

The estimated half-lives, λ^s, for the significant adjustment speeds by Po range from 8.31 semi-weeks for Po-Kumasi, the market pair with the highest speed of adjustment, to 9.55 semi-weeks for Po-Accra and Po-Tamale, which have the lowest adjustment speed.??? Note that the estimated half-lives make economic sense for only significant adjustment speeds; hence no half-life estimates exist for λ^c when Po is specified as the producer market under the standard VECM.

When we specify Techiman as the producer market, then the significant estimated α^s range from -0.069 (6.9%) to -0.141 (14.1%), averaging -0.102 (10.2%), with half-lives between 4.56 for Techiman-Kumasi and 9.69 for Techiman-Tamale. The significant adjustment, α^c, by the consumer markets including Po (Navrongo as proxy), on the other hand range from 0.030 (3%) to 0.094 (9.4%) averaging 0.063 (6.3%) with associated half-lives between 7 and 22.76 semi-weeks.

The findings indicate that though dynamic price relationships do exists between Po in Burkinabe and the Ghanaian tomato markets, the degree of price transmission is generally

[5] Though averages here have no econometric importance, they are computed here just for comparative purposes.

[6] Actual, physical direct flow of tomato between Techiman and Po does not really exist. Po is therefore "proxied" by Navrongo which is located near the border to Burkina-Faso when Techiman is the producer market.

[7] Whenever Po is considered a producer market, Techiman theoretically becomes a consumer market. The reverse is true.

lower (about 7%) when the Burkinabe market (Po) is the production centre than when tomato is locally produced in Techiman, in which case the speed of price transmission averages about 10%. This means when there is no border between producer-consumer market pairs, or where the border effect between the markets is removed, the speed of price transmission would be expected to improve by about 3%. Similarly, price adjustment by the producer market to correct disequilibrium following shocks would be faster without border effects than with same. The evidence is that the time required to eliminate half of the market disequilibrium following shocks would reduce from 10 semi-weeks with Po as the producer market (5 weeks) to about 7 semi-weeks (3.5 weeks) when Techiman is the source of tomato supply to the marketing system.

We suspect that the higher adjustment speeds by both the producer markets with Kumasi (i.e. Po-Kumasi (0.080) and Techiman-Kumasi (0.141)) may be attributed to both distance and border effects. By removing the border, the dynamic interaction of the producer market price with the Kumasi price increases from 8% to 14%. Despite the fact that Accra is the biggest tomato consuming market in Ghana and is expected, *ceteris paribus,* to exert the highest influence on price dynamics in the producer markets, Kumasi, the second largest consumer market in Ghana, appears to be playing this role because of its closer proximity to both producer markets than Accra (see Figure 1).

There is also evidence to suggest, in line with Ihle et al 2010, that the producer market prices seem to adjust more significantly towards correcting disequilibrium than prices in the consumer markets.

In fact, none of the consumer markets significantly error corrects when the source of supply is across the border in Po. This means the consumer markets are largely weakly exogenous (i.e. only future price dynamics in these markets are mostly significantly influenced by existing dynamics in the producer markets). Overall, the results however demonstrate that price adjustment processes that correct deviations from long-run equilibrium relationships are bidirectional (i.e. prices on both markets tend to respond to deviations from their common equilibrium more in the absence of a border between markets than they do when there is a border separating the markets).

4.2.2 Results of the Regime-Dependent VECM

Even though the above results of the standard VECM paint a general picture of the nature of inter-market price dynamics with and without the international border between the markets under study, the standard VECM is limited in not being able to distinguish between the seasonally-dependent regimes of tomato supply to the marketing system. By accounting for the regime effects in the regime-dependent VECM, we may obtain further evidence to verify the border effects identified above. This is because, with the tomato supply sources to the markets under study switching seasonally, then estimating the VECM with the speeds of adjustments specified as regime-varying parameters may yield more economically interpretable results than those of the standard VECM. The results of the regime-dependent VECM are presented in Table 5.

The speed of price adjustment (α_1^s) by the producer market, namely Po, under the Po regime ranges from -0.060 (6%) to -0.109 (10.9%) averaging -0.088 (8.8%). Like under the standard VECM, the adjustment speeds by the consumer markets (α_1^c) that correct deviations from equilibrium under the Po regime show significance at the 1% for only Kumasi. This estimate, however, has the wrong sign and is not economically interpretable. The corresponding adjustment half-lives by the producer market (λ_1^s) range from 6 to 11.20 semi-weeks, averaging 8.89 semi-weeks. Even though the half-life for the adjustment by Kumasi (λ_1^c) has been computed to be 11.81 semi-weeks, like its corresponding speed of adjustment estimate, this value too does not have economic meaning.

Table 5. Estimated speeds of price transmission and half-lives in the regime-dependent VECM

Po Regime	Price Adjustment Parameters and Half-lives			
Market Pair	α_1^s	λ_1^s	α_1^c	λ_1^c
Po – Accra	-0.092** [-2.570]	7.18	-0.051 [-1.730]	-
Po – Kumasi	-0.060** [-1.967]	11.20	-0.057*** [-3.123]	11.81
Po – Tamale	-0.109*** [-3.147]	6.00	0.046 [1.134]	-
Po - Techiman	-0.092*** [-2.570]	7.18	0.035 [1.434]	-
Average Adjustment	-0.088	8.89	0.057	11.81
Techiman Regime	α_2^s	λ_2^s	α_2^c	λ_2^c
Techiman – Accra	-0.151*** [-5.041]	4.23	0.057** [1.920]	11.81
Techiman – Kumasi	-0.138*** [-3.905]	4.67	0.042 [1.285]	-
Techiman – Tamale	-0.050*** [-3.031]	13.51	0.064*** [3.603]	10.48
Techiman - Po	-0.016 [-1.797]	-	0.032*** [3.301]	21.31
Aver. Adjustment	-0.113	7.47	0.051	14.53

Notes: The asterisks ** and *** denote significance at the 5% and 1% level respectively. Half-lives, measured in semi-weeks, are only computed for significant speeds of price transmission (adjustment parameters). Student t-values are in the parenthesis.
Source: Author's Own Computation.

Under the Techiman regime, the significant speeds of price transmission from the producer market (α_2^s) range from -0.050 (5%) to -0.151(15.5%) averaging -0.113 (11.3%). Unlike under the Po regime, three of the consumer markets – Accra (0.057), Tamale (0.064) and Techiman (0.032) - now respond significantly with the expected sign to error correct towards equilibrium following shocks.

The average consumer-market speed of transmission (α_2^c) is 0.051 (5.1%). The half-lives of price adjustment by the producer market (Techiman) under this regime range from 4.23 to 13.51 semi-weeks, averaging 7.47 semi-weeks. The half-lives of adjustment parameters associated with the consumer markets range from 10.48 to 21.31 semi-weeks, averaging 14.53 semi-weeks.

The results of the regime-dependent VECM tell us that differences in the speeds of price transmission across the two regimes exist. These differences result in principle from changes in shocks to fresh tomato prices following shifts in the source of tomato supply from a local producer market that is averagely nearer to and not separated from the consumer markets by a border, to a foreign producer market across a border and averagely further from, especially the large consumer markets. Alternatively stated, the empirical results hint at the likelihood of border and distance as the major underlying factors affecting the rate of price transmission and the timeliness of price adjustment between the markets under study. The results are thus consistent with the nature of cross-border trade in the ECOWAS region.

As noted earlier, even though the ECOWAS protocol on trade excludes any form of trade restrictions on member countries, practical impediments to smooth arbitrage processes exist at the borders and partly mitigate the effects of the protocol. The results in this way are of high theoretical and practical relevance. For instance, they demonstrate the extent of the integration of tomato markets between Ghana and Burkina-Faso and in the ECOWAS trade block in general. This allows policy makers to adopt the necessary steps to promote cross-border trade in the sub-region.

CONCLUSION AND OUTLOOK

The results of the above analysis show that the semi-weekly prices of fresh tomato largely co-move over the period of the investigation. Overall, the net consumer markets under both the Burkina-Faso and non-Burkina-Faso regimes are weakly exogenous (i.e. they either do not react at all or react weakly to deviations from their long-run equilibrium with the net producer markets). The net producer markets on the other hand exhibit stronger and significant response to deviations from this equilibrium. This weak exogeneity of the consumer markets may be attributed to the use of market power by traders to create price transmission asymmetry – the transmission of, especially positive price changes at the farm gate to retail and consumer markets while withholding positive price changes in the consumer markets from reaching the farm gate.

It can also be concluded that an increase in geographic distance and presence of borders between markets appear to weaken the speed of price transmission between markets separated by the borders, all other things being equal,. This is because other determinants of price transmission viz. transaction costs will generally increase with distance, making arbitrage more costly and increasing the average time required to complete a transaction (Ihle et al, 2010). Also the crossing of borders by traders in low income countries usually involves formal (and sometimes informal) costs, as well as delays. This effect is suspected to be particularly important in the West Africa Sub-region due to high costs incurred by traders at borders, the cumbersome nature of customs procedures and little transparency and automation of such procedures.

In conclusion, some evidence of the link between international borders and distance on the one hand, and the speed of price transmission on the other, has been obtained using prices of tomato at the production centre at Burkina-Faso with prices on Ghanaian tomato markets. These results suggest that though borders in West Africa do not completely curtail cross-border trade, price transmission and consequently market integration, they nevertheless weaken these processes. Whether the observed reducing effects are solely border and distance dependent or whether other microdrivers including poor communication, marketing, transportation, exchange rate, language and other practical difficulties are involved, is however difficult to unravel.

What is clear from the findings is that more can be done to improve the speed of price transmission between Ghanaian and Burkinabe tomato markets. Such an improvement is necessary to ultimately lead to an enhanced welfare for tomato producers in both countries.

A methodological improvement of this analysis would be to extend the geographic coverage to consider more markets in Burkina-Faso and Ghana. Also, including in such analysis other factors such as concessionary trade conditions, not reflected in the prices would be useful.

REFERENCES

Abdulai, A. (2000). Spatial Price Transmission and Asymmetry in the Ghanaian Maize Market, *Journal of Development Economics* 63:327-349.

Aker, J. C. (2008). Information from Markets Near and Far: The Impact of Mobile Phones on Grain Markets in Niger**,** *BREAD Working Paper* No. 177.

Von Cramon-Taubadel, S. (1998). Estimating Asymmetric Price Transmission with Error Correction Representation: An application to the German Pork Market, *European Economic Review* 25: 1-18.

Von Cramon-Taubadel, S., Ihle, R., and Amikuzuno, J., (2009). Price Transmission in the Eastern Africa Maize Market, *Report prepared for the World Bank, mimeo.*

Eduardo, L. M. and A. M. Mario (2001). Seasonal Price Transmission in Soybean International Market: The Case of Brazil and Argentina, *Pesquisa and Debate, SP* 12 1(19): 92-106.

Goetz, L. and S. Von Cramon-Taubadel (2006). Considering Threshold Effects in the Long-run Equilibrium in a Vector Error Correction Model: An Application to the German Apple Market, paper presented in the 12th Congress of the European Association of Agricultural Economists (EAAE) 2008 "People, Food and Environments: Global Trends and European Strategy" held in Ghent-Belgium, 26-29.08.2008.

Goodwin B. K., T. J. Grennes and L. A. Craig (2002). Mechanical Refrigeration and the Integration of Perishable Commodity Markets, *Exploration in Economic History* 39: 154-82.

Ihle, R., J. Amikuzuno and S. von Cramon-Taubadel (2010), "L'intégration des marchés avec et sans échanges commerciaux directs: le cas de la tomate au Ghana", *Revue D'Économie Du Développement* 24(1): 21-46.

Ihle, R., S. Von Cramon-Taubadel, and S. Zorya (2010). Country and Border Effects in the Transmission of Maize Prices in Eastern Africa: Evidence from a Semi-Parametric Regression Model, paper presented at the Joint 3rd African Association of Agricultural Economists (AAAE) and 48th Agricultural Economists Association of South Africa (AEASA) Conference, Cape Town, South Africa, 19-23.09.2010.

Rin Africa (2009). GHANA: Plummeting profits drive tomato farmers to suicide http://www.irinnews.org/Report.aspx?ReportId=83980 Accessed on 09. March 2009.

Jensen, R. (2007). The Digital Provide: Information (Technology), Market Performance, and Welfare in the South Indian Fisheries Sector, *Quarterly Journal of Economics* 122(3): 879-924.

Lütkepohl H. and M. Krätzig (2004). Applied Time Series Econometrics, Cambridge University Press.

Meyer, J. and S. Von CRAMON-TAUBADEL (2004). Asymmetric price transmission: A survey, *Journal of Agricultural Economics* 55: 581-611.

Motamed M., K. A. Foster and W. E. Tyner (2008). Applying Cointegration and Error Correction to Measure Trade Linkages: Maize Prices in the United States and Mexico, *Agricultural Economics* 39: 29-39.

Parrott S. D., D. B. Eastwood and J. R. Brooker (2001). Testing for Symmetry in Price Transmission:An Extension of the Shiller Lag Structure with an Application to Fresh Tomatoes, *Journal of Agribusiness* 19 (1): 35 – 49.

Padilla-Bernal L., D. D. Thilmany and M. L. Loureiro (2003). An Empirical Analysis of Market Integration and Efficiency for U.S. Fresh Tomato Markets, *Journal of Agricultural and Resource Economics* 28(3): 435 – 450.

Stephens, E.C., Mabaya, E., Von Cramon-Taubadel, S. and C.B. Barrett (2008). Spatial Price Adjustment with and without Trade. Cornell University Working paper.

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EFFECT OF PRODUCTION ON LARGE CARDAMOM PRICE VARIABILITY IN NEPAL

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ABSTRACT

During the last decade, price variability was a serious problem of large cardamom industry in Nepal. The objective of this study was to investigate the effects of production of large cardamom in Nepal and India and small cardamom in India on large cardamom price volatility in Nepal. A GARCH specification was applied to analyze the price volatility by employing monthly wholesale price data from January 2001 to October 2009. As a result, the production volume of large cardamom in Nepal and India played a positive role in reducing price volatility, whereas small cardamom production in India played no effect on price volatility. The results are helpful to formulate mitigating measures for price volatility problems focusing on production variation.

Keywords: price volatility, GARCH specification, large cardamom, small cardamom

JEL Codes: C22, Q11, Q17

INTRODUCTION

As an agricultural commodity, large cardamom is the second most export earning industry next to lentil in Nepal. The export statistics show large cardamom production been primarily export oriented as about 90% of production volume has been exported to international markets. In 2008, large cardamom export valued US$21 million, which explained 2.2% of total export earnings (MoCS 2010) in Nepal. However, the price of large cardamom was highly volatile during the last decade. The average monthly price of large cardamom fluctuated from a height of US$ 3.86 per kilogram in January 2001 to as low as

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US$ 1.57 in May 2005, and rose to UD$ 5.08 in October 2009 (AEC 2010). Consequently, small-farmers growing large cardamom were seriously affected by high fluctuations in price.

Most studies on price volatility of agricultural commodities concentrated on the effect of government policies (Gemech and Struthers 2007; Hudson and Coble 1999), seasonality (Hudson and Coble 1999; Kenyon et al. 1987), and futures contract maturity (Milonas 1991). Most importantly, these studies applied major commodities such as coffee, cotton, corn, soybean, wheat, and so on for price volatility analysis (Kenyon et al. 1987; Milonas 1991; Hudson and Coble 1999; Sarris 2000; Gemech and Struthers 2007). Unfortunately, the studies have hardly covered the supply effect of substitute commodity on price volatility. Very few studies have utilized quantity effect on price volatility. Kenyon et al. (1987) analyze the supply quantity effect on corn and soybean in futures markets. Shively (1996) describes the effect of domestic and regional production of maize on price variability in Ghana. However, understanding the effects of supply of own and substitution commodities on price volatility has not been studied so far and remained unclear.

For this study, production is assumed an important factor for price variation in Nepal. Therefore, large cardamom production in Nepal is considered as an important factor. Moreover, the production of large and small cardamom in India is also taken as an important factor for two reasons. Firstly, India is the largest market of large cardamom of Nepal. Production in Nepal and India share a common market in India, thus, the production situations in India are also important to analyze production effect precisely. Secondly, small cardamom is a substitute commodity for large cardamom (Varadarasan and Biswas 2002; CAP 2009) and India herself is a major supplier of small cardamom. Therefore, considering the effects of large cardamom production in Nepal and India and small cardamom production in India, the objective of this study was to investigate the large cardamom price volatility in Nepal.

This paper has been organized as follows. Firstly, cardamom in Nepal is presented along with price and production situations. Empirical methodology for GARCH specification, data sources and diagnostic are explained. The empirical results, conclusions, and policy implications are presented in the subsequent sections.

LARGE CARDAMOM PRODUCTION AND PRICE IN NEPAL

Large cardamom is considered a high potential industry for rural poverty reduction in Nepal; consequently, some Non-Governmental Organizations (NGOs) and Government of Nepal have been trying to explore the large cardamom industry for poverty reduction. The industry contributes to self-employment generation to the small farm families by accounting higher share of labor in inputs for the production process. Moreover, large cardamom is convenient for transportation because of its low volume and high value. Therefore, this industry is gaining popularity among the small hold hilly farmers in Nepal. Accordingly, this crop has expanded to 39 districts out of the 75 districts of Nepal.

In recent years, both planted area and production have rapidly increased in Nepal (George, Munankami, and Bijl 2007). The planted area of large cardamom was 8,782 hectares in 1994/95; it grew to 11,486 hectares in 2005/06, and reached to 14,370 hectares in 2008/09. On average, it increased by 400 hectares annually. Similarly, the production expanded more than twice during the same period from 3,010 m.t. in 1994/95 to 7,037 m.t. in 2008/09 with 288 m.t. annual growth on average (George, Munankami, and Bijl 2007; MoAC 2010).

Large cardamom in Nepal showed fluctuating pattern of production during the last nine years because of weather factors (rainfall and temperature) and viral diseases (Chapagain 2011). Similar to Nepal, cardamoms in India also showed production variation during the same period (Narayanan 2004).

Table 1 shows the descriptive statistics of production of large cardamom in Nepal and production of cardamoms in India. The statistics in Table 1 present 12 per cent coefficient of variation for large cardamom production in India and 8 per cent for both large cardamom production in Nepal and small cardamom production in India, which illustrate the existence of the production variation in both cardamoms.

Table 1. Descriptive Statistics of Cardamoms Production (m.t.) in Nepal and India, 2000-2008

	Large cardamom in Nepal	Large cardamom in India	Small cardamom in India
Mean	6458.78	5220.56	11222.67
Standard Deviation	520.13	648.11	871.79
Skewness	0.40	-0.21	-0.76
Kurtosis	-1.99	-0.86	1.46
Coefficient of variation	0.08	0.12	0.08

Source: Authors' estimation based on data.

As explained before, price of large cardamom is highly variable within and between the years. For instance, the nominal average wholesale price in Birtamod[1] market was US$ 4.29 per kg in 2000, 1.57 per kg in 2005, and 3.57 per kg in 2007. Moreover, the price of large cardamom started to increase sharply from October of 2009.

Price fluctuation during 2001 to 2009 was not realized only in Birtamod market, but also in other markets in Nepal and India[2] (George, Munankami, and Bijl 2007; AEC 2010; Spice Board India 2010).

In addition, the price fluctuation of large cardamom is reasonably higher compared to the prices of other important crops in Nepal. Figure 1 shows the price of large cardamom as highly unstable compared to medium fine rice and processed lentil.[3]

[1] A main market center for large cardamom in Nepal.

[2] Birtamod and Dharan are the main markets for large cardamom in Nepal, whereas Delhi, Silguri, and Gantok are in India. Birtamod is the biggest market in Nepal followed by Dharan for collection and export to India. Similarly, Delhi is the biggest market in India for consumption and exportation. Silguri is the biggest market for collection and distribution to different markets in India, and Gantok is collection market near production area in India.

[3] Rice is the major food crop and lentil is the first exporting agricultural commodity by value. The price of large cardamom is the monthly wholesale price in Biratnagar market, whereas prices of rice and lentil are monthly national average retail prices. Although direct comparison is difficult, however, it gives some knowledge of fluctuation patterns.

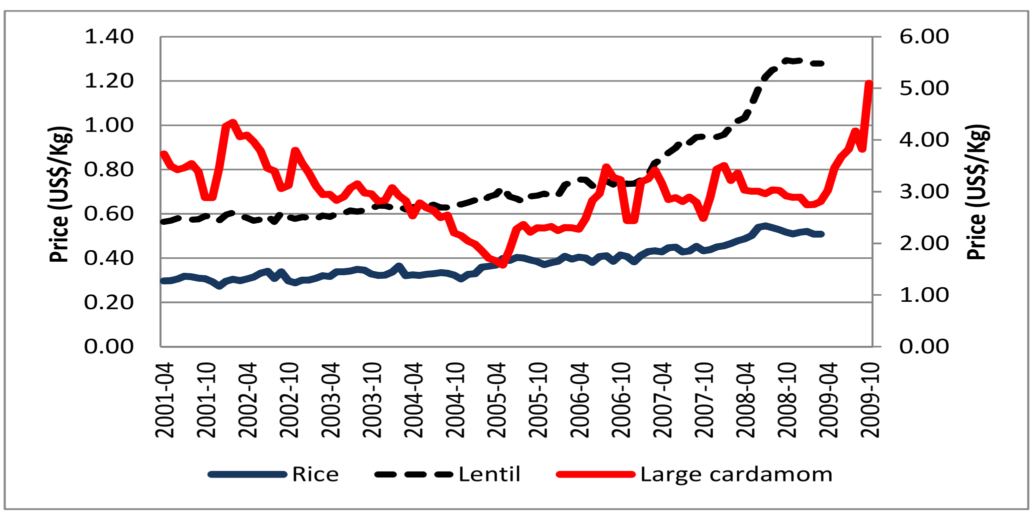
SUPPLY EFFECT ON PRICE VOLATILITY

Production of cardamoms is an important factor to induce large cardamom price volatility in local as well as in global markets. The current price of large cardamom in Nepal specially depends on current production in Nepal and India.

It is assumed that the production is independent and identically distributed (i.i.d.) and demand is non-stochastic. In absence of storage, an equilibrium condition is accounted by $D(p_t) = y_t$. Thus, the current period price is determined by the current supply and prices are also i.i.d. In the case of storage possibility, however, the price is determined by current supply and inventory. Consequently, the probability distribution of prices will no longer be i.i.d.

Under rational expectations, price volatility is explained through storage model. Price volatility in commodity can be resulted from the inherent nonlinearity introduced by the inability of market to carry negative stocks (Gustafson 1958). An increase in price induces inventory holders to sell their stocks and thereby leads to more volatility in the subsequent period prices. Conversely, a low price causes the inventory holders to keep or increase inventories and results less volatility in subsequent period prices. Therefore, large cardamom prices in Nepal are serially correlated and the current price is influenced by past ones through inventories.

Moreover, storage and export demands compete for large cardamom in Nepal. A high current production in India reduces the incentives to export due to low price in India. Consequently, stockholders in Nepal increase their inventory, which causes to reduce the price volatility. However, if the production in India is lower, the reverse is true. Thus, we consider price volatility of large cardamom in Nepal is influenced by the current production in India. Similar justification applies to small cardamom because it is considered a substitute commodity for large cardamom.



Source: Original data from Agribusiness Promotion and Marketing Development Directorate (ABPMDD), 2008-2009, and Agro Enterprises Centre (AEC), 2010.

Figure 1. Price Trends of Medium Fine Rice, Processed Lentil (left Y-axis), and Large Cardamom (right Y-axis) in Nepal, 2001- 2009.

ECONOMETRIC MODEL

Price variability in agricultural commodities has been widely modeled as autoregressive conditional heteroscedasticity (ARCH) and general autoregressive conditional heteroscedasticity (GARCH) (Aradhyula and Holt 1988; Shively 1996; Gamech and Struthers 2007; Ghoshray 2010). The ARCH model was first introduced by Engle in 1982. Later, Bollerslev (1986) extended the model as GARCH by introducing lag variance terms in the variance equation. The GARCH model provides a framework for the estimation of conditional mean and variance over time, where the test of hypothesis regarding non-constant conditional variances is done.

The general specification of GARCH process is

$$(1)\ y_t = \mu_0 + \sum_{l=1}^{k} \mu_l y_{t-l} + \varphi' X_t + \varepsilon_t$$

$$(2)\ h_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i \varepsilon_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j} + \psi' Z_t$$

Equations (1) and (2) describe the conditional mean of the process over time and the evolution of the conditional variance equations, respectively. Moreover, equation (1) is used to explain the large cardamom price levels conditional on X_t, whereas equation (2) is used to explain the price variance conditional on Z_t. In this study, X_t denotes the matrix of predetermined trend variables, and measures production of cardamoms; whereas Z_t contains predetermined variables that influence the residual variance. Here, ε_t is error term and μ_0, μ_l, α_0, α_i, β_j, φ', and ψ' are parameters.

In this study, a GARCH specification was applied to capture the possible effects of the production factors on large cardamom price volatility. Large cardamom production in Nepal and India and small cardamom production in India were applied as production variables.

$$(3)\quad p_t = \mu_0 + \mu_1 p_{t-1} + \varphi_1 L_t^N + \varphi_2 L_t^I + \varphi_3 S_t^I + \varepsilon_t$$

$$(4)\quad h_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i\, \varepsilon_{t-i}^2 + \sum_{j=1}^{p} \beta_j h_{t-j} + \psi_1 L_t^N + \psi_2 L_t^I + \psi_3 S_t^I$$

$$(5)\quad \sum_{i=1}^{q} \alpha_i + \sum_{j=1}^{p} \beta_j < 1$$

$$(6)\ \alpha_0 > 0,\ \alpha_i \geq 0,\ \beta_j \geq 0$$

In equation (3), p_t is the first difference of the log of price of large cardamom in Nepal, p_{t-1} represents its first lag value and ε_t is the error of p_t. Large cardamom production in Nepal (L_t^N), large cardamom production in India (L_t^I), mall cardamom production in India (S_t^I), and p_{t-1} are considered the influencing factors for mean price of large cardamom in Nepal. Price volatility is represented by the conditional variance (h_t) in equation (4), which is specified as a linear function of lagged square error, lagged value of conditional variance, large cardamom production in Nepal (L_t^N), large cardamom production in India (L_t^I), and small cardamom production in India (S_t^I). Here, α_i and β_j are parameters for ARCH and GARCH terms, respectively. Additionally, μ_0, α_0, μ_1, φ_1, φ_2, φ_3, ψ_1, ψ_2, and ψ_3, are parameters.

Equation (5) describes the sufficient condition for volatility decay over time as the sum of the coefficients is less than unity; however, the constraints in equation (6) imply for strict positivity of conditional variance.

DATA

The study utilized the monthly wholesale price data of large cardamom obtained from the Agro Enterprises Centre (AEC) under the Federation of Nepalese Chambers of Commerce and Industry (FNCCI), Nepal. Monthly wholesale prices were taken in Birtamod, the biggest market for large cardamom in Nepal. The price data were converted to UD\$[4]. Large cardamom prices constitute uninterrupted time series data from January 2001 through October 2009. Similarly, large cardamom production data in Nepal were collected from Ministry of Agriculture and Cooperatives, whereas large and small cardamom productions in India were obtained from a website of Spice Board of India under the Ministry of Commerce and Industry. The large cardamom price and production data were expressed in US$ per kg and m.t., respectively.

The study utilized the nominal price rather than the consumer price index (CPI) deflated price for the analysis. Goodwin (2000) pointed out the limitations of deflating by CPI and applied the nominal price in a similar study. The price series was augmented[5] with the annual production data of large cardamom and small cardamom produced in Nepal and India as applied by Shively (1996).

The study examined some diagnostic tests for time series data before proceeding actual GARCH analysis. Generally, the unit root test is done to test whether the series is stationary. Dickey Fuller (DF) unit root test (Dickey and Fuller 1979) is well established in recent literatures.

A generalized model of Dickey Fuller -- Augmented Dickey Fuller (ADF) unit root test was employed in this analysis. The test results rejected the null hypothesis of a unit root of large cardamom price series with a sufficient negative value[6] at first difference and random

[4] The conversion rate taken for the analysis is 1UD$= 70NRs, which is an almost average conversion rate.

[5] Annual production statistics were scrutinized with monthly price series. For the sake of production effect, new productions were applied from November for both cardamoms, because more than 90% large cardamom harvesting is done through August to October in both countries, whereas October to November is the peak harvesting period in case of small cardamom.

[6] The ADF unit root test result of log price series was -8.4170, against the critical values -4.0495, -3.4540, and -3.1526 at 1%, 5%, and 10% significance level, respectively.

walk with drift. Furthermore, partial autocorrelation results indicated that there is no autocorrelation in the series, and autoregressive process can be preceded. In addition, Lagrange Multiplier (LM) test was conducted to test ARCH effect in the equations. The LM results rejected[7] the null hypothesis of homoscedastic conditional variance in favor of ARCH effect at second lag based on ARCH (1) specifications.

RESULTS AND DISCUSSION

After a preliminary analysis of GARCH (p, q) specifications based on an Akaike information criterion, GARCH (1, 1) was found to be a good approximation for data generation process for equation (4).

In Table 2, the mean equation showed the inefficient results. The inefficient results may be due to the presence of ARCH effect (Shively 1996). However, results from the variance equation were quite interesting. The coefficients of the GARCH, large cardamom production in Nepal and India were statistically significant. With the assumption made in equation (6), the signs of the coefficients were correct.

Table 2. Maximum Likelihood Estimates of GARCH Model

Independent Variables	Estimated Values	z-Statistics	p-Value
Mean Equation			
Constant	-0.2471 (0.4556)	-0.5423	0.5876
Lagged price	0.1237 (0.1271)	0.9731	0.3305
Large cardamom Nepal	1.96×10^{-05} (3.79×10^{-05})	0.5169	0.6052
Large cardamom India	-7.36×10^{-06} (2.66×10^{-05})	-0.2764	0.7822
Small cardamom India	1.42×10^{-05} (1.19×10^{-05})	1.1936	0.2326
Variance Equation			
Constant	0.0026 (0.0043)	0.6038	0.5459
ARCH(1)	0.1505 (0.1330)	1.1314	0.2579
GARCH(1)	0.5997 (0.2734)	2.1928**	0.0283
Large cardamom Nepal	$- 8.95 \times 10^{-07}$ (3.56×10^{-08})	-25.1382***	0.0000
Large cardamom India	$- 1.45 \times 10^{-06}$ (8.23×10^{-07})	-1.7672*	0.0772
Small cardamom India	1.14×10^{-06} (7.2×10^{-07})	1.5896	0.1119
Log likelihood	110.6447		
N	105		

The numbers in parentheses are standard errors.
***, **, and * indicate significance at 1%, 5% and 10% level, respectively.

The coefficient of lag conditional variance (GARCH) is 0.60 and is significant at 95% confidence level, implying that lag variance had a higher influence on the variance at time t. Moreover, the sum of estimated ARCH and GARCH coefficients is 0.75, which is sufficiently lesser than unity, implying the volatility of price was not persistent.

[7] LM test results for F statistics was 3.3727 at (2, 99) degrees of freedom and χ^2 statistics was 6.650 at 2 degrees of freedom. Both test statics were significant at 5% level.

Most importantly, the variance equation results indicate that the coefficient of large cardamom production in Nepal has a negative sign and is significant at 99% confidence level. Similarly, the production coefficient of large cardamom in India also shows a negative sign and is significant at 90% confidence level. The findings indicate that price volatility of large cardamom is negatively influenced by the production of large cardamom in Nepal and India during sample period. The production volume of large cardamom in Nepal and India presents a good explanatory power to the large cardamom price volatility in Nepal. The results are explained by the storage model, where higher (lower) stocks reduce (increase) the price volatility of commodities. The findings are consistent with Shively (1996) who reported the production of maize in Ghana and in neighboring countries negatively influence price volatility in Ghana.

The result shows the coefficient of small cardamom production in India is positive; however, it is insignificant even at 90% confidence level. It could be the presence weak of substitutability between these two cardamoms. The result indicates production of small cardamom in India plays no role in price variation of large cardamom in Nepal.

Outcomes described by GARCH results may naturally be helpful to describe the price volatility situation of large cardamom in Nepal. Moreover, it can be useful to enhance market performance and reinforce government efforts towards price volatility problem. The results may not depict other crops; however, it will stimulate more empirical work in this exciting and important direction.

CONCLUSION

The price of large cardamom in Nepal showed a greater fluctuation compared to other important crops such as rice and lentil during the last decade. Supply factors were taken as major price volatility influencing factor for analysis. With the assumption of substitutability of large cardamom for small cardamom, cardamoms produced in Nepal and India shared common market. The collective supply of cardamoms influences market price in Nepal because prices in India and Nepal are highly correlated.

The study employed the GARCH (1, 1) model to examine the price volatility of large cardamom in Nepal. The results from GARCH analysis derive some general conclusions. First, the volatility of large cardamom price in Nepal at period ‘t’ is highly influenced by one lag variance; and the volatility is not persistent for longer period. The production volume of large cardamom in Nepal and India has a negative effect on price volatility. The result can be justified from the storage model; as higher (lower) stock decreases (increases) the price volatility. Similar results were observed by Shively (1996) in maize in Ghana. As a result, the second conclusion is the production of large cardamom from both Nepal and India played a positive role in reducing the large cardamom price volatility in Nepal. In contrast, the coefficient for small cardamom showed positive sign but insignificant result, implying that it has insignificant effect on price volatility of large cardamom in Nepal.

The results provide knowledge on the significance of production of large cardamom in Nepal and India for price variability in Nepal. Thus, policy maker can formulate price volatility mitigating measures taking the production of large cardamom into consideration. Based on the findings, some mitigating policies are suggested. Production fluctuation is the main cause of price variations in Nepal. Thus, some measures like disease and pest prevention and control and irrigation facility provision can be good solutions to minimize large cardamom production fluctuations in Nepal. In case of price fluctuation due to large cardamom production fluctuation in India, however, a price forecasting system based on demand and supply can be a helpful instrument.

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REFERENCES

Agribusiness Promotion and Marketing Development Directorate (ABPMDD). 2008-2009. “Agricultural Marketing Information Bulletin, Special Issue 2008-2009.” Department of Agriculture, Ministry of Agriculture and Cooperatives, Government of Nepal. http://www.agribiz.gov.np/download.html (accessed October 9, 2010)

Agro Enterprises Centre (AEC). 2010. “Monthly Prices of Agricultural Commodities 2001-2010.” Federation of Nepalese Chambers of Commerce and Industry (FNCCI), Nepal.

Aradhyula, S.V., and M.T. Holt. 1988. “GARCH Time Series Model: An Application to Retail Livestock Process.” *Western Journal of Agricultural Economics* 13: 365-374.

Bollerslev, T. 1986. “Generalized Autoregressive Conditional Heteroscedasticity.” *Journal of Econometrics* 31: 307-327.

Center for Agricultural Policy with Prosperity Initiative (CAP). 2009. “Small-scale Review of Cardamom 2009.” Center for Agricultural Policy with Prosperity Initiative (CAP), Vietnam. http://prosperityinitiative.org/images/stories/cardamom-small-scale_reporpdf (accessed November 5, 2010)

Chapagain, D. 2011. "Assessment of Climate Change Impact on Large Cardamom and Proposed Measures in Eastern Hill in Nepal. Paper Submitted to National Adaptation Program of Action (NAPA) Project Office, Ministry of Environment, Government of Nepal, Singhadarbar, Kathmandu, Nepal. climatenepal.org.np/main/downloadFile.php?fn=l5smn4oiz1e.pdf (accessed May 10, 2011)

Dickey, D.A., and W.A. Fuller. 1979. “Distribution of Estimators for Autoregressive Time Series with a Unit Root.” *Journal of American Statistical association* 74: 427-431.

Engle, R.F. 1982. “Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation.” *Econometrica* 50: 987-1008.

Gemech, F., and J. Struthers. 2007. “Coffee Price Volatility in Ethiopia: Effects of Market Reform Programmes.” *Journal of International development* 19: 1131-1142.

George, C.K., R. Munankami, and B. Bijl. 2007. “Nepal Advisory Service on Export Development of Priority Sectors: Sector Study on Large Cardamom.” International Trade Center, UNTAD/WTO. www.intracen.org/atf/nepal/Docs/cardamom-final.pdf (accessed October 10, 2010)

Ghoshray, A. “The Extent of World Coffee Market.” *Bulletin of Economic Research* 62 (2010): 97-107.

Goodwin, B.K., M.C. Roberts, and K.H. Coble. 2000. “Measurement of Price Risk in Revenue Insurance: Implications of Distributional Assumptions.” *Journal of Agricultural and Resource Economics* 25: 195-214.

Gustafson, R.L. 1958. “Carryover Levels for Grains: A Method for Determining Amounts That are Optimal under Specified Conditions,” Technical Bulletin 1178, USDA, Washington DC.

Hudson, D., and K. Coble. 1999. “Harvest Contract Price Volatility for Cotton.” *The Journal of Futures Markets* 19: 717-733.

Kenyon, D., K. Kling, J. Jordan, W. Seale, and N. McCabe. 1987. “Factor Affecting Agricultural Futures Price Variance.” *Journal of Futures Markets* 73: 73-91.

Milonas, N.T. 1991. “Measuring Seasonalities in Commodity Markets and the Half-Month Effects.” *The Journal of Future Markets* 11: 331-345.

Ministry of Agriculture and Cooperatives (MoAC). 2010. “Statistics on Nepalese Agriculture Sector 2000-2008.” Government of Nepal, Singhadarbar, Kathmandu, Nepal. http://www.moac.gov.np (accessed October 9, 2010).

Ministry of Commerce and Supply (MoCS). 2010. “Nepal Trade Integration Strategy: Executive Summary and Action Matrix.” Government of Nepal, Singhadarbar, Kathmandu, Nepal.

Narayanan, S. 2004. “Commodity Prices and Their Effects on Investment Viability and Repayment: The Case of Indian Cardamom,” Occasional Paper 30, 2004, Department of Economic Analysis and Research, National Bank for Agriculture and Rural Development, Mumbai, India.

Sarris, A.H. 2000. “Has World Cereal Market Instability Increased?” *Food Policy* 25: 337-350.

Shively, G.E. 1996. “Food Price Variability and Economic Reform: An ARCH Approach for Ghana.” *American Journal of Agricultural Economics* 78: 126-136.

Spice Board of India. 2010. “Production of Major Spices 2000-2008”. Ministry of Commerce and Industry, Government of India. http://www.indianspices.com (accessed October 9, 2010)

Varadarasan, S., and A. K. Biswas. 2002. “Large Cardamom *(Amomum Subulatum Roxb.).”* Cardamom, edited, P. N. Ravindran and K. J. Madhusoodanan. CRS Press.