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ADDITIONALITY OF CREDIT GUARANTEES FOR WHEAT EXPORTS

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ABSTRACT

Credit guarantees play an integral role in world trade. These guarantees are important marketing tools in the world wheat trade, both to compete in existing markets and to develop new markets. This study estimates additionality from export credit guarantees in the international wheat market. A conceptual framework is developed to distinguish between subsidy and non-subsidy aspects of guarantees. Empirical models were estimated using data pooled across importing countries and models were estimated for the principal exporting countries providing export credit guarantees: the United States, Canada, and France. The results indicate that U.S. and Canadian credit guarantees have opposing effects of similar magnitudes on U.S. exports. They also suggest that Canada's guarantee program has done more to displace U.S. sales than it has increased Canadian sales. The results suggest prospective benefits to the U.S. from multilateral reductions in subsidy levels.

Key Words: *Additionality, Export Credit Guarantees, Price Subsidy, GSM-102, EEP, Canada, United States, France.*
JEL Classification: F3

Credit guarantees are important tools in world trade and for wheat trade in particular. Exporting countries use guarantees to develop new markets and to compete in existing markets. Governments of exporting countries typically assume the default risk of importing countries when offering export credit guarantees. This has the effect of reducing the importer's cost of financing and may increase trade. Defaults represent an expected cost to

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the guarantor and have raised questions concerning the effectiveness of guarantee programs. One justification for these programs is that additional grain is sold when guarantees are provided. Additionality is measured as the change in the volume of trade associated with the credit guarantee (Smith and Ballenger 1989). Guarantees are an important component of the export programs in the United States, Canada, and the European Union (EU). However, when multiple countries offer credit, additionality in selected markets may be reduced.

The effectiveness of credit guarantee programs is an important issue confronting policymakers. Programs used by the United States have been scrutinized, both domestically and abroad [U.S. Government Accounting Office (GAO) 1997] and have been the subject of multilateral trade negotiations. However, the extent that credit guarantees produce additionality is not without dispute (Gudger 1998, pp. 11, 75).

WTO negotiations pertaining to agriculture have focused extensively on export subsidies and competition which were thought to be a "pillar" of discussions (WTO 2004). This pillar covers export subsidies (such as the U.S. Export Enhancement Program or EEP), all credit programs (such as the U.S. General Sales Manager Guarantee Programs, GSM-102 and GSM-103), and state trading enterprises or STEs. Specifics for resolution are still slated for further negotiation and need an implementation or ending date.

Producers, exporters, and importers are concerned with the effectiveness of guarantee programs, as they are the principal beneficiaries of increased sales. Guarantees may alleviate importers' credit constraints or allow exporters to charge a lower interest rate, thereby lowering the cost of financing and potentially expanding exports. Estimating additionality requires the subsidy amount be quantified and included in the analysis. The objective of this study is to derive estimates of additionality of credit guarantees on wheat exports. Import functions are estimated using pooled data for a group of importers that have been recipient countries of credit guarantees for wheat from the United States, Canada, and France. Results provide measures of additionality from credit guarantees for each of these wheat exporters' programs. In addition, we evaluate the effect of guarantee subsidies versus direct price subsidies on imports, and we compare the effectiveness of guarantee subsidies across competitor countries.

BACKGROUND

The United States, France, Canada, Australia, and some smaller wheat exporting countries each have some form of an export credit program. A government guarantee relieves exporters' banks of the risk that an importer will default. Guarantees are used widely by importing countries, due to the high cost of alternative financing. Importers incur financing fees to cover administration costs, but guarantees often provide an implicit subsidy to the importing country. The most popular programs are government-sponsored guarantees of private loans. Harris (1990), Dahl et al. (1995), Ray (1995), and World Perspectives, Inc. (1995) provide comprehensive reviews of these programs.

The Organization for Economic Cooperation and Development (OECD 2000) examined the history of programs for member nations from 1995-1998. They found that the United States, Canada, Australia, and the EU were the largest extenders of export credits, that use of export credits had increased from 1995 to 1998, and that these countries also had the largest subsidy rates for export credits. The OECD then sought to facilitate constructive disciplines on exporting countries about their credit programs, but would have been unable to enforce any conclusion.

Several aspects of the WTO agriculture negotiations are relevant to this study. In the August 2004 framework the members agreed they would like to curb (and ultimately eliminate) subsidy amounts (WTO 2004). Similarly, there is concern that credit programs contain subsidies and these should be controlled. The length credit is extended was proposed to be capped at 180 days. In addition, fees should be market-based. The exceptions raised for credit programs focus on their inherent ability to serve the needs of countries facing financial difficulties. Finally, the main issue surrounding STEs is transparency. When prices or credit terms are not observed, they may contain a subsidy.

While the negotiations continue, WTO policies and rulings have directly affected U.S. programs. In a ruling on U.S. cotton exports to Brazil, U.S. credit guarantees were classified as subsidies and therefore restricted in the same manner as direct price subsidies (Hudson et al. 2005). The reason for the classification was the cap or limit on fees (under U.S. law) charged under the GSM programs and the lack of a risk-based fee structure (Dispute Settlement 267). In response, U.S. Agriculture Secretary Johanns indicated that fees under GSM-102 would be assessed on a risk-based formulation. He also announced that no new applications would be accepted under GSM-103, although no specific mention of the terms were brought out in the settlement (USDA-FAS News Release, 2005). Further, Johanns forwarded legislative proposals to lift the cap on fees and eliminate the GSM-103 program entirely.

There are several motives for offering credit guarantees. These include increasing sales by relaxing an importer's foreign exchange constraint (Smith and Ballenger 1989), supporting specific sectors of an economy and correcting market failures (Raynauld 1992), and competing with other guarantors (Baron 1983). An importer's valuation of a credit guarantee affects their response and, ultimately, additionality. If credit simply relieves exchange shortfalls or reduces short-run debt servicing difficulties, additionality might be limited (Eaton 1986, p. 137). In some cases, the guarantee benefit may not be transmitted to the importer (e.g., because it is captured by the importer's bank), negating additionality. Thus, the extent of additionality in each import market is an empirical question.

A more prominent export strategy used by the United States during the study period was the EEP. As a direct price subsidy, the EEP functioned differently from any subsidies implied in credit guarantees. Numerous studies addressed the EEP, with some focusing on issues of additionality. Ackerman and Smith (1993) discussed factors that influence additionality attributable to EEP. Gardner (1994), though skeptical about the actual effects of improving U.S. wheat exports, attributed the success of the 1993 General Agreement on Tariffs and Trade (GATT) negotiations in reducing agricultural subsidies to the EEP. The effect of the EEP has been analyzed using different approaches with two general sets of results. Studies by Makki, Tweeten, and Miranda (1996), Chambers and Paarlberg (1991), Koo and Karemera (1991), and Seitzinger and Paarlberg (1990) show that the EEP had limited effects. In contrast, those by Johnson and Wilson (1995), Bailey (1989), and Haley (1989) report that the EEP had considerable positive effects on barley and wheat exports from the United States.[3]

[3] This is not intended to be an exhaustive survey of EEP studies. Patterson, Abbott, and Stiegert (1996) analyzed the impact of the EEP on firm-level entry decisions in world poultry, wheat, and wheat flour markets. Studies,

Wang and Sexton (2004) examined how the United States sets up the EEP bonus structure and suggested that any inefficiency in transferring the subsidy to importers would likely result in less additionality. None of these studies assessed the effects of credit guarantee programs or considered interactions with other export programs or those of competing exporters. If credit was addressed at all, it was treated simply as a price reduction or dummy variable.

Other studies addressed the impact of credit on demand (e.g., Benson and Clay, 1998). The effects of credit programs have been modeled with limited attention to the value of the guarantee to importers. Quantities imported under GSM-102 were used by Fleming (1990) to analyze rice demand. Haley (1989) analyzed the Commodity Credit Corporation (CCC) guarantees subsidy in a trade flow model. The subsidy used was the proportion of claim to the volume of loans guaranteed over the four years before 1986-1987. Koo and Karemera (1991) included a dummy variable for credit sales and the EEP when modeling wheat trade flows. Using dummy variables implies shifts in demand attributable to credit and EEP sales. Skully (1992) treated the guarantee subsidy as a price discount or pure price subsidy. Yang and Wilson (1996b) used a multinomial logit model to derive the marginal effects of changes in loan volume under CCC guarantees. They found the elasticity of own credit was significant, but declined with the introduction of the EEP in the late 1980s. Yang and Wilson (1996a) found that allocations are also influenced by competitors' credit sales.

Dahl, Wilson, and Gustafson (1999) show that the size of the guarantee subsidy depends upon the repayment terms of the loan, the size of the loan guaranteed, and importer risk of default. Guarantee subsidies, along with other subsidies, from EEP and PL-480, may allow sellers to price discriminate and Vercammen (1998) showed that a seller can use credit guarantees to price discriminate. Diersen and Sherrick (2005) developed a framework to model different aspects of managing a portfolio of guarantees and measured guarantee values for a cross-section of GSM recipients.

A major problem in analysis of credit guarantees is that there is limited information about terms of credit offerings in the private market. Few comparable short-term loans are made to these countries and terms are usually not reported (Raynauld 1992; Seilor 1990; Baron 1983). Thus, little information exists on alternative measures of the importer specific costs of borrowing for guaranteed loans. Hyberg et al. (1995) estimated the market interest rate for importers by mapping institutional investor country risk ratings to yields on a subset of countries that receive Moody's credit ratings. Using the relation between country risk ratings and yields, premiums were estimated for all GSM recipients. While Hyberg et al. (1995) recognized the need to measure additionality, their focus was to estimate the subsidy. Previous studies alluded to the effect that guarantees have on importer behavior. Guarantees may relax foreign exchange shortfalls, may result in better financing terms, and may be seen as equivalent to a price subsidy. However, none of these studies have shown how a credit guarantee influences imports. In this article, implicit guarantee subsidies are defined and used to evaluate their impact on imports.

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such as Roberts and Whish-Wilson (1993), analyze the implications of the U.S. EEP on wheat prices in targeted markets and, consequently, on the Australian wheat industry. Alston, Carter, and Smith (1993) compared the effects of export subsidies and production subsidies and evaluated export subsidies as a policy tool. Anania, Bohman, and Carter (1992) found evidence that subsidized exports replaced commercial exports. Bohman, Carter, and Dorfman (1991) analyzed the welfare effects of targeted export subsidies using a general equilibrium framework.

CONCEPTUAL FRAMEWORK AND EMPIRICAL ANALYSIS

Credit guarantees may result in an implicit subsidy which can be inferred as a source of savings for the importer. A conceptual model is used to illustrate the impacts of a direct subsidy and a guarantee on consumption. The procedure to measure the subsidy levels is described to provide an estimate of the benefits that could result in additionality. The data and unique aspects of guarantee subsidies that affect the empirical specification are then shown.

Conceptual Analysis of Additionality

It is necessary to examine how credit influences consumption to describe how export credit guarantees may generate additionality. The potential effect of a credit guarantee on consumption is ambiguous because it allows a shift between consumption and saving. Understanding the effect that guarantees have on consumption provides the framework for moving to the aggregate level where guarantees may influence trade. Other conceptual representations of the impact of credit are contained in Vercammen (1998) and Kehoe and Levine (1993).

Consumption model: Consider a consumer with demand for a single consumption good available at a price of one in two periods, as illustrated in figure 1. Let consumption be q_0 in period 0 and q_1 in period 1. Assume that income in both periods, Y_0 and Y_1, and the discount rate for future consumption are known. A consumer maximizes utility by choosing the amount to consume in each period. In the absence of credit, consumption would occur at point a, where $q_0=Y_0$ and $q_1=Y_1$. Credit allows consumption to change across the two periods. With credit, the consumer's budget constraint, B^m, has a slope of $-(1+r_m)$, where r_m is the private market rate of interest. The consumer would borrow at r_m in period 0 and consumption would occur at point b, where indifference curve U is tangent to B^m.

A direct price subsidy, offered in period 0, affects consumption in each period. Assume credit is available and the consumer begins at point b. A direct price subsidy in period 0 rotates the budget constraint around the q_1 intercept to B^p (the dashed line in figure 1). The consumer increases utility by moving to point c, where the indifference curve U' is tangent to B^p. As the consumer moves from point b to point c, consumption increases from q_0^b to q_0^c. The increase in consumption, $q_0^c - q_0^b$, would constitute additionality attributable to the direct price subsidy.

A credit guarantee may have a similar outcome on consumption, but through different means. If the guarantor has a comparative advantage managing credit risk relative to the private market, then a guarantee results in a lower interest rate, r_g, for the consumer. The lower interest rate rotates the budget constraint counterclockwise around a point (shown to the left of point a in figure 1) to B^g. The consumer increases utility by moving to where indifference curve U' is tangent to B^g, point c. At this higher utility level, consumption in period 0 increases from q_0^b to q_0^c. The consumption increase in period 0 would be attributable to the implicit guarantee subsidy and would constitute additionality.

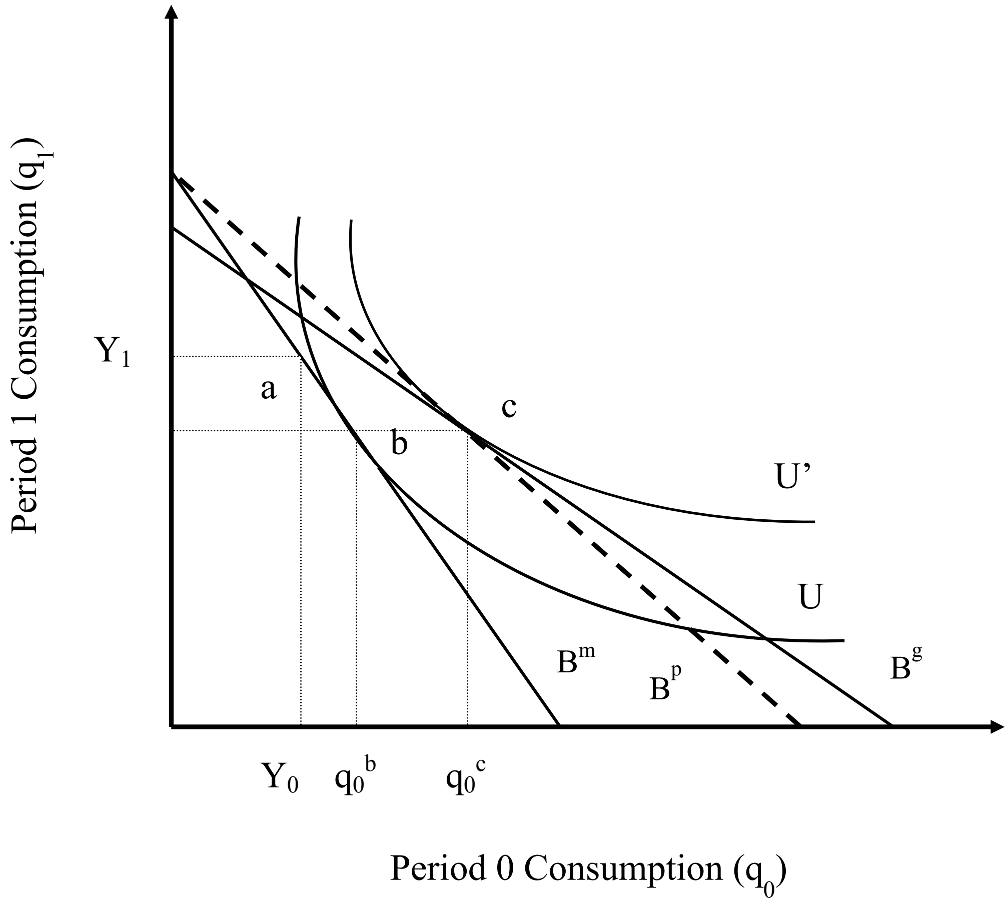

Figure 1. Additionality of a credit guarantee subsidy

The amount of additionality from the price subsidy and the guarantee are equivalent in this example. This stylized feature was intended to show how the source of additionality differs for the two instruments, while the outcome could be equivalent. It seems reasonable, therefore, to quantify the benefits the consumer receives from those instruments and relate them to changes in imports.

Additionality attributable to guarantees depends on the guarantors' comparative terms. Sources of advantage include: a lower cost of funds, more financial resources, greater risk-assessment ability, and, if the guarantor is a government, sovereign power. Regardless, guarantors assume the credit risk of borrowers and then charge fees to offset any contingent liability. The effect of any fee is to shift or rotate the budget line back toward the origin, in a way that depends on how the fees are structured (paid up front or capitalized). If a guarantor charges a high enough fee, it could reduce or eliminate additionality.

A difference between a direct price subsidy and a credit guarantee is the level of additionality that would occur under market failure, e.g., when private credit is constrained.

Without any available credit, consumption occurs at point a, where $q_0=Y_0$ and $q_1=Y_1$. If credit is constrained and a price subsidy is offered, consumption in period 0 would increase to some level less than q_0^c, where the budget constraint B^p is reached. In contrast, if credit is constrained and a guarantee is offered, period 0 consumption could again reach q_0^c.

Import model: Consider a country that is not self-sufficient producing wheat. Aggregate demand for wheat is influenced by income, prices, and benefits from trading tools used by wheat exporting countries, e.g., direct price subsidies and guarantees. The aggregate inverse demand for wheat is specified as:

$$P^d = f(Q^D, Y, V),$$

where P^d is the domestic price of wheat, Q^D is the quantity of domestic demand, Y is income, and V is the value of any direct or indirect exporter subsidy.

Traders handle wheat from exogenous domestic production, Prod, and imports, D. Assume that both sources of wheat are available to merchandisers at world price, P^w. Merchandisers combine domestic and imported wheat in a production function, *h*(Prod, D), in order to handle Q^S, the quantity of wheat supplied domestically. Their actions may include sourcing, cleaning, blending, and distributing wheat. Traders choose the level of imports to maximize their profit, specified as:

$$\pi = \{P^d Q^S - (P^w \text{ Prod} + P^w D) \mid Q^S = h(\text{Prod}, D)\}.$$

Upon optimization, the envelope theorem is used to obtain the input demand function for imported wheat, $D^* = D(P^d, P^w, \text{Prod})$. Under market equilibrium, $Q^D = Q^S$, the domestic market is satisfied, allowing the substitution of P^d resulting in $D^* = D(Y, V, P^w, \text{Prod})$. Hence, imports depend on domestic income, the subsidy from exporters, the world wheat price, and domestic production.

Empirical Specification

The model was stylized to account for features of the international wheat trade. Consumers, in a typical wheat importing country, face multiple goods and time periods and multiple exporters that extend direct price subsidies and credit guarantees. A complicated trade flow model is infeasible in the presence of guarantees, but the benefit of guarantees can still be modeled to assess additionality. Most EEP payments and guarantees were used on only a portion of the imports for most countries. Thus, the subsidy or benefit is only applicable for the volume of trade under a given tool and is measured as dollars transferred to importers. There were also instances of sales under both credit guarantees and EEP.

The value or benefits from credit guarantees should reflect the total savings a wheat importer receives from the guarantee. The guarantee subsidy, V^k_j, is defined as the product of the subsidy rate, S^k_j, and the loan volume, L^k_j, under guarantees extended by exporter k to importer j. The guarantee subsidy reflects the total implicit discounted savings, in dollars, that the importer receives from the guarantee. This allows for a direct test of the significance of guarantee and price subsidies.

The subsidy rate is derived using a formula measuring the savings importers receive from a guaranteed relative to a non-guaranteed loan on a per dollar of loan basis. Raynauld's formula [1992, equation (1) (p. 42)] for the subsidy rate S is

$$S = 100(1\frac{r}{i})[1-\frac{1-\frac{1}{(1+i)^{T}}}{iT}] \,, \tag{1}$$

where r is the interest rate on guaranteed loans, *i* is the market (discount) rate for the importing country, and T is the term of the loan being guaranteed. This yields a guarantee subsidy as a percentage of the loan volume guaranteed. The formula discounts the implicit interest differential (which exists over the life of the guaranteed loan) to the current period. Other than transaction fees, the interest rate charged on guaranteed loans is comparable to the cost of capital in less risky countries. The market rate, *i,* represents the opportunity cost of funds for the guarantee recipient. This would be the rate of interest the importer would pay for private financing with similar terms. The market rate varies by importer and reflects a premium over the rate with a guarantee.

The presence of multiple guarantors may result in "crowding-out." It is possible that competition among guarantors is so intense that they purely subsidize importers and displace sales that would have occurred anyway. Guarantor competition could reduce additionality attributable to guarantees, but guarantors typically face constraints on their activity through budgets, legislation, and internal rules such as country limits. Hence, the presence of multiple guarantors warrants inclusion of the different guarantee subsidies available to each importing country. Given the composition of our sample countries and for these reasons, we account for the prospective differential effects across export country programs.

The empirical import demand model is specified as

$$D_{jk} = f(P^{US}, P^{CA}, P^{FR}, V^{eep}{}_{j}, V^{US}{}_{j}, V^{CA}{}_{j}, V^{FR}{}_{j}, PROD_{j}, GNPPC_{j}) + e_{jk} \tag{2}$$

where D_{jk} is wheat volume imported by country j from exporter k; P^{US}, P^{CA}, and P^{FR} are FOB export prices for the United States, Canada, and France, respectively (P^{FR} includes the export restitution); $V^{eep}{}_j$ is the total EEP subsidy ($V^{eep}{}_j = B_j \cdot Q^{eep}$ where B_j and $Q^{eep}{}_j$ are bonus and volume of wheat sold under the EEP to country j), $V^{US}{}_j$, $V^{CA}{}_j$, and $V^{FR}{}_j$ are the values of the total guarantee subsidy (further described below) derived for CCC, Canadian Wheat Board (CWB), and COFACE (la compagnie Francaise D'assurance pour le commerce exterior), respectively; $PROD_j$ is wheat production in the importing country; $GNPPC_j$ is gross national product per capita for the importing countries; and e_{jk} is the error term. Domestic production comprises a considerable portion of wheat used in these countries and its treatment is similar to Koo and Karemera (1991), Arnade and Davison (1987), and Lent and Dusch (1994).

SCOPE OF ANALYSIS AND DATA SOURCES

A pooled cross-sectional time-series model of imports by a group of importing countries was estimated. Separate models were estimated for the exporting countries of the United

States, Canada, and France. Credit allocations and acceptances have been sporadic across countries and time. Six countries receiving guarantees were chosen for analysis: Algeria, Brazil, Egypt, Mexico, Morocco, and Tunisia. Taken together, these countries provide sufficient observations for the econometric analysis. Each country has at least one competing guarantor. Time series data for 20 years are used, which should give robust estimation with the pooled sample. Each model has 132 observations and 117 degrees of freedom. The last year for which observations for all variables are available is 1992.[4] Credit was used extensively since 1981. Since the mid-1990s, there have been major changes in the structure of the programs and much of the necessary data, particularly about risk premiums, are not available. Though the time period of the study is dated, it contains observations both before and after the credit guarantees and thus makes it useful to test the hypothesis about their impact on demand.[5]

Aggregate trade data were taken from several sources. Quantities imported (in 1,000 mt), production (1,000 mt), and all prices ($/mt) were from the International Wheat Council's *World Grain Statistics* and *World Wheat Statistics* (IWC 1992, 1993, 1994). These are the most reliable and for these purposes suffice, though it is recognized that transaction prices would be preferred. Canadian exports were taken from the CWB *Annual Reports*, and U.S. quantities were from the U.S. Department of Agriculture, Economic Research Service (USDA-ERS 1995, 2000). In cases where these sources were inconsistent with IWC data, the latter were used.

Exports under credit guarantees for the CCC (in $1,000) were taken from the *Notice to Exporters* [USDA, Foreign Agricultural Service (FAS)] and for the CWB (in 1,000 mt) were taken from the CWB *Annual Reports*. French shipments under COFACE guarantees (in 1,000 mt) were taken from the *Grain Market Report* (IWC 1992, 1993, 1994) and the *Secretariat Report* (IWC 1988). EEP data (bonus in $/mt and quantity in 1,000 mt) are from *Agricultural Export Assistance Update Quarterly Report* (USDA-FAS). Per capita income was obtained from *World Tables 1994* developed by the World Bank (1994).

An approximation was used to derive the market interest rate for different importers. Risk premiums described in Hyberg et al. (1995) were obtained from one of the authors (Skully 1994). Their method adjusted the premium to terms comparable to these guarantees. Importer specific market rates, *i*, were then computed as the London Interbank Offer Rate (LIBOR), plus the risk premium. LIBOR was obtained from the International Monetary Fund (1995). While approximated, the market rates fell within the sporadic observed rates for these and similar risky countries on bonds and other instruments.

The U.S. subsidy rate, S^{US}_{j}, was computed using terms on CCC guarantees. CCC guarantees typically covered 98% of the loan principal. Thus, lenders still face some exposure and charge a fee. Interest rates on CCC guaranteed loans are a small spread over LIBOR. During the study, the spread was between 25 and 100 basis points (Vanderbeek 1994). Because those spreads are not available for specific importing countries, the interest rate was set at LIBOR+25 basis points for all importers. A full term of three years was assumed, along with annual payments and discounting. For reference, risk premiums, the interest rate

[4] Specifically, data for market rates in the form we used were only available through 1993.

[5] Indeed, a natural extension of this study would be to revise the analysis to be inclusive of the more recent years. These are characterized by no EEP payments, but a radically different structure of credit arrangements making it very challenging to quantify the value of these instruments.

charged, the CCC subsidy rate, and the total guarantee subsidy are shown in table 1 for observations with U.S. guarantees.

Table 1. CCC Guarantee Subsidy Parameters and Estimation

Guarantee Recipient	Year	Risk premium	LIBOR + Premium	LIBOR + 25 Basis Points	CCC Subsidy Rate (S^{US})	CCC Subsidy (V^{US})
		%				$1,000
Algeria	86	1.02	7.97	7.20	1.36	1,529
	87	1.67	9.28	7.86	2.45	4,311
	88	3.03	11.44	8.66	4.66	10,228
	89	2.21	11.52	9.56	3.28	4,307
	90	2.72	11.17	8.70	4.15	6,557
	91	3.25	9.54	6.54	5.17	8,072
	92	3.78	7.93	4.40	6.23	9,000
Brazil	82	0.71	14.40	13.94	0.74	2,483
	83	1.75	12.55	11.05	2.47	8,861
	84	2.68	14.50	12.07	3.90	17,283
	85	3.62	12.73	9.36	5.54	2,898
	86	5.44	12.39	7.20	8.58	2,984
	87	5.10	12.71	7.86	7.98	456
	88	4.80	13.21	8.66	7.43	1,255
	89	4.43	13.74	9.56	6.78	381
Egypt	82	3.06	16.75	13.94	4.37	7,208
	83	2.08	12.88	11.05	3.00	1,895
	84	0.33	12.15	12.07	0.13	44
	85	3.10	12.21	9.36	4.72	6,158
	86	6.10	13.05	7.20	9.58	6,451
	87	6.15	13.76	7.86	9.56	17,395
	88	5.46	13.87	8.66	8.43	24,801
	89	5.00	14.31	9.56	7.64	19,950
Mexico	86	5.96	12.91	7.20	9.37	125
	87	5.75	13.36	7.86	8.96	8,109
	88	4.67	13.08	8.66	7.23	5,525
	89	4.14	13.45	9.56	6.33	2,950
	90	3.13	11.58	8.70	4.81	1,591
	91	2.44	8.73	6.54	3.82	1,776
	92	1.80	5.95	4.40	2.82	2,782
Morocco	80	0.50	13.94	13.69	0.40	207
	81	0.93	17.06	16.38	1.05	804
	82	2.63	16.32	13.94	3.72	3,995
	83	2.43	13.23	11.05	3.56	6,151
	84	3.50	15.32	12.07	5.15	3,498
	85	4.65	13.76	9.36	7.13	7,093
	86	7.01	13.96	7.20	10.92	10,306
	87	6.28	13.89	7.86	9.75	9,539
	88	5.33	13.74	8.66	8.23	9,273
	91	4.00	10.29	6.54	6.39	459
	92	3.69	7.84	4.40	6.08	9,656
Tunisia	82	1.34	15.03	13.94	1.74	370
	83	1.07	11.87	11.05	1.37	944
	84	0.46	12.28	12.07	0.35	112
	85	1.82	10.93	9.36	2.65	171
	86	4.35	11.30	7.20	6.88	2,619
	87	3.26	10.87	7.86	5.08	1,859
	90	2.38	10.83	8.70	3.60	745
	91	2.76	9.05	6.54	4.35	401
	92	3.00	7.15	4.40	4.91	807

The interest rate and other terms of CWB guarantees were similar to CCC guarantees (Harris 1990) during the study period. Hence, for the same importer in the same year the subsidy rate would be the same for CCC and CWB, but the total guarantee subsidy could vary depending on the loan amount. COFACE guaranteed loans typically have longer terms. The subsidy rate, S^{FR}, was computed with a term of seven years and would be higher than CCC and CWB guarantees in the same year to the same importer.

The loan amount guaranteed was reported differently across exporting countries. CCC loan amounts, in dollars, under guarantees are published and the values of L^{US}_j were available for each importing country. CWB and COFACE report the ex-post amount of wheat sold (in mt) under guarantees. Loan volumes for both of these guarantors, L^{CA} and L^{FR}, were approximated as the product of wheat quantity under the guarantee (in mt) and the price for the period (in \$/mt).

STATISTICAL RESULTS AND HYPOTHESIS TESTS

The sample of time series observations was pooled across importers. For each exporting country, the linear fixed effects were significant. The models that were most robust are presented. Results are presented for each exporting country (table 2). Empirical estimates are discussed along with measures of additionality, where applicable.

United States

Model Choice and Statistical Results: An F-test for the U.S. model suggests rejection of the null hypothesis that all dummy variables (importing countries) are zero (the F-statistic was 22.54). Thus, these variables were retained. Each of the price variables had the *a-priori* sign. GNPPC is not significant, but PROD is highly significant, reflecting its importance as a determinant of import demand.

Effects of both V^{eep} and V^{US} are significant and positive on U.S. exports. V^{CA} is significant, indicating that CWB credit subsidies adversely affect demand for U.S. wheat. V^{FR} is insignificant, suggesting that the effect of COFACE subsidies on U.S. exports may not be important. Insignificance of P^{US} and P^{FR} may be attributable to the strong influence of the EEP.

Pooling across countries and over time assumes a constant variance among groups and over time. The residuals were tested for heteroscedasticity, correlated errors among countries, and autocorrelation [see Diersen et al. (1997) for a description of these tests and statistical results]. While the testing indicated the presence of these effects, a different estimation procedure (two-step least squares) failed to correct for the possible inappropriate assumptions. The parameter estimates were generally robust, with the V^{eep}_j coefficient being smaller and the V^{CA}_j coefficient being larger using two-step estimation.

Table 2. Parameter Estimates: United States, Canada, and French Models

Independent variable	Parameter estimates		
	United States	Canada	France
GNPPC	-0.012	0.137**	0.088*
	(0.091)[a]	(0.033)	(0.050)
PROD	-0.233**	-0.089**	-0.082**
	(0.057)	(0.021)	(0.032)
P^{US}	-2.449	0.257	-1.995
	(2.996)	(1.103)	(1.656)
V^{eep}	0.015**	-0.001-	0.003
	(0.002)	(0.001)	(0.001)
P^{CA}	5.918**	-0.177	5.218*
	(2.860)	(1.053)	(1.581)
P^{FR}	0.086	0.751	-2.128**
	(1.681)	(0.619)	(0.929)
V^{US}	0.057**	-0.004	0.006
	(0.014)	(0.005)	(0.008)
V^{CA}	-0.072**	0.046**	0.029
	(0.032)	(0.012)	(0.018)
V^{FR}	0.036	-0.004	-0.002
	(0.044)	(0.016)	(0.024)
Intercept[b]	214.679	200.732*	281.356*
	(302.940)	(111.586)	(167.468)
Brazil	1,210.717**	506.858**	-478.768**
	(205.963)	(75.865)	(113.858)
Egypt	1,345.195**	-29.865	885.116**
	(228.875)	(84.304)	(126.524)
Mexico	298.589	-182.302*	-488.349
	(212.873)	(78.410)	(117.678)
Morocco	209.424	-137.992**	110.509
	(210.909)	(77.687)	(116.592)
Tunisia	-424.408**	-359.468*	-325.033**
	(177.309)	(65.310)	(98.018)
Adj. R^2	.69	.76	.75
Dependent Variable	D_{US}	D_{CA}	D_{FR}

Note: [a] The standard errors are in parentheses
[b] The intercept term corresponds to Algeria
* Indicates significance at the 0.10 level
** Indicates significance at the 0.05 level

Additionality Estimates of U.S. Programs: These results indicate additional sales occurred in these importing countries due to credit guarantees. The estimated coefficients for the subsidy variables measure importers' responses to changes in subsidies. V^j are measured as the combined effects of changes in loan volumes guaranteed and the interest subsidy rate. Additionality is interpreted as the product of the subsidy coefficient and the average subsidy for the sample. This provides a direct measure of additionality, as opposed to the measure reported in the U.S. GAO (1995), and isolates the subsidy effect from that of other programs and prices.

The guarantee subsidy accounts for a significant portion of variability in U.S. exports and is an indicator of additionality. In addition, though not shown here, it remains significant and similar in magnitude regardless of the error term assumption. Results indicate that a $1,000 change in the subsidy value (a subsidy unit) resulted in a 57 mt (0.057*1,000 mt) change in imports (1,000 mt is the unit of quantity imported), on average, during the sample period. The effect of the subsidy was quantified over time for the 50 observed guaranteed loans to the six importing countries. The average CCC subsidy was $5.1 million. Loan guarantees averaged $105 million with an average subsidy rate of just over 5% of loan volume. Thus, on average, the subsidy accounted for 292,000 mt of additionality. This is about 23% of the average (1,261,000 mt) of total wheat exports to the sample of importing countries and 33% of the average (877,000 mt) guaranteed quantity. The estimated parameter for the credit subsidy can be used to measure additionality for each importing country on an annual basis (table 3). In 1986, Algeria had 87,000 mt of additional imports attributable to the credit guarantee. Additionality was greatest for Egypt, with 4.8 mmt of additionality over eight years. The lowest total was for Tunisia, which was the smallest importer in this study.

The EEP parameter was significant and similar in value to that of the CCC subsidy. Results indicate that a $1,000 change in the EEP payment (bonus times quantity) results in an estimated 15 mt change in imports. The average bonus value for these countries was $32 per mt on EEP volumes of 905,000 mt. Thus, the average impact of EEP subsidies was 492,000 mt, roughly 54% of the EEP sales.[6] Estimates of additionality were derived for each year (table 4). The credit guarantee generated an additional 14.6 mmt to this sample of countries from 1980 through 1992. The year with the greatest additionality was 1988. In comparison, additionality due to EEP in these countries over 1985 to 1992 was 19 mmt.

Hypotheses Tests: Results show that both credit guarantees and the EEP added to wheat exports and that Canada's guarantee program adversely affected U.S. exports. To evaluate the relative effectiveness of the programs, several hypotheses were formulated and statistical tests conducted.

Table 3. Estimates of Additionality: CCC Guarantee Program by Year and Importer

	Additionality by Importing Country (1,000 mt)					
Year	Algeria	Brazil	Egypt	Mexico	Morocco	Tunisia
80					12	
81					46	
82		142	411		228	21
83		505	108		351	54
84		985	2		199	6
85		165	351		404	10
86	87	170	368	7	587	149
87	246	26	992	462	544	106
88	583	72	1,414	315	529	
89	246	22	1,137	168		
90	374			91		42
91	460			101	26	23
92	513			159	550	46
Total	2,509	2,087	4,783	1,303	3,476	457

[6] The V^{eep} coefficient times the average total subsidy (0.015 * 32,188).

Table 4. Estimates of Additionality: United States Export Programs by Year

	Program Additionality (in 1,000 mt)		
Year	EEP	GSM-102, 103, 105	Total
80		12	12
81		46	46
82		802	802
83		1,018	1,018
84		1,192	1,192
85	1,093	930	2,023
86	2,528	1,368	3,896
87	4,244	2,376	6,620
88	1,556	2,913	4,469
89	531	1,573	2,104
90	2,066	507	2,573
91	3,915	610	4,525
92	3,111	1,268	4,379
Total	19,044	14,615	33,659

One function of the CCC guarantee program is to compete with other guarantors' programs. A test was conducted to evaluate the relative effects of these programs on U.S. exports. A t-test failed to reject the null hypothesis that the guarantee subsidy coefficients have the same magnitude (the t-statistic was -0.47 with 120 d.f.). Thus, CWB subsidies have an equal and opposite impact of CCC subsidies on U.S. wheat exports. Interpretation of this result is that a dollar of CWB credit subsidy has an equal and opposite effect of a dollar of CCC subsidy, implying that their effects are offsetting in terms of changing U.S. exports.

The effect of credit guarantees versus direct subsidies is an important policy question. Previous empirical studies (e.g., Skully 1992; Haley 1989) assumed credit guarantees to be equivalent to direct price subsidies. However, the theoretical model suggests this is not necessarily true. The validity of this assumption was tested with these results. The variables V^{eep}_j and V^k_j were defined on a comparable basis, making the coefficients directly comparable. The null hypothesis is that they are equal, i.e., a dollar of guarantee subsidy is equivalent to a dollar of price subsidy. The t-test rejected the null hypothesis (the t-statistic was -2.87 which was greater than the table t-value).

The effect of the guarantee subsidy is not equivalent to the effect of a direct price subsidy on U.S. exports. The guarantee subsidy accrues as an interest savings across the loan volume guaranteed, is indirect, and affects consumption over time. In contrast, EEP subsidies are direct payments, which are transferred to importers via a lower selling price. These results indicate that importers do not respond to these subsidies in a similar manner. A dollar of credit guarantee subsidy has a greater impact in terms of additional exports than does a dollar in EEP subsidy.

There are two important distinctions in interpreting these conclusions. First, besides providing an implicit price subsidy, credit guarantees could also relax credit constraints, making it more valuable than a direct price subsidy. Second, the test is of the equivalence of estimated parameters for two variables which are interpreted as the dollar value of the subsidy. The effect of the EEP to an individual country depends on the bonus and quantity, and the effect of a guarantee depends on the subsidy rate and loan volume. The two correspond to the importers' trade-off between substitution of wheat and other commodities for the amount of bonus and substitution of current consumption to future consumption for the amount of a credit subsidy, respectively.

Some importing countries were targeted with both programs' subsidies. In the sample, V^{eep} and V^{US} are correlated with r = .45 suggesting possible interaction among the subsidies. To account for this, an interaction term among the subsidies was added to the basic model. The results indicated this effect was not significant. The F-statistic in comparing the restricted and unrestricted models was 1.98.

Canada

The Canadian model explained 76% of the variation and depends heavily upon the CWB credit subsidy. PROD was significant with a negative sign. GNPPC was significant and positive, indicating that as incomes rise, more Canadian wheat is demanded. V^{CA} was the only significant guarantee subsidy variable. Prices have *a-priori* expected signs, but are not significant. Testing for the inclusion of importers' specific dummy variables in the Canadian model yielded an F-statistic of 33.191, large enough to reject the null hypothesis that all dummy variables are zero.

An interesting aspect of these results is the number of insignificant parameter estimates. V^{eep} and V^{US} are not significant, suggesting that U.S. subsidies have not adversely affected exports from Canada in these markets. Additionality was measured using the V^{CA} parameter estimate. The higher V^{CA} parameter estimates and lower loan volume, on average, relative to the United States, yields about the same level of additionality for Canada and the United States (table 5). Algeria and Brazil accounted for most of Canada's additionality of 4.4 mmt.

Table 5. Estimates of Canadian Additionality by Year and Importer

	Additionality by Importing Country (1,000 mt)			
Year	Algeria	Brazil	Egypt	Mexico
82		93		4
83		260	161	72
84		363	5	
85		455	183	
86	12	422	126	90
87	125	237		92
88	146			
89	171	117		
90	223	125		
91	147	340		
92	273	151		
Total	1,097	2,563	475	258

France

The dummy intercept variables were significant in the French model. The F-statistic for inclusion of these effects is 35.91, large enough to reject the null hypothesis that all dummy variables are zero. Results for France differ from those of the United States and Canada. V^{eep}, P^{FR}, and P^{CA} are all significant with *a-priori* signs as were income and domestic production. Since V^{FR} was not significant, estimates of additionality were not derived. These results suggest that price is more important than credit which differs from the results for the other exporting countries. The V^{eep} is marginally significant, and the magnitude of the coefficient is smaller than in the U.S. model; thus, EEP increases U.S. exports more than it harms French exports. The competing credit subsidy variables were not significant.

SUMMARY AND IMPLICATIONS

One of the important problems confronting competition among exporters' credit programs is estimating the volume of trade that can be attributed to guarantees. Any subsidy element associated with a guarantee program is implicit, and the benefits accrue over time, as opposed to directly affecting price. Most major competitors use similar programs, potentially dissipating the effects of a single country's credit programs. These indirect subsidy programs ultimately have to compete with the direct price subsidies. The purpose of this study was to analyze the additionality attributable to export credit guarantees relative to other programs.

Results indicated that the GSM programs have resulted in additional exports that would not have occurred without the programs. Additionality of CCC guarantees totaled approximately 14.6 mmt to the sample importing countries over 13 years. Additionality attributed to Canadian credit guarantees was 4.4 mmt to the sample countries.

For the WTO and OECD, these results provide reasons why individual exporters are reluctant to revise or reduce the scope of their credit programs. The results show that providing credit impacts demand and has strategic implications. For policymakers in individual countries, an important issue has been the effectiveness of credit guarantee programs relative to direct price subsides.[7] These results indicate the credit subsidy (from guarantees) provided more additionality than the EEP when the subsidy is measured on an equivalent dollar value. Thus, though both of these programs had the effect of increasing exports, on a per dollar basis of subsidy equivalence, credit guarantees had a greater impact on exports. The effectiveness of these programs relative to competing country programs is an important strategic consideration. The results suggest Canada's guarantee program may do more to displace U.S. sales than it does to help Canadian sales. These results indicate that the effect of COFACE subsidies on competing countries was not substantial.

Guarantee programs were criticized for their high cost (U.S. GAO, 1992) and have recently been identified as a major agenda item in the next WTO round, following efforts by the OECD to resolve disputes. For the United States, it is important to note that the EEP subsidy was not as favorable as credit guarantees when compared to expected program costs.

[7] In previous studies, credit guarantees were viewed as providing a default subsidy (Haley 1989) and a pure price subsidy (Skully 1992). Vercammen (1998) illustrated that export credit guarantees facilitated price discrimination.

Guarantee programs are important to Canada, both strategically and to induce sales in competitive situations where credits play an important role.

Finally, policy decisions on credit offerings by export countries are becoming increasingly interdependent. This was not true during the time period of this study when credit allocation decisions were made by agencies independent of selling firms and organizations and made prior to the commencement of a marketing year. In recent years these practices have changed. The inability of these competitors to agree on credit guidelines suggests their interdependency has escalated. Consequently, an important area of future research would be to explore these policy decisions (in this case, credit allocation decisions) in some form of game theoretic model.

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WELFARE EFFECTS OF SUPPLY EXPANSION WITH TRADE RESTRICTIONS: THE CASE OF SALMON*

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ABSTRACT

In response to protectionist pressures from United Kingdom salmon producers, in 1996 Norway instituted a feed quota to limit its production. The feed quota made the Norwegian and world supply curves for salmon less price elastic. This contributed to price instability, magnified the downward pressure on world salmon prices associated with the 71% increase in supplies of farmed salmon from Chile between 2000 and 2002, and increased Norway's incidence of US salmon tariffs. Our analysis suggests these unintended consequences of the feed quota policy are nontrivial, amounting to some $62 million in lost surplus to UK producers between 2000 and 2002, equivalent to 9.5% of export value.

Key words: *quotas, supply control, tariffs, unintended consequences*.
JEL classification Codes: Q13, M30, M37

INTRODUCTION

Traditionally, Norway has been the world's largest producer of salmon, accounting for some 46% of total world production in 1995 (Asche 1997). However, Norway's dominant position is being challenged by emerging low-cost producers, most notably Chile. In 2002 Chile accounted for 25% of the world's farmed salmon production, up from 20% in 1995.

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Perhaps most telling is that between 2000 and 2002 the ratio of Chile's price to Norway's price declined from 0.99 to 0.74 (table 1).

Table 1. Baseline Data for the World Salmon Model

Variable	Definition	Value			
		2000	2001	2002	Average
P_N	Norway price	3.78	2.77	3.07	3.21
P_{CH}	Chile price	3.76	2.68	2.26	2.90
P_{UK}	UK price	4.06	2.96	3.25	3.42
P_{CAN}	Canada price	5.61	5.24	4.81	5.22
P_{FI}	Faroe Islands' price	3.73	2.28	2.55	2.85
P_{RS}	ROW price (producing countries)	2.93	3.18	3.43	3.18
P_{EU}	EU price	3.93	2.77	3.12	3.27
P_{US}	US price	5.22	4.67	4.36	4.75
P_J	Japan price	6.23	4.76	5.33	5.44
P_{RD}	ROW price (consuming countries)	4.98	3.40	3.37	3.92
X_N	Norway's net exports	409	399	438	415
X_{CH}	Chile's net exports	138	203	236	192
X_{UK}	UK's net exports	63	67	60	63
X_{CAN}	Canada's net exports	46	56	65	56
X_{FI}	Faroe Islands' net exports	31	44	42	39
X_R	ROW's net exports	13	19	18	17
ΣX_i	*Total net exports*	*700*	*788*	*859*	*782*
M_{EU}	EU's net imports	392	409	410	404
M_{US}	US' net imports	179	224	259	221
M_J	Japan's net imports	39	43	38	40
M_R	ROW's net imports	90	112	152	118
ΣM_j	*Total net imports*	*700*	*788*	*859*	*782*

Source: Norwegian Seafood Export Council and government statistics.

Note: Prices are fob expressed in constant US 2001 dollars per kilogram; quantities are raw weight expressed in 1000 metric tons.

Chile's growing price advantage presents a challenge to the wild-caught salmon sector, which is located chiefly in North America. But it also poses a challenge to European producers of farmed salmon, especially producers in the United Kingdom where production costs may be higher than in Norway.[8] It should not be surprising, therefore, that producers in the UK and US have sought protection through tariffs and other measures designed to reduce the supply of Norwegian and Chilean salmon on world markets. In particular, in response to pressures from UK producers, in 1996 Norway instituted a feed quota in an effort to reduce supply and strengthen market prices. About the same time the United States imposed countervailing and dumping duties on Norwegian salmon averaging 26.1%, which were followed by similar duties on Chilean salmon averaging 5.5%. More recently, the European Commission has imposed minimum import prices and a tariff-rate quota on farmed salmon imports and a provisional anti-dumping duty of between 6.8% and 24.5% on imports from Norway in an effort to protect Scottish and Irish producers (Bendz 2005; Lem 2005).

Ordinarily supply control is used to protect *domestic* producers by taking advantage of an inelastic demand (Hertel and Tsigas 1991; Kola 1993; Edelman, Langley and White 2003). Norway's feed quota is unique in that it is designed to protect *foreign* (namely UK) producers by limiting the supply of Norwegian salmon in the European Union market. Because most studies suggest the export demand for salmon is price elastic (Bjørndal, Salvanes, and Andreassen 1992; Hermann, Mittlehammer and Lin 1993; Asche, Bjørndal and Salvanes 1998), and Norway exports over 90% of its production, the supply restriction is not likely to benefit Norway's producers from a revenue perspective. At issue here, however, are the more subtle impacts stemming from policy interactions and induced supply growth.

The purpose of this research is to determine the unintended welfare losses from Norway's feed quota policy. As noted by Houck (1986, p. 181) "most...trade policy schemes accentuate the price inelasticity of world supply and demand functions." In the case of the feed quota, we show that the inelasticity introduced into the world supply curve for salmon aggravated price instability, magnified period-to-period swings in trade volumes, undermined US tariff policy, and intensified producer losses associated with increased supplies of farmed salmon from Chile. The feed quota also diluted the spillover benefits that UK and other international competitors enjoy as a result of the US tariffs targeted against Norway and Chile.

The analysis proceeds by first specifying a partial equilibrium model of the world salmon market similar to the one used by Kinnucan and Myrland (2000, 2002) to assess the Norway-EU salmon agreement. The model is then simulated to determine the price, trade, and welfare effects of Chile's supply expansion and US tariffs with and without the feed quota. A concluding section summarizes the key findings.

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[8]Although Scottish producers claim their production costs are higher than in Norway due to limits on fish farm sites and extra tax burdens, no data exist to support the claim. A study has been commissioned to determine the comparative costs of salmon farming in different countries (Callender McDowell 2005), which should help to clarify the issue.

MODEL

Specification

The model is based on two simplifying assumptions: the world salmon market is sufficiently competitive that the Law of One Price (LOP) holds, and different salmon species (e.g., Pacific and Atlantic) are sufficiently close substitutes that the generalized Composite Commodity Theorem holds. Empirical support for LOP for *long-run* analysis (the major focus of this paper) is provided in the studies by DeVoretz and Salvanes (1993), Asche, Bremmes, and Wessels (1999), Steen and Salvanes (1999), and Asche (2001). As for whether it is valid to treat different salmon species as a single good as implied by CCT, Asche, Bremmes, and Wessels (1999, p. 579) state "when investigating long-run market issues ... studies that aggregate salmon from species-specific to a generic product category are capturing much of the information about the international market."

With the foregoing assumptions, the supply side of the model consists of six equations as follows:

$$X_N{}^* = \acute{\varepsilon}_N P_N{}^* \qquad \text{(with feed quota)} \tag{1a}$$

$$X_N{}^* = \varepsilon_N P_N{}^* \qquad \text{(without feed quota)} \tag{1b}$$

$$X_{CH}{}^* = \varepsilon_{CH} \ (P_{CH}{}^* - \gamma_{CH}) \tag{2}$$

$$X_i{}^* = \varepsilon_i P_i{}^* \qquad (3) - (6)$$

where equations (1) and (2) refer to export supplies from Norway and Chile, respectively, and equations (3) - (6) refer to export supplies from UK, Canada, Faroe Islands, and Rest-of-World (ROW). (For variable definitions and numerical values, see table 1.) The asterisks (*) in these equations denote percentage changes (e.g., $X_i{}^* = dX_i/X$), the ε_i refer to excess supply elasticities, and the γ_{CH} parameter in equation (2) refers to the vertical shift in Chile's excess supply curve over the study period (to be discussed later). The excess supply curve for Norway is specified in two equations to accommodate the supply elasticity reducing aspect of the feed quota wherein $\acute{\varepsilon}_N < \varepsilon_N$. The disaggregation of supply into six regions permits a detailed assessment of the unintended consequences of the feed quota.

The demand side of the market is separated into four regions as follows:

$$M_j{}^* = -\ \eta_j P_j{}^* \qquad (7) - (10)$$

where the j subscript refers to the EU, US, Japan, and ROW. The η_j parameters represent excess demand elasticities, which are expressed in absolute value. The EU dominates among the customer markets for salmon with a trade share of 52%, followed by the US, ROW and Japan with shares of 28%, 15%, and 5% respectively as shown in table 2.

In the early 1990s the US imposed tariffs on imports from Norway averaging 26.1%, which were later followed by tariffs on imports from Chile averaging 5.5%. These tariffs are modeled using the following identity implied by LOP:

$$P_{US} = (P_k + C_k)(1 + \tau_k)$$

where $k = N$, CH denotes Norway and Chile, P_{US} is the US price of salmon inclusive of transportation costs and the tariff, P_k is price in the named country exclusive of transportation costs and the tariff, C_k is the per-unit transportation cost, and τ_k is the *ad valorem* tariff.

Table 2. Mean Values for Quantity Share Parameters, 2000-2002

Item	Definition	Value [a]
kx_N	Norway's share of world exports (= $X_N / \sum X_i$)	0.531
kx_{CH}	Chile's share of world exports (= $X_{CHL} / \sum X_i$)	0.246
kx_{UK}	Scotland's share of world exports (= $X_{UK} / \sum X_i$)	0.081
kx_{CAN}	Canada's share of world exports (= $X_{CAN} / \sum X_i$)	0.071
kx_{FI}	Faroe Islands' share of world exports (= $X_{FI} / \sum X_i$)	0.050
kx_R	ROW's share of world exports (= $X_R / \sum X_i$)	0.021
km_{EU}	European Union's share of world imports (= $M_{EU} / \sum M_j$)	0.516
km_{US}	United State's share of world imports (= $M_{US} / \sum M_j$)	0.282
km_J	Japan's share of world imports (= $M_J / \sum M_j$)	0.051
km_R	ROW's share or world imports (= $M_R / \sum M_j$)	0.151
x_N	Share of Norway's production exported (= X_N/S_N)	0.97
x_{CH}	Share of Chile's production exported (= X_{CH}/S_{CH})	0.95
x_{UK}	Share of Scotland's production exported (= X_{UK}/S_{UK})	0.53
x_{CAN}	Share of Canada's production exported (= X_{CAN}/S_{CAN})	0.90
x_{FI}	Share of Faroe Islands' production exported (= X_{FI}/S_{FI})	0.99
x_R	Share of ROW's production exported (= X_R/S_R)	0.60
m_{EU}	Share of EU's consumption imported (= M_{EU}/D_{EU})	0.96
m_{US}	Share of US's consumption imported (= M_{US}/D_{US})	0.93
m_J	Share of Japan's consumption imported (= M_J/D_J)	0.59
m_R	Share of ROW's consumption imported (= M_R/D_R)	0.90

Source: Norwegian Seafood Export Council.

Taking the logarithmic differential of this equation holding C_k constant yields:

$$d \ln P_{US} = (P_k/P_{US})(1 + \tau_k)\, d \ln P_k + (\tau_k/(1 + \tau_k))\, d \ln \tau_k,$$

which may be written more compactly as:

$$P_{US}^* = (1 + \tau_k)\, \varphi_k\, P_k^* + \zeta_k\, \tau_k^*$$

where $\varphi_k = P_k/P_{US}$ are price transmission elasticities and $\zeta_k = \tau_k/(1 + \tau_k)$ are tariff transmission elasticities. If tariffs in the initial equilibrium are zero, $\tau_k = 0$ and the above equation simplifies to:

$$P_{US}^* = \varphi_k\, P_k^* + \tau_k' \qquad (11) - (12)$$

where τ_k' are the tariff rates in post-tariff equilibrium. Setting $\tau_N' = 0.261$ and $\tau_{CH}' = 0.055$, equations (11) and (12) imply the US price will rise by 26% in response to the Norwegian tariff and 5.5% in response to the Chilean tariff *provided the tariffs have no effect on prices in the targeted countries.* In reality, the tariffs will cause prices in the targeted countries to decline as import demand from the US decreases. Hence, the US price is expected to rise by less than the amount of the tariffs. A critical issue in this research is the extent to which the feed quota reduces the ability of US tariffs to raise the US price.[9]

Equations (11) and (12) link the US price to prices in Norway and Chile. Price linkages to the remaining countries are given by:

$$P_{US}^* = \varphi_l\, P_l^* \qquad (13) - (19)$$

where l indexes countries other than Norway, Chile and US and $\varphi_l = P_l/P_{US}$ are price transmission elasticities computed in an analogous fashion as for equations (11) and (12). (See table 3 for numerical values.) The model is closed by the following market-clearing condition:

$$\textstyle\sum_i^6 kx_i\, X_i^* = \sum_j^4 km_j\, M_i^* \qquad (20)$$

where kx_i and km_j are export quantity and import quantity shares, respectively, as given in table 2.[10] Equilibrium, therefore, requires that the changes in exports weighted by their respect export shares equal changes in imports weighted by their respective import shares.

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[9]The price wedge associated with a tariff equals the tariff only if the price transmission elasticity is unity. This may be seen by setting $P_{US}^* = 0$ in equations (11) and (12) and solving for foreign price to yield $P_k^* = -\tau_k'/\varphi_k$. In this equation P_k^* is the maximal decline in foreign price, i.e., the decline when the tariff's entire incidence is shifted to foreign producers. As can be seen, $P_k^* = -\tau_k'$ only if $\varphi_k = 1$. If $\varphi_k < 1$, as is true in this study, it is possible for the price declines in Norway and Chile to exceed their respective tariffs, a fact to bear in mind when interpreting results.

[10]The export and import data used in this analysis are *net* in that they take into account exporting countries' imports and importing countries' exports. For brevity, in the narrative we omit the "net" qualifier.

Table 3. Parameter Definitions and Values

Parameter	Definition	Value
$\acute{\varepsilon}_N$	Norway's excess supply elasticity with feed quota	0.44
ε_N	Norway's excess supply elasticity without feed quota	1.60
ε_{CH}	Chile's excess supply elasticity	1.64
ε_{UK}	Scotland's excess supply elasticity	3.89
ε_{CAN}	Canada's excess supply elasticity	1.80
ε_{FI}	Faroe Island's excess supply elasticity	1.53
ε_R	Rest-of-world's excess supply elasticity	3.30
γ_{CH}	Vertical shift in Chile's excess supply curve between 2000 and 2002[a]	-0.58
η_{EU}	European Union's excess demand elasticity	1.31
η_{US}	United States' excess demand elasticity	1.40
η_J	Japan's excess demand elasticity	3.08
η_R	Rest-of-world's excess demand elasticity	1.50
φ_N	US-Norway price transmission elasticity	0.67
φ_{CH}	US-Chile price transmission elasticity	0.61
φ_{UK}	US-UK price transmission elasticity	0.72
φ_{CAN}	US-Canada price transmission elasticity	1.10
φ_{FI}	US-Faroe Islands' price transmission elasticity	0.60
φ_{RS}	US-ROW exporters' price transmission elasticity	0.67
φ_{EU}	US-EU price transmission elasticity	0.69
φ_J	US-Japan price transmission elasticity	1.14
φ_{RD}	US-ROW importers' price transmission elasticity	0.82
τ_N'	Norway's average tariff rate	0.261
τ_{CH}'	Chile's average tariff rate	0.055

[a] Computed using text equation (22); see narrative for details.

Altogether the model contains 20 endogenous variables: ten to represent region-specific changes in prices, six to represent region-specific changes in exports, and four to represent region-specific changes in imports. There are three exogenous variables, two to indicate price wedges associated with US tariffs (the τ_k' parameters in equations (11) and (12)), and one to indicate Chile's supply expansion (the γ_{CH} parameter in equation (2)).

Parameterization

Numerical values for the excess supply and demand elasticities in the model were developed using the following formulas:

$$\varepsilon_i = (e_S + (1 - x_i)\, e_D)/x_i \tag{21a}$$

$$\eta_j = ((1 - m_j)\, e_S + e_D)/m_j \tag{21b}$$

where x_i are domestic export shares for the six exporting regions and m_j are domestic import shares for the four importing regions as defined in table 1.[11] The e_S and e_D parameters in equation (21) are domestic supply and demand elasticities for salmon within the specific exporting and importing regions. Given the dearth of region-specific estimates for these parameters, they are assumed to be uniform. Specifically, we set e_S to 1.5 in all regions except Norway, where e_S is set to 0.39 to reflect the elasticity-reducing aspect of the feed quota quantified by Kinnucan and Myrland (2002, p. 219). The value of 1.5 for e_S is based on Steen, Asche, and Salvanes's (1997) estimate of Norway's long-run supply elasticity prior to the feed quota, the only known empirical estimate of domestic supply response for salmon. Similarly, we set e_D to 1.2 in all regions, a value suggested by Bjørndal, Asche, and Steen's (1996) literature review.[12]

The foregoing procedure resulted in excess supply elasticities ranging from 0.44 for Norway to 3.89 for the UK and excess demand elasticities ranging from 1.31 for the EU to 3.08 for Japan (see table 3 for complete listing). The trade share weighted average excess demand elasticity is 1.45, which suggests increases in world exports can be absorbed without reductions in export value, an hypothesis to be examined later.

According to data given in table 1 Chile's exports between 2000 and 2002 increased by 71% and Chile's price declined by 40%. These price and quantity changes were converted to the implied shift in Chile's excess supply curve using the following approximation formula (see Appendix for derivation):

$$\gamma_{CH} \approx (\varepsilon_{CH}\, P_{CH}^* - X_{CH}^*)/(\varepsilon_{CH} + X_{CH}^*) \tag{22}$$

where γ_{CH} indicates the *vertical* shift, i.e., the shift in the price direction holding quantity constant. Substituting $X_{CH}^* = 0.71$, $P_{CH}^* = -0.40$ and and $\varepsilon_{CH} = 1.64$ into expression (22) yields $\gamma_{CH} = -0.581$. Hence, our analysis is based on the assumption that Chile's excess supply

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[11]Houck (1986, pp. 33-34) derives the same equations in a somewhat different form.

[12]Given the dynamic nature of salmon markets in recent years, a reviewer questioned whether the domestic demand elasticity estimate of 1.2 is still valid. For example, a recent study by Fousekis and Revell (2004) suggests the demand for salmon in Great Britain is price inelastic, a result consistent with supply increasing along a static linear demand curve. Whether this result can be generalized to other customer markets is an issue for further research. Sensitivity analysis with e_D set to 0.5 indicated little effect on results reported later.

curve between 2000 and 2002 shifted down by 58.1%.[13] Since the shift is measured in the price direction, the 58.1% may be interpreted as the estimated reduction in Chile's per-unit production costs over the period in question. By way of comparison, Anderson (2002, p. 145) reports cost reductions of some 45% in the early years of salmon aquaculture development in Norway. At issue is the extent to which the associated supply expansion in Chile harmed competing exporters, and the extent to which the feed quota magnified the harm, especially with respect to UK producers.

A caveat in using expression (22) is that it is an approximation formula. To minimize approximation error Piggott (1992, p. 133) suggests limiting displacements to about 10%. However, in Marsh's (2003) analysis where a similar formula is used a 66% displacement was modeled with no apparent ill effects. Still, the potential for approximation error must be borne in mind when interpreting results.

Price and Trade Effects

Houck (1986, p. 181) asserts that the inelasticity introduced into world supply or demand curves due to trade restrictions exacerbates price instability, but also magnifies fluctuations in trade volumes. To test this, and to provide a basis for welfare measurement, the price and trade effects of the US tariffs and Chile's supply expansion with and without Norway's feed quota are computed by expressing the model in matrix notation as follows:

$$\boldsymbol{A}\,\boldsymbol{y} = \boldsymbol{B}\,\boldsymbol{x}, \tag{23}$$

where $\boldsymbol{A}$ is a 20 x 20 matrix of parameters corresponding to the model's endogenous variables; $\boldsymbol{y}$ is a 20 x 1 vector containing the model's endogenous variables; $\boldsymbol{B}$ is a 20 x 3 matrix of parameters corresponding to the model's exogenous variables; and $\boldsymbol{x}$ is a 3 x 1 vector containing the model's exogenous variables. Inverting $\boldsymbol{A}$ and pre-multiplying both sides of equation (23) by $\boldsymbol{A}^{-1}$ yields:

$$\boldsymbol{y} = \boldsymbol{E}\,\boldsymbol{x} \tag{24}$$

where $\boldsymbol{E} = \boldsymbol{A}^{-1}\boldsymbol{B}$ is a 20 x 3 matrix containing the desired price and trade effects. In computing $\boldsymbol{E}$ we set ε_N alternatively to 0.44 and 1.60 to assess the impact of the feed quota. For brevity, the price and trade effects for the Chilean tariff are not reported as they are qualitatively similar to those of the Norwegian tariff.

Results support the hypothesis that the feed quota contributes to instability in both prices and trade volumes (table 4).

[13]The implied horizontal shift, i.e., the shift in the quantity direction with price held constant, is 95%. The price-constant supply increase of 95% is larger than the actual supply increase of 71% due to the depressing effects of the 40% price decline on Chile's production.

Table 4. Simulated Price and Quantity Effects of Chile's Supply Expansion and US Tariffs on Salmon Imports from Norway with and without Norway's Feed Quota

Endogenous Variable	Supply Expansion Effect		Ratio	Tariff Effect		Ratio
	With Feed Quota (%)	Without Feed Quota (%)		With Feed Quota (%)	Without Feed Quota (%)	
P_N*	-9.7	-7.8	1.26	-35.2	-28.0	1.26
P_{CH}*	-10.7	-8.5	1.26	4.2	12.1	0.35
P_{UK}*	-9.1	-7.2	1.26	3.5	10.2	0.35
P_{CAN}*	-5.9	-4.7	1.26	2.3	6.7	0.35
P_{FI}*	-10.9	-8.7	1.26	4.2	12.2	0.35
$\sum_{i=1}^{6} kx_i P_i^*$	*-9.7*	*-7.7*	*1.26*	*-16.9*	*-9.7*	*1.73*
P_{EU^*}	-9.5	-7.5	1.26	3.7	10.7	0.35
P_{US}*	-6.5	-5.2	1.26	2.5	7.4	0.35
P_J*	-5.7	-4.5	1.26	2.2	6.4	0.35
$\sum_{j=1}^{4} km_j P_j^*$	*-8.2*	*-6.5*	*1.26*	*3.2*	*9.3*	*0.35*
X_N*	-4.3	-12.4	0.35	-15.5	-44.8	0.35
X_{CH}*	77.6	81.1	0.96	6.8	19.8	0.35
X_{UK}*	-35.2	-28.1	1.26	13.7	39.7	0.35
X_{CAN}*	-10.7	-8.5	1.26	4.2	12.0	0.35
X_{FI}*	-16.6	-13.2	1.26	6.5	18.7	0.35
$\sum_{i=1}^{6} kx_i X_i^*$	*11.7*	*9.3*	*1.26*	*-4.5*	*-13.1*	*0.35*
M_{EU}*	12.4	9.9	1.26	-4.8	-14.0	0.35
M_{US}*	9.1	7.3	1.26	-3.6	-10.3	0.35
M_J*	17.6	14.0	1.26	-6.8	-19.8	0.35
$\sum_{i=1}^{4} km_j M_j^*$	*11.7*	*9.3*	*1.26*	*-4.5*	*-13.1*	*0.35*

Note: to conserve space price and quantity effects for ROW are not presented.

Specifically, without the feed quota Chile's supply expansion would have caused export prices worldwide to decline by 7.7% and import prices to decline by 6.5%; with the quota the estimated declines are 9.7% and 8.2% respectively. Similarly, without the feed quota Chile's supply expansion would have caused exports and imports worldwide to increase by 9.3%; with the quota the estimated increase is 11.7%. Overall, these estimates suggests the feed quota enlarged price and trade volume fluctuations by 26%.

Results also support the hypothesis that the feed quota undermines US tariff policy. In particular, without the feed quota the rise in US price associated with the 26.1% Norwegian tariff is 7.4%; with the quota the price rise dwindles to 2.5%. The reason for the tiny price effect is that most of the tariff's incidence is borne by Norwegian producers. Specifically, the decline in the Norwegian price associated with the tariff increases from 28.0% to 35.2% as the feed quota is applied. Hence, the feed quota shifts more of the tariff's incidence onto Norwegian producers, but also widens the price wedge (from 35.4% to 37.7%).[14] These estimates are consistent with the analyses by Asche (2001) and by Kinnucan (2003), which suggest anti-dumping tariffs do more to punish producers in the targeted exporting countries than to assist producers in importing countries.

Tariffs provide an implicit subsidy to exporters not subject to the tariff. However, in the present case the implicit subsidies are modest, with no single exporter enjoying a price rise greater than 4.2% due to the Norwegian tariff (table 4). (For the 5.5% Chilean tariff, the largest price rise enjoyed by Chile's international competitors is 1.7%.) Still, removal of the feed quota would increase the implicit subsidy by a factor of three; hence, foregone spillover benefits for UK producers in particular are not inconsequential.

Overall, Chilean supply expansion reduced export prices worldwide by 9.7% and increased imports worldwide by 11.7% (table 4). This implies that the "total" import demand curve (Pearce 1952; Buse 1958) for salmon is elastic at -1.2. Hence, supply control measures that reduce exports are counterproductive in that they erode export value. Indeed, the added supplies from Chile over the period in question *raised* export value by 2%.

WELFARE EFFECTS

To place welfare measures on the unintended consequences, we used the following formulas:

Producer impacts:

$$\Delta PS_{CH} = P_{CH}\, X_{CH}\, (P_{CH}^* - \gamma_{CH})(1 + ½\, X_{CHi}^*) \tag{25a}$$

$$\Delta PS_i = P_i\, X_i\, P_i^*\, (1 + ½\, X_i^*) \qquad i = N, UK, CAN, FI, ROW \tag{25b}$$

Consumer impacts:

$$\Delta CS_j = -\, P_j\, M_j\, P_j^*\, (1 + ½\, M_j^*) \qquad j = EU, US, J, ROW \tag{26}$$

Total impact:

$$\Delta TS = \textstyle\sum_{i=1}^{6} \Delta PS_i + \sum_{j=1}^{4} \Delta CS_j \tag{27}$$

[14]The widening of the price wedge is due to the non-unitary price transmission elasticities (see footnote 2). If all price transmission elasticities in the model are set to one, the simulated price wedges equal the respective tariffs, and are unaffected by the feed quota. Incidence, however, is affected in the manner indicated.

where ΔPS_i is the change in "producer" surplus in the ith exporting country, ΔCS_j is the change in "consumer" surplus in the jth importing country; and ΔTS is the change in total surplus.[15] In these equations P_i and P_j are average export and import prices, respectively, over the evaluation period as reported in table 1, and X_i and M_j are the corresponding export and import quantities.[16] The asterisked variables in equations (25) - (27) indicate the price and quantity impacts of the exogenous variables computed from equation (24). Since an increase in Chile's supply reduces the demand for salmon from competing exporters, equation (25b) measures competing supplier impacts by shifting the respective excess demand curves along each exporter's stationary excess supply curve. Conversely, the own impact of supply expansion is measured via equation (25a) by shifting Chile's excess supply curve along its stationary excess demand curve. (The term (P_{CH}* - γ_{CH}) > 0 in equation (25a) indicates the extent to which the cost reduction associated with the supply shift exceeds the price reduction.) Consumer gains are measured via equation (26) by shifting the excess supply curves facing the importing countries along their respective stationary excess demand curves. All gains and losses are reported in constant (2001) US dollars.

Supply Expansion Effects

Chile's supply expansion generates a total surplus gain of $1.30 billion (table 5). Most of the gain ($1.05 billion) accrues to Chilean producers, as might be expected since the implied cost reduction (58.1%) far exceeds the associated price decline (10.7%, see table 4). With lower prices consumers gain $771 million. The losers, of course, are Chile's international competitors, who suffer a combined loss of $525 million. Most of this loss, $381 million, is absorbed by Norway's producers, with Canadian and UK producers a distant second at $49 million apiece.

Removal of the feed quota leaves the total welfare gain unchanged at approximately $1.30 billion, but shifts the incidence in favor of producers. Specifically, with the feed quota in place the producer share of the $1.3 billion total welfare gain is 41%; with the feed quota removed the producer share of the gain rises to 54%. The reduced consumer incidence may be explained by the fact that quota removal makes supply more elastic, which reduces the price decline associated with Chile's supply expansion. With price reduction attenuated, consumers gain less, and Chile's international competitors lose less, both of which help to tilt total welfare gain in favor of producers. Overall, estimates in table 5 suggest the feed quota enlarged benefits to consumers from the supply shift by 28%, but at a cost in increased losses to producers (exclusive of Chile) of between 20% and 31%. Losses to UK producers, the major focus of this analysis, decline from $49 million to $40 million with quota removal. Thus, for the period in question UK producers appear to have suffered excess losses from Chilean supply expansion due to the feed quota of about 20%.

[15] Technically, surplus changes are measured off *excess* supply and demand curves and thus the "producer" and "consumer" surpluses are properly interpreted as "exporter" and "importer" surpluses. However, since trade accounts for the bulk of production or consumption in most of the regions (see table 2), the distinction in the present case is unimportant.

[16] A reviewer questioned whether there was any implications from using average prices, noting that for about half the regions prices bottomed in 2001. The only implication is a modest loss in precision relative to measures based on annual prices.

Table 5. Welfare Effects of Increased Salmon Supplies from Chile on Salmon Importers and Exporters with and without Norway's Feed Quota, 2000-2002

Item	Welfare Gain		Ratio
	With Quota	Without Quota	
Exporters:	---------------- (Million US $) ------------		
Norway	-381	-291	1.31
Chile	1,054	1,119	0.94
United Kingdom	-49	-40	1.20
Canada	-49	-39	1.24
Faroe Islands	-33	-27	1.23
ROW	-13	-11	1.21
All	529	711	0.74
Importers:			
European Union	399	314	1.27
United States	215	169	1.27
Japan	41	32	1.28
ROW	116	92	1.27
All	771	606	1.27
ΔTS	1,299	1,317	0.99

Tariff Effects

Results indicate that total welfare losses from the Chilean tariff ($155 million) are modest in relation to losses from the Norwegian tariff ($1.44 billion). Hence, discussion will focus on the Norwegian tariff unless indicated otherwise.

With the feed quota in place the lion's share of the total welfare loss, $1.30 billion, accrues to Norwegian producers (table 6). This should not be surprising given that 93% of the tariff's incidence is borne by Norwegian producers (recall table 4). The only beneficiaries of the tariff are Norway's international competitors, who, owing to the induced rise in their prices, enjoy a combined welfare gain of $138 million. Most of this gain, $72 million, accrues to Chilean producers, with UK benefitting by a more modest $24 million. The spillover benefits, however, are more than offset by consumer losses totaling $277 million. Hence, the tariff against imports from Norway exacts a high cost in relation to benefits.

Table 6. Welfare Effects of U.S. Tariffs on Salmon Importers and Exporters with and without Norway's Feed Quota, 2000-2002

Item	Tariff Against Norway			Tariff Against Chile		
	With Quota	W/O Quota	Ratio	With Quota	W/O Quota	Ratio
Exporters:	------ (Mil $) ------			------- (Mil $) -------		
Norway	-1,298	-869	1.49	61	49	1.25
Chile	72	222	0.33	-116	-121	0.96
United Kingdom	24	79	0.31	9	7	1.26
Canada	21	62	0.33	8	6	1.26
Faroe Islands	15	45	0.33	6	5	1.26
ROW	6	21	0.31	2	2	1.26
All	-1,160	-441	2.63	-29	-52	0.56
Importers:						
European Union	-143	-393	0.36	-62	-49	1.26
United States	-78	-219	0.36	-46	-37	1.25
Japan	-14	-38	0.37	-7	-5	1.25
ROW	-42	-115	0.36	-12	-10	1.25
All	-277	-766	0.36	-126	-100	1.26
ΔTS	-1,436	-1,206	1.19	-155	-152	1.02

Lifting the feed quota reduces the total welfare loss to $1.21 billion and redistributes losses away from Norwegian producers to consumers (table 6). The most dramatic effect in terms of the research objectives of this paper is the enlargement of the implicit subsidy conferred by the tariff on Norway's international competitors. In particular, UK producer benefits from the tariff rise from $24 million to $79 million when the quota is removed, for a net gain of $55 million. Comparing this gain with the $9 million gain from quota removal estimated in table 5, tariff effects are more consequential than supply expansion effects. Adding the tariff-based gain of $9 million to the supply expansion-based gain of $55 million yields an estimated gross gain to UK producers from quota removal of $64 million. This is a *gross* gain because, according to table 6, feed quota removal would reduce the spillover benefit that UK producers receive from the Chilean tariff by $2 million. Taking this loss into account, the estimated net gain to UK producers from quota removal is $62 million, equivalent to 9.5% of UK export revenue over the evaluation period.

SENSITIVITY ANALYSIS

As Bredahl, Meyers, and Collins (1979) suggest, price transmission elasticities have important implications for trade analysis. Accordingly, since the classical trade model assumes transmission elasticities are unity, we re-computed our welfare measures with the nine transmission elasticities in the model set to one to determine how inferences might be affected. Results show the *magnitude* of the impacts being affected, but not their *direction* (Table 7).

Table 7. Sensitivity of Welfare Effects to Price Transmission Elasticities

Item	Chile's Supply Expansion			US Tariff Against Norway		
	With Quota	W/O Quota	Ratio	With Quota	W/O Quota	Ratio
Exporters:	------ (Mil $) ------			------- (Mil $) -------		
Norway	-342	-268	1.28	-903	-655	1.38
Chile	1,113	1,162	0.96	39	119	0.33
United Kingdom	-47	-40	1.18	15	49	0.31
Canada	-70	-58	1.21	20	62	0.33
Faroe Islands	-27	-22	1.21	8	24	0.33
ROW	-12	-10	1.19	4	12	0.32
All	614	764	0.80	-817	-389	2.10
Importers:						
European Union	366	294	1.24	-89	-255	0.35
United States	291	234	1.24	-70	-202	0.35
Japan	65	51	1.26	-14	-39	0.36
ROW	129	104	1.24	-31	-89	0.35
All	851	684	1.24	-205	-584	0.35
ΔTS	1,465	1,448	1.01	-1,022	-973	1.05

Note: results are based on setting all price transmission elasticities in the model equal to one.

For example, the total welfare losses from the Norwegian tariff decline from $1.44 billion to $1.02 billion when the price transmission elasticities are set to one. However, the overall conclusion that the feed quota increases the tariff's cost to Norwegian producers and reduces spillover benefits to Norway's international competitors is unaffected. Similarly, the total welfare gain of Chile's supply expansion increases from \$1.30 billion to \$1.46 billion when the price transmission elasticities are set to one. But the overall conclusion that the feed quota enlarges gains to consumers at the expense of Chile's international competitors is unaffected.

The simulations confirm that price wedges equal tariff rates when price transmission elasticities are unity (26.1% for Norway and 5.5% for Chile). In the original model the price wedges were 37.7% and 8.4% (see footnote 2), which explains why tariff impacts implied by that model are larger than in the sensitivity analysis. Still, our overall conclusion that the feed quota entails significant costs in terms of unintended consequences is not much affected by the transmission elasticities.

Concluding Comments

The basic theme of this research is that trade restrictions can have unintended consequences. In the case of the feed quota instituted by Norway to assist United Kingdom salmon producers, unintended consequences include steeper price declines associated with increased supplies of salmon from Chile, and a magnified Norwegian incidence of US tariffs on salmon imports from Norway. The latter undermined the tariffs' ability to benefit US producers, and decreased the implicit subsidies enjoyed by Norway's international competitors as a result of higher world prices induced by the tariffs. For UK producers, these unintended consequences resulted in some \$62 million in foregone surplus between 2000 and 2002, equivalent to 9.5% of export revenue. Hence, the feed quota may well have been ineffectual in its main purpose, which was to benefit UK producers by restricting the supply of Norwegian salmon into the EU market.

This conclusion, coupled with the finding in this study that the general equilibrium demand curve for salmon is elastic at -1.2, suggests trade restrictions are not an effective policy instrument for assisting salmon producers. One policy alternative that may have merit is market promotion. Research suggests Chile is a major free rider on Norway's salmon promotion efforts (Kinnucan and Myrland 2003). Since these efforts are aimed chiefly at markets in the European Union and Japan, Chile could be encouraged to invest in promotion in the United States, perhaps in league with US producers. In addition to addressing the free-rider issue, the added promotion would help strengthen market prices, thereby moderating price declines associated with ongoing productivity improvements in salmon farming (Asche 1997; Anderson 2002). If Chile's promotion program were funded by an export tax, as is the case for Norway's program, the scheme would have the added benefit of reducing market supply. The point is that given the unintended consequences of supply control, demand-side policies might offer a better long-term solution to the problem of low or declining prices.

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APPENDIX: SUPPLY SHIFT MEASUREMENT

Marsh (2003, p.902) defined shifts in the US retail demand curve for beef as the "percentage differences between observed retail beef prices and estimated retail beef prices holding demand constant." This definition can be adapted to the measurement of shifts in Chile's excess supply curve by reference to Figure 1.

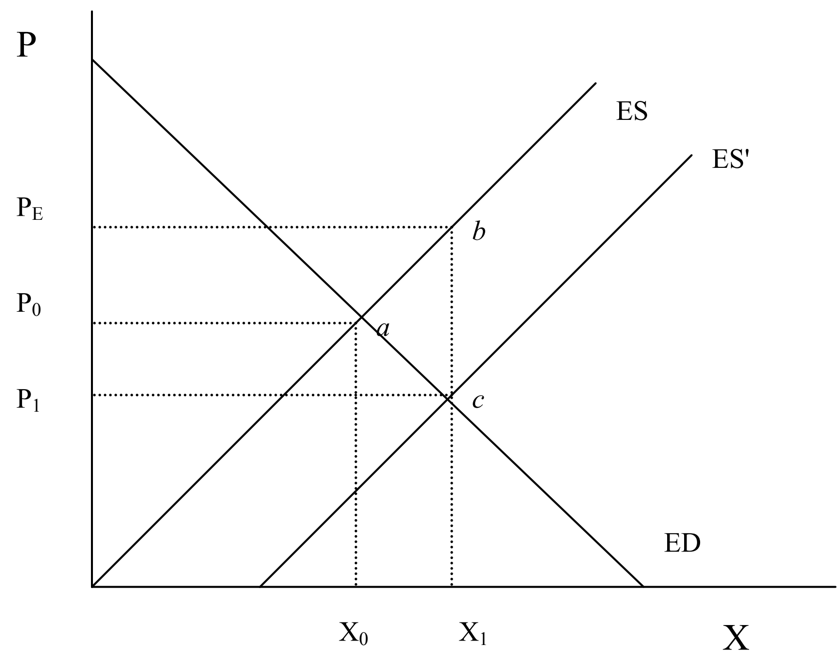

Figure 1. Measurement of the Vertical Shift in Chile's Excess Supply Curve

Let X_0 and P_0 represent Chile's observed exports and price in 2000 and let X_1 and P_1 represent the observed exports and price in 2002. The supply shift corresponding to these price/quantity coordinates is identified by letting P_E be the price that would obtain if the observed increase in exports was due to a demand shift rather than a supply shift. Specifically, in Marsh's terminology P_E is the estimated price holding supply constant, i.e., if demand increased along ES from point *a* to point *b*. Dropping a line from P_E to P_1 gives the absolute vertical shift in the excess supply curve, labeled in the diagram as distance $bc = |P_1 - P_E|$.

The corresponding relative vertical shift is:

$$\gamma = (P_1 - P_E)/P_E \tag{A1}$$

where $\gamma < 0$ since the shift is downward.

In Marsh's (2003) paper a numerical estimate for γ was obtained using a four-step procedure. Here, we derive an analytical expression that permits the calculation to be done in one step.

From Figure 1, $P_E = P_1 + (P_0 - P_1) + (P_E - P_0)$, which implies:

$$P_E = P_0 + \Delta P \tag{A2}$$

where $\Delta P = P_E - P_0$. Substituting equation (A2) into (A1) yields:

$$\gamma = (P_1 - P_0 - \Delta P)/(P_0 + \Delta P).$$

Multiplying the right-hand side of this expression by P_0/P_0 yields:

$$\gamma = (P^* - P_E^*)/(1 + P_E^*) \tag{A3}$$

where $P^* = (P_1 - P_0)/P_0$ is the observed percentage change in price and $P_E^* = (P_E - P_0)/P_0$ is the percentage change in price measured along the original excess supply curve *ES* from point *a* to point *b*. Following Marsh (2003), P_E^* can be approximated as follows:

$$P_E^* \approx X^*/\varepsilon$$

where $X^* = (X_1 - X_0)/X_0$ is the observed percentage increase in exports and ε is the excess supply elasticity corresponding to *ES* . Substituting the above expression into equation (A3) yields:

$$\gamma \approx (\varepsilon P^* - X^*)/(\varepsilon + X^*), \tag{A4}$$

which is identical to text equation (22).

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IMPLICATION OF U.S. RICE EXPORT PROMOTION PROGRAMS: THE CASE OF SELECTED LATIN AMERICAN COUNTRIES

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The U.S. domestic rice market will not likely be able to absorb projected increases in U.S. rice supply. Without export development, rice production would be constrained by limited U.S. domestic consumption. The main objective of the study is to analyze the effect of promotion programs and exchange rates with respect to rice exports to the three major U.S. long grain rough rice importers in Latin America. Ordinary Least Square (OLS) estimation is conducted as the first step. In addition, a Generalized Method of Moments (GMM) fitness test is conducted to correct the efficiency problem in the OLS. The study showed promotion expenditures to be significant only for Mexico and Honduras. The average return of each one dollar investment would result in $5 and $39 in the Mexican and Honduran markets using the Argentine exchange rate, and $3 and $45 in these respective markets using the Uruguayan exchange rate.

Key words: *Latin America, export promotion programs, U.S. rice exports.*

More than 20 percent of total agricultural products in the United States have gone to export markets during recent decades. Export markets have become an important element in government decision-making related to economic development in the U.S. agricultural industry. The value of exports for agricultural products has grown rapidly over the past two decades. U.S. agricultural export revenues accounted for 20-30 percent of U.S. farm income during the last 30 years and are projected to remain at this level (USDA/FSA, *Fact Sheet*).

Assuming constant or increasing returns to scale in rice production, combined with improved technical expectations, the domestic rice markets would not be able to absorb the total domestic rice supply (Figure 1). Without export development, rice production would be constrained by the domestic rice consumption level.

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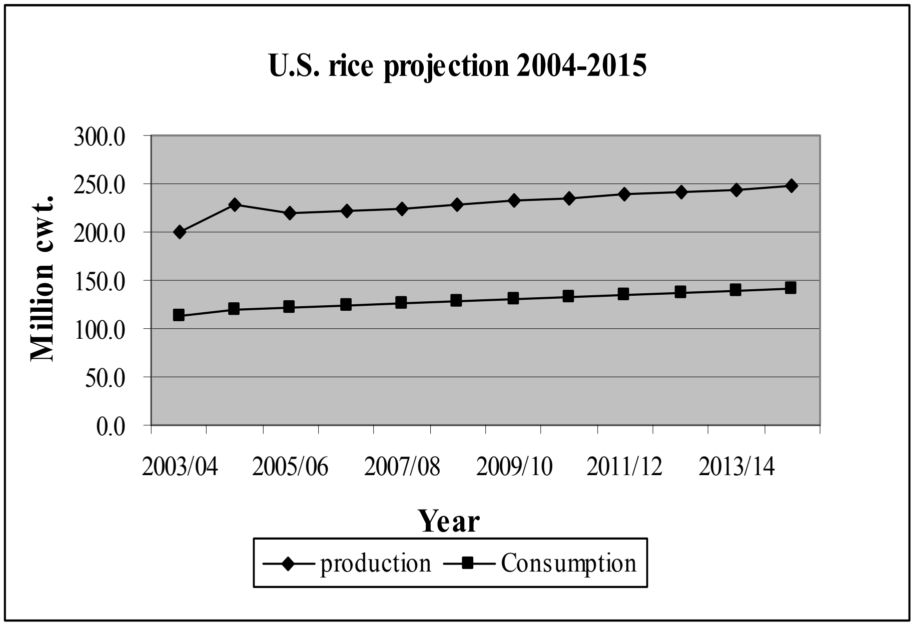

Source: USDA; 2005 USDA projections. Rice Outlook (2005), USDA

Figure 1. Rice Production and Consumption in the United States: 2004 - 2015

If the government fails to provide an acceptable support price, rice farmers could be driven out of business. Comparatively, the cost of a price-support policy would be much higher than the cost of export promotion projects, such as The Foreign Market Development Program (FMDP) and the Market Access Program (MAP). Those promotional policies for rice, with additional regional trade agreements and world trade agreements being considered, would create more opportunities for U.S. rice producers with an additional gain in social welfare.

Much of the literature focuses on specific commodities in different countries and regions and adopts various methodologies: seemingly unrelated regression (SUR) (Ward and Tang, 1978; Fuller, Bello, and Capps, 1992), single equation based econometric analysis (Binkely, 1981; Halliburton and Henneberry, 1995), and cross sectional time series analysis (Rosson, Hamming, and Jones, 1986; Le, Kaiser, and Tomek, 1998). Instead of focusing on specific methodologies, some studies analyzed the effects of exchange rates on U.S. export promotion program and expenditures (Schuh, 1975; Rosson, Adcock, and Hobbs, 2001; Adhikari et al., 2003).

The exchange rate is an essential element that influences the trade and competitiveness of U.S. agricultural products. Stable exchange would affect the long run U.S. agricultural economy. The existence of unpredictable fluctuation of exchange rates is an important source of risk to farmers and ranchers, processors, and traders. However, as a key force that influences the competitiveness of U.S. rice exports to the selected Latin American countries, competitors' exchange rates are considered in this study to understand how the competitors' exchange rate fluctuations would impact the U.S. rice exports to the selected Latin American countries. Empirically, own exchange rates are not included in the study because the prices of importing countries' rice are deflated by the importing countries' CPI and converted into U.S. dollar which reflects own exchange rate effects.

Over the past decade, the United States has lost substantial market share in Sub-Saharan Africa and the Middle East. However, rapid expansion in U.S. rough rice exports has offset much of the decline. Growth markets for rough rice exports are mainly in the Latin American region. Four countries in Latin America have been listed among the top ten U.S. rice buyers during the past 15 years. They are Mexico, Brazil, Costa Rica and Honduras. With the exception of Brazil, the others share common characteristics. U.S. rice exports to any of these three countries has taken more than 90% of the countries' import market share during the past fifteen years, along with an increasing import trend and a decrease in domestic production. Furthermore, during 1989-2003, the ratios of U.S. exports with respect to the total domestic consumption in Mexico, Costa Rica, and Honduras are 48%, 26%, and 60%, respectively, compared to approximately 2% for Brazil (USDA/FAS, *FATUS data base*). Therefore, this study focuses on these three countries: Mexico, Costa Rica, and Honduras.

The main objective of the study is to analyze the effect of promotion programs and exchange rates with respect to rice exports to the three major U.S. long grain rough rice importers in the selected Latin American Countries. Specific objectives are to develop a theoretical framework to analysis the economic factors which would affect the import demand in the three importing countries, to develop a model to estimate the import demand for Mexico, Costa Rica, and Honduras, and to evaluate the effectiveness of the promotion programs in expanding rice exports in the target markets.

The outline of the study is organized as follows: the next sections provide general information regarding U.S. rice trade with selected Latin American countries. The theoretic framework and procedures are then presented. Finally, the results and implications of the promotion programs are discussed.

U.S. Rice Trade With Selected Latin American Countries

Rice production and marketing is a multi-billion dollar activity in the United States. Rice is produced on over 3 million acres in the U.S. and accounts for $1.4 billion in farm revenues. Almost half of the rice produced in the U.S. is exported (USDA/ERS, *Outlook and Situation Yearbook*), which makes it critical for U.S. rice producers and exporters to fortify and develop international market activities (USDA/FAS, *Attaché reports*).

In terms of individual countries, U.S. exports of long grain rough rice to Mexico and Central America has become the major growth market. With none of the major Asian exporting countries allowing rough rice exports, the United States is the only major rice exporter that exports rough rice. The primary market for U.S. rough rice is Mexico, and Central American countries. Argentina, Uruguay, and Guyana ship small amounts of rough rice to Latin America. Under this situation, U.S. rough rice exports have expanded substantially since 1990/91, accounting for more than 30 percent in recent years (USDA/ERS, *Rice Outlook and Situation Year Book*).

The United States currently has a virtual monopoly on rice trade with Mexico, with an average of about 70 percent of Mexican rice imports coming from the United States during 1999-2003. Close geographic proximity and the existence of regional trade agreements, as well as sanitary and phytosanitary restrictions on Asian rice enacted by Mexico in 1993, solidify the monopoly power of the United States in this market (Zahniser and Link, 2002).

The North American Free Trade Agreement (NAFTA) benefits U.S. rice exports to Mexico, with a decreasing tariff for milled, rough, and broken rice. Before NAFTA, Mexico imposed import tariffs on U.S. rice, 20 percent on brown and milled rice and 10 percent on rough and broken rice. In 1990, the tariff rate for milled and brown U.S. rice was raised from 10 percent to 20 percent in response to demands from Mexican millers. Under NAFTA, Mexico was gradually lowering these rates to zero over the 9-year period that ended on January 1, 2003.

Honduras and Costa Rica are the primary U.S. rice importers in Central America. The United States has over a 90 percent share in both markets. Honduras had a dramatic decrease in its domestic production, which resulted from less government support and higher production costs compared to the price of importing rice. Rice imports from the United States increased during the past five years, with 90 percent of total domestic rice consumption being supplied by the United States. Costa Rica has kept a constant production level during 1989-2003, and domestic production is still a major source of satisfying domestic consumption. U.S. rice holds a nearly 99 percent import market share in Costa Rica.

Costa Rica and Honduras prefer rough rice imports to milled rice imports in order to support their domestic milling industries. According to the CAFTA agreement, Costa Rica gives a 50,000 metric ton (MT) duty-free quota to U.S. rough rice, increasing by 2 percent annually; the quota for milled rice is 5,000 MT, growing at 5 percent annually. Honduras provides U.S. rough rice a 90,000MT duty–free tariff-rated quota (TRQ) with 2 percent annual growth, and an 8,500MT duty-free TRQ for milled rice with 5 percent annual growth (USDA/FAS). Rice is a sensitive commodity to bring into trade negotiation, given that the domestic milling industries are highly protected in the targeted countries. This policy would provide little benefit to the U.S. milling industry although it provides the largest gains to U.S. rice producers. However, this current policy would benefit the U.S. rice industry if the sanitary restrictions are relaxed in the future. Asian exporters would still not receive market access, since they are not seeking to export rough rice.

U.S. RICE EXPORT PROMOTION PROGRAMS

The United States Department of Agriculture (USDA) has four general categories of export promotion: consumer promotion, trade servicing, technical assistance, and export credit guarantee programs. Consumer promotion focuses on the retail level market. The intent is to raise the consumers' awareness for the products, build lasting preference, and increase the final consumption level through marketing activities. The Foreign Market Development Program (FMDP) and the Market Access Program (MAP) are the primary source of support for federal rice promotion (USDA/FAS).

The Foreign Market Development Program (FMDP) was created in 1955 and includes the Cooperator Market Development program (CMDP) and the Export Incentive Program (EIP), also administered by the Foreign Agricultural Service (FAS) of the USDA. The goal of the program is to develop, maintain, and expand long-term export markets for U.S. agricultural products, primarily through trade service and technical assistance programs. The program fosters a trade promotion partnership between USDA and U.S. agricultural producers and processors who are represented by non-profit commodity or trade associations called cooperators. Under this partnership, the USDA and cooperators pool their technical and financial resources to conduct market development activities outside the United States.

Since 1993, the Market Access Program (MAP) replaced the former Market Promotion Program (1991-1993) and the Targeted Export Assistance Program (TEA) (1986-1990). The MAP is a cost-share program that uses Commodity Credit Corporation (CCC) funds to support U.S. producers, exporters, private companies, and other trade organizations' promotional activities for U.S. agricultural products.

The U.S. Rice Federation (USARF) annual report indicates that their promotional activities in Mexico have resulted in a 200 percent increase of U.S. rice consumption in restaurants and doubled the U.S. rice long grain exports to Mexico from 1998 to 2003, including a 24 percent increase in 2001-2003 (USARF, www.usarice.com).

Through industry groups like USARF and the U.S. rice producer association (USRPA), U.S. rice producers, millers and exporters have joined together to develop export promotion effectiveness, increasing the international market demand of a variety of U.S. rice forms and types, including rough, brown, white, parboiled, and long, medium, and short grain.

THEORETICAL FRAMEWORK

This study assumes that, based on the different domestic and general import situations (imports from United States and the rest of world), the efficiency of promotion programs would differ among these three countries. The primary hypothesis to be tested is that U.S. export promotion expenditures have had a positive impact on the rice imports of selected Latin American countries. To test this hypothesis, two sets of econometric models have been applied. First, the single equation model tests the effect of the promotional programs and competitors' exchange rate fluctuation for each country. Analysis of exchange rate fluctuations shows how the competitor's exchange rates affect U.S. rice exports to the selected Latin American countries. The single-equation structural demand equation has been a popular method and remains in use for three reasons. The first reason is that demand for the studied commodities can be modeled independently with variables deemed necessary to determine demand for the commodity. Second, the data required is quite flexible. The final advantage is computational ease. Another popular method in the literature is the Almost Ideal Demand system (AIDs). The independent variable of AIDs is market share (Jan, Huang, and Epperson, 1999). However, the U.S. share of rice imports in three of the four countries is over 90% in the period studied (1989-2003). This implies that, AIDs may not be appropriate in these cases.

Binkley (1981) suggested that single-equation methods are appropriate for estimating import demand when the supply faced by importers is exogenous, i.e., the importer is a price-taker. He also suggested that in many cases in which demand (supply) are estimated, use of single-equation methods are justified on the basis that because of the highly elastic nature of supply (demand), simultaneous effects are of no practical consequence. In addition, economic theory offers little guidance on appropriate measures of variables which are included in the import demand function or on the appropriate functional form. An appropriate model is defined as one which generates unbiased and efficient elasticity estimates. Therefore, the precise specification of import demand is largely an empirical issue (Thursby and Thursby). This condition could be recognized in international rice trade where most of the rice producing countries are the major consumers, and leave only a small margin to be traded on the world market. In previous research, a single equation model has been specified to analyze the impact of promotion programs on export demand for several agricultural commodities, such as grapefruit (Fuller, Bello, and Capps, 1992), almonds (Halliburton and Henneberry, 1995; Onunkwo and Epperson, 2001), pecans (Epperson, 1999; Florkowski and Timothy, 1999; Onunkwo and Epperson, 2000), and poultry meat (Jan, Huang, and Epperson, 2002).

Ordinary Least Square (OLS) estimation is normally conducted as the first step to observe the relationship between explanatory variables and dependent variables. It holds strong estimation power in econometric analysis. The collinearity, heteroscedasticity, and autocorrelation tests are conducted to determine the efficiency of the single equation system result. A Generalized Method of Moments (GMM) fitness test is conducted to correct the efficiency problem caused by heteroscedasticity in the OLS.

The models considered in the study for the single country analysis are shown in equations (1) and (2), respectively.

$$y_{i,t} = \beta_0 + \beta_1 tds_{i,t} + \beta_2 tdc_{i,t} + \beta_3 pr_{i,t} + \beta_4 pw_{i,t} + \beta_5 pro_{i,t} + \beta_6 uex_{i,t} + e_{i,t} \tag{1}$$

$$y_{i,t} = \beta_0 + \beta_1 tds_{i,t} + \beta_2 tdc_{i,t} + \beta_3 pr_{i,t} + \beta_4 pw_{i,t} + \beta_5 pro_{i,t} + \beta_6 aex_{i,t} + e_{i,t} \tag{2}$$

where the subscript t refers to time, and subscript *i* represents the importing countries, Mexico, Costa Rica or Honduras. β_0 is the intercept and $e_{i,t}$ is the error term. The dependent variable $y_{i,t}$ represents the total import amount of U.S. rice exports to selected country *i* in year *t*.

For the explanatory variables, $pro_{i,t}$ denotes the promotional expenditure invested on country *i* in year *t* deflated by the United States CPI index; $pr_{i,t}$ and $pw_{i,t}$, denote the unit import prices in importing countries' currency for rice and wheat, respectively, deflated by the importing countries' CPI and converted into U.S. dollars; $tds_{i,t}$ denotes the total importing countries' annual domestic supply, including the initial stock and the domestic production. $tdc_{i,t}$ denotes the total importing countries' annual domestic consumption. $uex_{i,t}$ and $aex_{i,t}$ denote the real exchange rates of the competing exporting countries' currency in term of the U.S. dollar (Uruguayan Pesos and Argentine Pesos; See Table 1 for variable description).

Table 1. Variable Description

Variable	Source	Unit	Description
Y	USDA/FATUS	1,000 MT	Volume of U.S. exports.
Pro	USARF and USRPA	$1,000	Annual promotional programs expenditure, adjusted by the U.S. CPI (Consumer Price Index) [a]. TEA/MPP/MAP and FMDP included.
Pr	USDA/FATUS	$/MT	Export unit value [b] of rice-paddy, milled, adjusted by the U.S. CPI [a].
Pw	USDA/FATUS	$/MT	Export unit value [b] of wheat, unmilled, adjusted by the U.S. CPI [a].
TDS	PSandD	1,000 MT	Total annual domestic rice supply of the importing country, the sum of initial stock and domestic production.
TDC	PSandD	1,000 MT	Total annual domestic rice consumption of the importing country.
UEX	ERS	index	The real annual exchange rate between the Uruguay pesos and U.S. dollars.
AEX	ERS	index	The real annual exchange rate between the Argentina pesos and U.S. dollars.

Notes: a. CPI base year 1995 = 100. b. The Unit Value is calculated by dividing the sum of the value by the sum of the quantity converted to the FAS unit of measure to three decimal places (USDA/FATUS)

DATA

Annual data for the period 1989 through 2003 were used to estimate the export promotion effects. The Federal promotion program expenditure data (FMDP and MAP) used in the model was obtained from the USARF and USRPA, who are the only two recipients of the USDA/FAS promotional programs funds during the study period.

The data represent the aggregate expenditures distributed by the United States Department of Agricultural (USDA). Since MAP programs are processed by market-year (July-June) and FMDP programs are processed by calendar year, the MAP data were adjusted into a calendar-year basis data set by applying a two-year moving average method. Then, the total MAP and FMDP expenditures are summed to represent the rice export promotional expenditure. Only federal promotion program investments were included in the study.

The rice export, import, and unit values are obtained from the publication Foreign Agricultural Trade of the United States (USDA/FATUS). The export unit value derived from the total import value and quantity of each importing country is used as a proxy for the U.S. rice export price in each market. The CPI for the importing countries and the United States are collected from International Financial Statistics, published by International Monetary Fund.

ESTIMATION RESULTS

The parameter estimation of the export demand equation for U.S. rice is shown in Tables 2-4 for Mexico, Costa Rica, and Honduras, respectively. In general, F-values for most of equations are significant, and the measures of goodness-of-fit for the estimated equation are high. The adjusted R^2 value for Mexico, Costa Rica and Honduras are 0.97, 0.93, and 0.97, respectively.

The Durbin-Watson statistic tests and the collinearity diagnostics tests do not indicate problems with autocorrelation and collinearity for the three markets. The heteroskedasticity test is obtained from the White test statistic. The null hypothesis for the test is that the variance of the residuals is homogenous. However, the chi-square value ranges from approximately 10.0-13.0, compare with the critical value 27.58. This indicates that there is a heteroskedasticity problem for the model. With the presence of heteroskedasticity, the covariance matrix is incorrect; the estimation is unbiased and consistent, but inefficient.

Table 2. Parameter Estimation for Mexico

	OLS		GMM	
	Equation (1)	Equation (2)	Equation (1)	Equation (2)
Intercept	-688.29*	-991.02**	-688.29**	-991.02**
	(-2.07)	(-2.95)	(-3.02)	(-2.54)
Rice Export Price	-0.78	-0.72	-0.78*	-0.72*
	(-1.50)	(-1.16)	(-2.09)	(-1.91)
Wheat Export Price	1.08**	1.54**	1.08**	1.54***
	(2.32)	(2.92)	(2.65)	(4.25)
Total Mexican Domestic Rice Supply	-0.29	-0.435	-0.29	-0.435
	(-1.03)	(-1.24)	(-0.96)	(-1.01)
Total Mexican Domestic Rice	1.48**	1.70***	1.48***	1.70**
Consumption	(3.26)	(3.38)	(4.06)	(2.45)
Promotion Expenditures	0.014	0.009	0.014*	0.009
	(1.59)	(0.90)	(1.85)	(1.08)
Exchange rate	77.62**		77.62***	
Argentina Pesos: U.S. dollars	(3.34)		(6.00)	
Exchange rate		11.55**		11.55**
Uruguay Pesos: U.S. dollars		(2.52)		(2.30)
Durbin-W	2.806	2.611	2.806	2.611
F value	65.37***	48.61***	-	-
Adj. R^2	0.9650	0.9533	0.9650	0.9533

Note: Single asterisk (*), double asterisk (**), and triple asterisk (***) denote rejection of H_0 at 0.10, 0.05, 0.01 significance levels respectively

Table 3. Parameter Estimation for Costa Rica

	OLS		GMM	
	Equation (1)	Equation (2)	Equation (1)	Equation (2)
Intercept	-193.69**	-263.42**	-193.69**	-263.42**
	(-2.33)	(-3.01)	(-2.50)	(-3.28)
Rice Export Price	-0.08*	-0.094*	-0.08***	-0.094***
	(-1.84)	(-2.27)	(-3.96)	(-5.17)
Wheat Export Price	0.53**	0.65***	0.53***	0.65***
	(2.82)	(3.58)	(4.12)	(4.17)
Total Costa Rica Domestic Rice Supply	-0.22	-0.13	-0.22	-0.13
	(-0.94)	(-0.60)	(-1.38)	(-0.85)
Total Costa Rica Domestic Rice Consumption	1.02***	1.06***	1.02***	1.06***
	(3.40)	(3.76)	(3.32)	(3.85)
Promotion Expenditures	0.17	0.18	0.17	0.18
	(1.27)	(1.44)	(0.98)	(1.08)
Exchange rate Argentina Pesos: U.S. dollars	15.64*		15.64**	
	(1.97)		(2.81)	
Exchange rate Uruguay Pesos: U.S. dollars		3.45**		3.45***
		(2.31)		(4.10)
Durbin-W	2.983	3.188	2.984	3.188
F value	8.66***	9.91***	-	-
Adj. R^2	0.9265	0.9325	0.9265	0.9325

Note: Single asterisk (*), double asterisk (**), and triple asterisk (***) denote rejection of H_0 at 0.10, 0.05, 0.01 significance levels respectively

With error variance relationship unknown, the GMM method is applied as a remedial measure to solve the problem. The correction to the standard error uses the Newey West correction.

For the most part, the estimated parameters displayed signs consistent with prior expectations. The own-price parameters are consistently negative and significant for the three countries; the demand and price are inversely related. The price for wheat has a positive sign, which indicates that wheat is a substitute for rice in these three markets. Generally, the import demand of the three countries is more responsive to own-price than promotion expenditures. The estimation of the wheat price suggests that wheat is an elastic substitute for rice in Mexico and Costa Rica, but less elastic for the Honduran market. This result partially supports a previous study which focuses on Mexican rice import markets (Salin et al., 2000).

The results indicate a negative relationship between total domestic supply and export demand. The greater domestic production and initial stocks, the less that rice would be demanded from the foreign market. Honduras shows a significant relationship between the domestic supply and import demand (-0.34/-0.56). The total domestic rice supply also shows a negative relationship in Mexico (-0.29/-0.435) and Costa Rica (-0.22/-0.13). The positive parameter estimation for total domestic consumption shows the demand for rice imports would increase with an increasing total domestic consumption level.

Table 4. Parameter Estimation for Honduras

	OLS		GMM	
	Equation (1)	Equation (2)	Equation (1)	Equation (2)
Intercept	-31.62	-61.19*	-31.62	-61.19*
	(-0.89)	(-1.87)	(-1.22)	(-2.23)
Rice Export Price	-0.12**	-0.11**	-0.12**	-0.11**
	(-3.12)	(-3.24)	(-2.57)	(-3.20)
Wheat Export Price	0.03	0.084	0.03	0.084
	(0.18)	(0.65)	(0.23)	(0.86)
Total Honduras Domestic Rice Supply	-0.34	-0.56**	-0.34*	-0.56***
	(-1.42)	(-2.46)	(-2.04)	(-3.73)
Total Honduras Domestic Rice	1.83***	1.89***	1.83***	1.89***
Consumption	(6.83)	(7.98)	(9.33)	(10.09)
Promotion Expenditures	0.033	0.042	0.033**	0.042**
	(1.26)	(1.77)	(2.46)	(2.68)
Exchange rate	8.79*		8.79**	
Argentina Pesos: U.S. dollars	(2.04)		(2.75)	
Exchange rate		1.29		1.29***
Uruguay Pesos: U.S. dollars		(1.56)		(4.01)
Durbin-W	2.850	2.951	2.850	2.951
F value	64.82***	81.86***	-	-
Adj. R^2	0.975	0.972	0.965	0.972

Note: Single asterisk (*), double asterisk (**), and triple asterisk (***) denote rejection of H_0 at 0.10, 0.05, 0.01 significance levels respectively

The promotion expenditures are positive and significant in the Mexican and Honduran markets. This indicates that each dollar of promotion expenditure generated a positive quantity of rice exports from the U.S. rice exporters to those two importing countries. However, it is important to note that only the federal government money is included, which may exaggerate the promotion programs' effect. If the promotion programs' expenditure is the same as the private group expenditures, the return of promotion programs could be divided by two to obtain the actual dollar return per dollar invested.

The exchange rate of the competitor is positive and significant. This suggests that a strong competitor currency would benefit U.S. rice exporters. The strong currency would increase U.S. exports. If we assume that rice produced in Uruguay and Argentina is the same quality as U.S. rice, if price differences are not large, then an appreciation of the Uruguayan and Argentine pesos would increase the demand for U.S. rice. Both countries are experiencing a strong currency against U.S. dollars, which may be part of the reason the United States holds such a large market share in the three importing rice markets after competition from Asian markets is eliminated.

Table 5. Elasticities Estimation for Own Price and Export Promotion

		Own price elasticity	Promotion Expenditure Elasticity
Mexico	UEX[a]	-0.97*	0.0134**
	AEX[b]	-0.96*	0.0240*
Costa Rica	UEX[a]	-5.15***	0.153
	AEX[b]	-5.09***	0.136
Honduras	UEX[a]	-0.838**	0.036*
	AEX[b]	-0.62	0.031*

Notes: Elasticities are calculated by the method of OLS with log-log model, show in each cell in three lines from up down respectively. Single asterisk (*), double asterisk (**), and triple asterisk (***) denote rejection of H_0 at 0.10, 0.05, 0.01 significance levels respectively, a: The elasticities are derived from the model with Uruguay exchange rate, b: The elasticities are derived from the model with Argentina exchange rate

The elasticities for own price and promotion are derived through transforming the previous model into log-log form. The results are shown in Table 6. The own price elasticities for Mexico are -0.97 and -0.96 for the model using the Uruguayan exchange rate and the Argentine exchange rate, respectively. Every one percent increase of the promotion programs would result in a 0.96 percent decreasing in the import demand. The promotion expenditures elasticities differ in the two equations, which are 0.0134 and 0.0240 for Uruguayan and Argentine equations, respectively.

For Honduras, the own price elasticities are -0.828 and -0.615, and the promotion expenditure elasticities are 0.036 and 0.031 for Uruguay and Argentina equation respectively. Costa Rica has very high price elasticity levels, with -5.15 and -5.09 for Uruguayan and Argentine exchange rates, respectively. The Costa Rica market acts more sensitively than the other two, since the U.S. rice import ratio only account an average 26 percent in the studying period. The promotion elasticity is higher than other two markets, 0.153 and 0.136 for Uruguay and Argentina, respectively.

The promotion expenditure return could be calculated using the elasticities obtained from this section. The average marginal return on promotion expenditure for each country could be obtained by multiplying the elasticity by mean value of export expenditures, and then dividing by the mean value of export promotion expenditures. Based on the elasticities derived in table 5, the return ranged from 3 to 5 dollars for the Mexican market, and the from 39 to 45 dollars for the Honduran market with respect to each one dollar investment in promotions from 1989-2003. However, this method simply provides the gross return without considering the production and export cost, and is overestimated by not including the private sector's promotion investment.

Table 6. Estimate of Lag Distribution: Decaying Promotion Effect

	Mexico		Costa Rica		Honduras	
Lag variable	Estimate	t value	Estimate	t value	Estimate	t value
lPro_0[a]	0.42	1.77	-0.14	-1.1	0.19***	108.07
lPro_1	0.25	1.01	0.02	0.33	0.07***	53.59
lPro_2	0.14	0.52	0.12	0.85	-0.01***	-7.29
lPro_3	0.08	0.36	0.13	0.91	-0.06***	-33.86
lPro_4	0.08	0.77	0.07	0.81	-0.07***	-44.23

Note: lPro_0, lPro_1, lPro_2, lPro_3 and lPro_4 represent the lag distribution from 0 to 4.

PROMOTION PROGRAMS LAGS

Lags in consumer behavior in response to advertising programs would occur between current sales and future sales. The effect of the advertising programs is assumed to show significant not only in the initial year, but has a finite declining procedure. Because of the habit persistence and lags in consumer behavior, time lags must be considered in the model. Distributed lags analysis is a specialized technique for examining the relationships between import demands and advertising programs that involve some delay.

As with other domestic advertising, international promotion programs also hold a decaying effect. A consistent investment in promotion programs would effectively increase the rice consumption demand and maintain the market share in the targeted importing countries in long run. Given factors in addition to the domestic market, the international promotional programs could require more work to analyze.

The polynomial distributed lag models are applied to the analysis of the impact of generic promotion on the demand for rice in three targeted markets. The current year promotion programs have much more impact on the rice import demand for Mexico and Honduras, then decay for the following years. For Costa Rica, there has been a slow lag time for the promotion programs to take effect. A possibility for this result is that the Costa Rican market is a mature market for rice consumption with a much higher per capita consumption level than compared to the other countries. A mature market could already have its own taste, consumption patterns, and loyalty to a certain type of rice. It would take a longer time for the new supplier to adjust to the market and learn its characteristics. The strategies to gain market share in such a market would be a consistent persuasive advertising pattern to attract and switch the consumer appetites and purchasing habits. However, only the estimated results for Honduras are significant, with the first year promotion effect of 0.19, decreasing to 0.07 in the second year. The insignificance of the results maybe due to data limitations; a short time-series of fifteen observations would inflate the error term and decrease the t value (Table 6).

IMPLICATIONS

The study showed promotion expenditures to be significant only for Mexico and Honduras. The average returns to Honduras show effectiveness of the promotion programs. This does not necessarily imply that promotion programs do not work in the other two markets. Given the returns over the promotion expenditure, the promotion programs should be continued in the targeted markets. The increase in the demand for U.S.-grown rice in those markets would continuously become a source of expanding U.S. rice sales.

Also, for markets with different levels of per capita consumption, the promotion programs should be designed separately. For the Mexican and Honduran markets, the promotion programs should continuously focus on promoting demand for U.S. rice. In markets which already consume large quantities of rice, the emphasis should be left on gaining more market share from local producers and other rice exporters. Given this, further study focusing on the price elasticity and other exporters' competitiveness would be important.

The U.S. rice exports to the three targeted countries, Mexico, Honduras and Costa Rica, hold over 90 percent of the import markets in recent years. If this situation changes due to the release of the previous sanitary restriction on Asian-grown rice, the Asian exporters could become strong competitors in those markets. The existing cost and price advantage of Asian rice exporters may shrink the current U.S. rice market share in the Latin American market. If the rice imported from the Asian markets are substitutable for U.S. rice, then price would become a dominant decision-making element for consumers and importers in importing countries. However, there is no indication as to when this sanitary restriction could be eliminated. Even if this restriction were eliminated, Asian access to this market may be limited due to preferential treatment received by U.S. producers under the NAFTA and CAFTA agreements. Given that rice is a sensitive commodity in the region, it is difficult to predict trade liberalization that may occur through the World Trade Organization negotiations.

MAP and FMDP focus more on the generic advertising than on brand sales, which means that although the consumption demand increased, the advertisers may not benefit from the money spent on promotion. Local rice producers and other exporters could easily act as free-riders and reduce the newly-gained market share. It may give promotion expenditure more flexibility by allowing promotional monies to be spent not only on generic advertising but also targeted at the brand level. Brand loyalty could be a more consistent motive than the general knowledge of the benefits of rice, such as better nutrition, ease of cooking and combining to local food taste.

CONCLUDING REMARKS

An import demand equation for the U.S. rice was estimated, with special focus on the two major export promotion programs: Foreign Market Development Program (FMDP) and Market Access Program (MAP). The Ordinary Least Squares method was used to estimate the equation based on annual exports to Mexico, Costa Rica and Honduras from 1989 to 2003. These countries represent the major export markets for U.S. long grain rough rice, which became the fastest increasing rice export market. Their preference for rough rice distinguishes their position in the world rice market, which would benefit U.S. rice exports.

The empirical analysis consists of three parts. The first part contains the regression estimation results and the interpretation of the promotion expenditure and exchange rate pass-through from competitor's currencies. In the second part, the elasticities for the own price and promotion expenditure were derived. Finally, the lagged effects of the promotion programs are tested to investigate the trend of the long-run decaying positive promotion effects. The empirical evidence from the regression analysis is supportive of the promotion programs and exchange rate effect of competitors' currency fluctuation with respect to U.S. dollars.

In the econometric model, the validity of the promotion effect for those importing countries was estimated. In addition, various lag lengths were considered, in addition to the original value of the promotion expenditures, in an attempt to test the promotion effect in the current year and long run. The results indicate a tendency for promotion to have the highest effect on current import demand in the Mexican and Honduran markets. In the case of Costa Rica, the promotion programs may come into effect in the long run. The results generally support the hypothesis that the promotion programs have a positive impact on U.S. rice exports to Mexico and Honduras, and Costa Rica to a lesser extent. Nevertheless, future research may include the supply and demand systems in world long grain rice trade, which would provide further understanding of the rice markets in the Latin American region. Details on the long grain rice production cost in the exporting countries and importing countries, the transportation costs, trade policies, and importing market's rice market development could be considered in future studies of the market.

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ENDOGENOUS TRADE PROTECTION IN THE MEXICAN CORN MARKET

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ABSTRACT

A conceptual structural model of endogenous trade protection is specified that links the international market to political factors influencing administrative trade barriers. The general objective is to increase the empirical knowledge of endogenous protection on agricultural trade. The specific objectives are to quantify the impact of corn import permits on international marketing margins between the U.S. and Mexico for corn and sorghum. Empirical findings support the conclusion that a government's policy of allocating import permits effects imports from the U.S. to Mexico.

Keywords: *Agricultural Trade; Protection; Trade Uncertainty; Marketing Margins.*

INTRODUCTION

It is widely recognized that technical trade barriers create numerous obstacles to the international exchange of agricultural products (Ndayisenga and Kinsey 1994; Roberts 1999; Weyerbrock and Xia 2000). The export of agricultural products from one country to an importing country normally requires the guarantee that the goods comply with all sanitary and phytosanitary requirements, as well as other requirements that the importing country might stipulate. These non-health related requirements range from moisture content to packaging standards.

Technical trade barriers are often attributed to growing demands by consumers for food safety, environmentally sound products, product differentiation, and product information (labeling). However, as tariffs are continuously lowered and more overt forms of protection are increasingly becoming illegal, many technical trade barriers have arisen because of the

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perception that they can be used to protect domestic producers from imports. Rules governing the use of technical trade barriers are in part set forth in the World Trade Organization (WTO) Agreement on the Application of Sanitary and Phytosanitary Measures and the WTO Agreement on Technical Barriers to Trade. These agreements give countries the flexibility to apply technical requirements to ensure the quality of exports, the protection of human, animal, and plant life or health, the protection of its environment, and for the prevention of deceptive practices. The agreements also stipulate that technical measures should not constitute a disguised restriction on international trade or be applied in an arbitrary or unjustified discriminatory manner.

Technical barriers can be very obscure with their effects hard to trace. Roberts (1999) suggests that economic analysis could advance the understanding of technical trade barriers especially in such areas as net cost of the barrier and distinction between justifiable and unjustifiable technical barriers. Most models of trade uncertainty in agricultural products due to technical trade barriers are theoretical in nature (Gallagher 1998; Ker 2000). Several give an overview of the technical trade barriers encountered in agricultural trade (Ndayisenga and Kinsey 1994; Roberts 1999; Weyerbrock and Xia 2000). Nevertheless, economic literature is sparse on empirical research about technical barriers. The empirical literature on technical barriers mostly consists of case studies that together provide only fragmentary evidence of the costs to the international economy (Ndayisenga and Kinsey 1994).

Many barriers to free trade can be attributed to the political process (Pincus 1975; Brock and Magee 1978; Anderson 1995; Hillman 1982; Hillman and Ursprung 1993; Goldberg and Maggi 1999). Grossman and Helpman (1995) concluded that politically motivated governments tilt negotiations in trade talks with other governments toward their most organized interest groups. Protection can be particularly high when the domestic interest group is strong and the foreign interest group is weak in relative strength. More often than not empirical studies ignore the political implications of trade protection.

Mexico uses an administrative trade barrier (import permit) to protect its producers from U.S. corn imports, potentially hindering the trade flow between the two countries. The issuance of these import permits is a highly political process in which different lobbying groups put pressure on the Mexican government to begin or stop issuing import permits. Politicians then decide to whom and when the licenses are issued. Uncertainty caused by these arbitrary decisions is thought to have an adverse effect on U.S. exporting firms and Mexican importing firms.

Our article contributes to the empirical literature on endogenous trade protection in several ways. Focusing on Mexican corn imports from the U.S., we develop a structural model of international marketing margins for an agricultural commodity and trade uncertainty that links behavioral processes from the private market to political factors influencing administrative trade barriers. In doing so, this systematically links trade models specified by Gallagher (1998) which characterize private markets under uncertainty but ignore direct influences from political markets, to work by Trefler (1993), Gawande and Bandyopadhyay (2000), and Grossman and Helpman (1994), which focus on endogenous trade protection. In other words, instead of modeling trade uncertainty induced in part by political processes in a purely probabilistic manner, the behavioral processes that determine the likelihood of trade are identified using concepts of trade protection and public choice analysis. The main objective of this article is to quantify the marketing margin attributable to technical trade barriers in the Mexican corn market. The intent is to show that a trade model with endogenous protection can help better explain uncertainty in the trade of agricultural commodities.

PROTECTION AND IMPORT PERMITS

The ratification of NAFTA in 1994 lowered tariffs on many agricultural products, but also allowed for special cases and exemptions. Corn was one of those special cases. When the negotiations ended, U.S. corn had restricted access into the Mexican market until 2009 at which time the quota restriction is scheduled to end. The tariff imposed on imports above the established quota was designed to be cost prohibitive. On the other hand, Mexico had to allow a three percent cumulative increase in the amount of corn imported every two years from the ratification of NAFTA (Josling 1997).

The Mexican government has a history of controlling agricultural imports through tariffs, licensing requirements, and other administrative barriers (Smith 1984; Mielke 1984; Josling 1997). Its preferred method of protecting the domestic corn market is to control the timing and quantity of imports through import permits. The Secretariat of Commerce and Industrial Development (SECOFI) has overall control of the licensing program, but issuance of import permits is done through a Tariff Commission or Cupo (import permit) committee.

The Tariff Commission meets on a quarterly basis. Import permits are issued to individual firms for a specific amount of time and quantity. The permits are classified by usage, i.e. white corn is for human consumption and yellow corn is for industrial use. Import companies are given differing amounts per permit depending on their lobbying efforts and their previous purchase of domestic production. For example, a corn processor who purchased a large amount of domestic production is more likely to receive an import permit on a consistent basis for the quantity they require.

The Mexican government uses import permits not only to distribute the quota among importers but also to avoid displacing domestic production (Mielke 1989).[17] However, one of the main problems of this licensing policy and the basis of this article is that the terms of the import permit are subject to change without notice, thereby adding uncertainty to the trade flow of corn between the U.S. and Mexico. This added trade uncertainty decreases export quantities in the U.S. and is hypothesized to increase import prices in Mexico.

Despite the import permit's economic cost, the political price of an open border policy for corn is controversial because of corn's central role as a subsistence crop for most of Mexico's producers. Mexico has the world's second highest per capita consumption of corn with 68% of all corn used for food, mostly in the form of tortillas (Morris 1998). Therefore, ending government intervention in the corn market could have a negative effect on consumers, since according to Garcia (1996) tortillas and corn flour may be subsidized by as much as 40%.

[17] The Mexican government identifies the industries that will need to import corn and ensures that they buy domestic corn. In order for these firms to be assured an import permit they must participate in the domestic market. In other words, they must buy a certain amount of domestic corn in order to be assured of getting an import permit.

Corn Interest Groups

Since the ratification of NAFTA, corn interest groups can be placed into two general categories; those who are against corn imports (and therefore do not want import permits issued) and those in favor of corn imports (and therefore want import permits to be issued). Those against corn imports and the issuance of import permits to importers fear that allowing imports of tariff free corn could damage domestic production. Those in favor of imports want access to quality, low cost inputs and at a minimum want a consistent import permit issuance policy.

Mexican corn and sugar producers have historically lobbied against tariff free corn imports allowed in by import permits for lower priced U.S. corn as they regard it as a rival in their market (Devadoss, Kropf, and Wahl 1995; Mena 1997)[18]. These producer groups have also lobbied SECOFI for inclusion in the decision making process. They have continually protested that the Cupo (import permit) Committee, which was made up of SECOFI officials, industry representatives, and the National Company of Popular Subsistence (CONASUPO), has always consulted with corn importers but left producers out of the process.

Sugar producers and processors claim that NAFTA negotiations worked against them since the U.S. restricts sugar imports through a quota that is tied to Mexican consumption of sugar and high fructose corn syrup (HFCS). As Mexican consumption of sugar and HFCS increases so does the U.S. import quota for Mexican sugar. This means that not only is their export market restricted but because HFCS is a substitute for sugar, tariff free corn allows HFCS manufacturers to reduce costs and increase their market share in the Mexican sweetener market.

In contrast, large corn processors lobby for the issuance of corn import permits. The industries that make up this group include the livestock industry, the tortilla industry, the corn flour industry, the starch industry and HFCS manufacturers. They claim that in order to compete in the market they must be allowed to import quality low-cost inputs and that includes corn. In its effort to appease producer groups and placate large corn processors, the Mexican government has tried to institute a policy of not issuing import permits during harvest in an attempt to time imports in a manner that avoids displacing domestic production but still fulfills its NAFTA obligations. In effect, the Mexican government is trying to protect two different groups made up of individual firms lobbying for their own interests.

Markets for Trade Protection

For illustrative purposes consider isolating a simple conceptual model of trade protection for the *i*th commodity ($i = w$ for white corn, y for yellow corn, and s for sorghum). This can be conceptualized as a demand and supply equilibrium model in the form:

[18] Mexican producers might have cause for concern as research by Bivings (1997) showed that the liberalization of the sorghum market in Mexico hurt Mexican producers who could not compete with U.S. sorghum imports. The Mexican government did not re-instate import permits in this case but did put in place seasonal tariffs on imports, domestic price controls, and a storage subsidy scheme (Bivings 1997).

$$Q_{EPi}^{d} = f_{EPi}^{d}(P_{EPi}, Z_{EPi}^{d}) \tag{1}$$

$$Q_{EPi}^{s} = f_{EPi}^{s}(P_{EPi}, Z_{EPi}^{s}) \tag{2}$$

$$Q_{EPi}^{d} = Q_{EPi}^{s} = Q_{EPi} \tag{3}$$

Here, Q_{EPi}^{d} is the quantity demanded of protection for the *i*th commodity, Q_{EPi}^{s} is the quantity supplied of protection, P_{EPi} is the price of protection, and $Z_{EPi}^{d}(Z_{EPi}^{s})$ is a matrix of exogenous factors of demand (supply). The last equation (3) is a simple market clearing condition.

In the case discussed above, the Mexican government is the supplier of protection from corn and sorghum imports. Domestic producers of corn and sugar demand protection and lobby for import permits not to be issued, especially not during harvest periods of June, July, November, and December. Alternatively, industrial users of corn made up primarily of the starch, tortilla, livestock, corn flour, and cereal industries lobby to have import permits issued in a more timely and consistent manner. Meanwhile, the Mexican government is trying to optimize its political support by placating both groups.

From the equilibrium model of trade protection, it is possible to employ the implicit function theorem to solve the system for the endogenous variables in terms of the exogenous variables. This yields the reduced form equations $Q_{EPi} = h_q(\mathbf{Z}_p)$ and $P_{EPi} = h_p(\mathbf{Z}_p)$ where $\mathbf{Z}_p$ is the matrix of exogenous variables influencing trade protection. As a result we can hypothesize that equilibrium quantity and price of protection are a function of exogenous factors such as macroeconomic, political, and production variables. An example of a political variable that can influence the level of protection is producer pressure on government officials to halt the issuance of import permits.

Extending the concept of trade protection to that of endogenous protection, hypothesizes that markets of imports are simultaneously determined with markets for trade protection (see Magee, Brock, and Young 1989; Trefler 1993; Gawande and Bandyopadhyay 2000). As pressure from producer associations in Mexico increases, the number of import permits issued decreases thereby increasing the level of protection. Under this assumption, a completely generalized simultaneous equation model of import demands can be defined as:

$$Q_i^d = g_i(\mathbf{P}, P_{EPi}, Q_{EPi}) \tag{4}$$

where the quantity demanded of the *i*th import commodity is a function of a vector of import prices **P** and endogenous protection variables P_{EPi}, Q_{EPi}. Similarly, each demand and supply function of protection for the other commodities could be re-specified to be simultaneously determined with quantity demanded and prices of imports.

Linking Import Demand to Trade Protection

Because the quantity and price of trade protection are not observable for Mexican corn and sorghum, reduced form equations are useful in linking import demand equations to political market variables. Zellner (1970) and Goldberger (1972) both discuss simultaneous equation models where potentially endogenous dependent variables are not observable. Zellner points out, that functional relationships containing unobservable variables can be compensated by using an instrumental variable approach. As a result, substituting the reduced form relationships from the trade protection market into the simultaneous import demand equations in equation (4) yields:

$$Q_i^d = g_i^*(\mathbf{P}, \mathbf{Z}_p) \tag{5}$$

The specification in equation (5) indicates that import quantities are functions of prices, as well as exogenous variables influencing trade protection.

DATA

Data were collected monthly from January 1994 to December 2000 using various sources, including the United States Department of Agriculture's (USDA) Federal Grain Inspection Service (FGIS) and the Food and Agricultural Organization (FAO) of the United Nations. Table 1 contains the definitions of data variables and descriptive statistics.

The quantity of white corn, yellow corn, and sorghum Mexico imported from the U.S. per month from 1994 to 2000 were obtained from FGIS. Total U.S. monthly exports to Mexico, which include agricultural and nonagricultural goods, were obtained from the U.S. Department of Commerce. Total exports from the U.S. to Mexico may capture macroeconomic effects on the trade flow between the two countries such as the import penetration of U.S. products into Mexico and currency devaluations.

Articles in the media concerning the use and allocation of import permits were collected from Lexis-Nexis and the USDA, Foreign Agricultural Service (FAS) for each month between January 1994 and December 2000. News items are used as a proxy for producer pressure on the Mexican government against import permits similar to the process used by Davis, Leeds, and Moore (1998).

Import permit allocations were obtained from FAS from 1994 to 2000. Import permits are categorized by the Mexican government as either for human consumption (white corn) or industrial use (yellow corn). Since each industry is issued import permits on a quarterly basis, if an import permit was issued to an industry for a certain quarter a 1 was placed for all three months in that quarter and 0 otherwise. Although individual firms may be given permits with different amounts from quarter to quarter, the corn volume allotted to each industry fluctuates little, from quarter to quarter. Import permits were allocated to the starch industry, the livestock industry, the cereal industry, the corn flour industry, the tortilla industry, and CONASUPO. Table 2 depicts the type of import permits issued to each industry.

Table 1. Descriptive Statistics and Variable Definitions

Variable	Definition	Mean	Std Dev	Observations
YCORN	Quantity of U.S. yellow corn inspected for export to Mexico in metric tons	264894.092	167057.856	84
WCORN	Quantity of U.S. white corn inspected for export to Mexico in metric tons	47217.0992	51838.9869	84
SORG	Quantity of U.S. sorghum inspected for export to Mexico in metric tons	169989.965	102399.647	84
TTLEXP	Total exports from U.S. to Mexico in million metric tons.	5980.3119	1861.9859	84
MXMZ	Total production of corn in Mexico	1516094.81	1245618.05	84
MXSG	Total production of sorghum in Mexico	457494.823	88947.4255	84
NEWS	Articles on producer protests	1.5833	2.0780	84
STCH	Starch industry import permit allocations (0=no, 1=yes)	0.7500	0.4356	84
LIVE	Livestock industry import permit allocations (0=no, 1=yes)	0.5952	0.4938	84
CEREAL	Cereal industry import permit allocations (0=no, 1=yes)	0.5119	0.5029	84
FLOUR	Corn flour industry import permit allocations (0=no, 1=yes)	0.6548	0.4783	84
SUPO	CONASUPO import permit allocations (0=no, 1=yes)	0.3571	0.4820	84
TORT	Tortilla industry import permit allocations (0=no, 1=yes)	0.4405	0.4994	84
CPYC	Price of yellow corn USD per metric ton	95.9781	24.3395	84
CPSG	Price of sorghum USD per metric ton	85.0348	23.7567	84
CPWC	Price of white corn USD per metric ton	116.5033	29.9183	84

Monthly cash prices in U.S. dollars for yellow corn and sorghum were obtained from National Agricultural Statistics Service (NASS), USDA. The cash price for white corn was obtained from the Global Risk Management Corporation.[19] The nominal daily exchange rate between the U.S. dollar and the Mexican peso was obtained from the Bank of Canada and averaged for each month.

Table 2. Import Permit Allocation by Industry

Industry	Type of Import Permit	Corn Type
Cereal Industry	Human Consumption	White
CONASUPO	Human Consumption Industrial Use	White Yellow
Corn Flour Industry	Human Consumption	White
Livestock Industry	Industrial Use	Yellow
Starch Industry	Industrial Use	Yellow
Tortilla Industry	Human Consumption	White

EMPIRICAL ANALYSIS

Import demand equations for Mexico for yellow corn, white corn, and sorghum are conceptually derived following Diewart and Morrison's (1986) production theory approach from a restricted quadratic profit function. The import demand models are simultaneously determined with the demand for protection using an instrumental variables approach. As discussed ahead the import demand equations are econometrically estimated equation by equation using a simultaneous tobit framework with monthly data from January 1994 to December 2000.

Consider, for example, the import demand model for white corn (Q_w^d), which is a function of the quantity of yellow corn imported (Q_y^d), quantity of sorghum imported (Q_s^d), quantity of import protection (Q_{EPw}), exogenous variables associated with the import demand of white corn (Z_w), price of white corn $\left(P_w\right)$, price of yellow corn $\left(P_y\right)$, price of sorghum $\left(P_s\right)$, and an error term e_w:

$$Q_w^d = \beta_1 Q_y^d + \beta_2 Q_s^d + \beta_3 Q_{EPw} + \mathbf{Z}_w \beta_4 + \beta_5 P_w + \beta_6 P_y + \beta_7 P_s + e_w \tag{6}$$

[19] However, the GMRC data for white corn was not complete and only covered the period from September 1997 to April 2000. This three year data set was in daily terms so the average for each month was used. In order to have a complete data set, yellow corn cash price was used to estimate white corn cash price using Ordinary Least Squares (OLS) and the estimated values for white corn were then used for the missing values.

The reduced form function for import protection of white corn is represented as:

$$Q_{EPw} = \mathbf{Z}_p \alpha_{EPw} + v_{EPw} \tag{7}$$

with error term v_{EPw}. Because the quantity of import protection for white corn is not observable, the reduced form relationship in equation (7) is substituted into equation (6) yielding

$$Q_w^d = \beta_1 Q_y^d + \beta_2 Q_s^d + \mathbf{Z}_p \left(\beta_3 \alpha_{EPw}\right) + \mathbf{Z}_w \beta_4 + \beta_5 P_w + \beta_6 P_y + \beta_7 P_s + u_w \tag{8}$$

where $u_w = \beta_3 v_{EPw} + e_w$. In effect this is analogous to Theil's (1971) approach to achieving consistent estimates for two stage least squares, which is a reduced form representation of the structural form model in equation (6). A limitation of this process of dealing with unobservable independent variables is the inability to identify parameter α_{EPw} from β_3 in the estimation process, since the quantity of import protection is not known. However, the model still yields a consistent estimate of $(\beta_3 \alpha_{EPw})$, which is particularly useful for prediction and simulation purposes.

Due to the use of import permits (and other potential trade barriers), the dependent variable in the import demand model is censored at zero. As a result, the import demand equation for white corn, for example, is specified as:

$$\begin{cases} Q_w^d = Q_w^{d*} & \text{if } Q_w^{d*} > 0 \\ Q_w^d = 0 & \text{otherwise} \end{cases} \tag{9}$$

where Q_w^{d*} is the latent amount demanded. Given censoring of the dependent variable, performing OLS on equation (8) would result in inconsistent coefficient estimates. To accommodate censoring and simultaneity problems discussed above, the method we apply is an equation by equation simultaneous tobit estimator.[20]

The matrix of import protection variables $\mathbf{Z}_p$ include producer pressure, Mexican government allocations of import permits to the corn flour, tortilla, starch, livestock, and cereal industries as well as to CONASUPO. Lagged variables for import protection are also added as import permits last month may have an effect on import demand for this month.

Hypothesized Effects

The public choice literature suggests different industries will lobby for or against import protection depending on whether they are consumers or producers of a product (Grossman and Helpman 1995; Levy 1997; Maggi and Rodriguez-Clare 1998). Grossman and Helpman

[20] In order to determine if the quantities of white corn, yellow corn, and sorghum are jointly determined, a test for exogeneity in a simultaneous equation tobit model is performed using the approach by Smith and Blundell (1986). We failed to reject the null hypothesis of weak exogeneity in all three models. This suggests that monthly white corn, yellow corn, and sorghum exports are not jointly determined.

(1994) and Gawande and Bandyopadhyay (2000) discuss how different industries form a lobby to interact with the government in order to maximize protection. Each industry lobbies to increase or decrease the level of import protection. The hypothesized results of these lobbying efforts are discussed below.

The import demand equations for white corn, yellow corn, and sorghum are a function of prices, production variables, and protection variables.[21] From economic theory one can state a priori the expected effects of price and production variables on the quantity demanded of corn and sorghum. For example, own-price effects should be negative. Cross-price effects will reflect the degree of substitutability or complementarity between imports of corn and sorghum, which are themselves interesting because they may influence the success of strategic trade outcomes.[22] In general, we expect that as the supply of protection increases, then import demand will decrease. Likewise, as the demand for protection decreases, then import demand will increase.

In a multiple market setting with permit allocation restrictions, protection in one market has the potential to directly influence that commodity or indirectly create spillovers onto other commodities. Spillovers can be created by induced substitution or complement effects across commodities. Moreover, with possible restrictions on the allocation of import permits across industry groups, spillovers may arise due to tradeoffs between yellow and white corn permit allocations.

Consider, for example, lobbying efforts by the tortilla industry. They should have a positive impact on the demand for white corn. In other words, if the tortilla industry increases its lobbying efforts for more import permits issued, then the Mexican government will likely decrease its supply of protection and import demand should increase. In contrast, if the starch industry increases its lobbying efforts for yellow corn one would expect that the import demand for white corn will decrease. Also, if the cereal industry receives an import permit, one would expect that the import demand for white corn will increase. If producer pressure increases for protection, then import demand for white corn will decrease.

Alternatively, consider yellow corn. If the yellow corn importing industries (i.e. starch and livestock industries) receive import permits, then it is expected that import demand for yellow corn will increase. However, if the tortilla, cereal, corn flour, and CONASUPO receive import permits then import demand for yellow corn should decrease as they will want import permits for white corn. Import demand will also decrease if producer pressure increases.

Finally, consider sorghum. In the sorghum market, if the demand for protection from imports of yellow corn increases then the import demand for sorghum increases. For instance, if sugar producers increase their lobbying efforts so that the Mexican government decreases the amount of import permits given for yellow corn, then those industries that can, will substitute sorghum for yellow corn (such as the livestock industry increasing their imports of sorghum). Alternatively, one would expect that if the livestock industry receives import permits for yellow corn, then the import demand for sorghum will decrease. When the livestock industry receives import permits for corn they will not import sorghum unless the price of sorghum is low enough to make up for the difference in protein.

[21] Although more applicable to price uncertainty, including the GSM102 and 103 program as a variable may be of interest for future research.

[22] For instance, Dong, Marsh, and Stiegert (2006) demonstrate that product differentiation in the global malt barley market dampened the ability/desire of state trading enterprises to pursue rent shifting objectives.

Results

The import demand models were estimated using LIMDEP (Greene 1998). The parameter estimates, standard errors, and p-values for the white corn, yellow corn, and sorghum models are presented in tables 3, 4, and 5, respectively.

White Corn

The estimated results from the white corn model provide some interesting insights into the import demand for white corn. For example, the results indicate that increases in yellow corn or sorghum imports are associated with increases in the import demand for white corn. Import permits allocated to the starch industry negatively relate to the import demand for white corn as does the Mexican production of sorghum. White corn price is not statistically significant.

Table 3. Mexican Import Demand for White Corn with Endogenous Protection

Variable	Coefficient	Standard Error	P[|Z|>z]	Marginal Effects
YCORN (metric ton)	0.143***	0.036	0.0001	0.126
SORG (metric ton)	0.202***	0.067	0.0027	0.178
TTLEXP (metric ton)	13.999	10.769	0.1936	12.312
MXMZ (metric ton)	-0.000	0.004	0.9938	-0.000
MXSG (metric ton)	-0.203*	0.122	0.0973	-0.178
NEWS (# of articles)	-601.041	2951.516	0.8386	
FLOUR (1=yes,0=no)	17631.005	17199.420	0.3053	
TORT (1=yes,0=no)	17401.665	12210.406	0.1541	
STCH (1=yes,0=no)	-44120.994**	17341.083	0.0109	
LIVE (1=yes,0=no)	16414.738	15269.205	0.2824	
SUPO (1=yes,0=no)	19715.739	17488.923	0.2596	
CEREAL (1=yes,0=no)	6191.903	15698.363	0.6933	
CPWC (Peso/MT)	6348.010	10528.770	0.5466	5583.074
CPYC (Peso/MT)	-3031.719	15744.780	0.8473	-2666.396
CPSG (Peso/MT)	-6946.475	6822.042	0.3086	-6109.424
LNEWS(1=yes,0=no)	1423.515	2492.527	0.5679	
LSTCH (1=yes,0=no)	16338.610	15969.066	0.3062	
LLIVE (1=yes,0=no)	-12457.562	16668.675	0.4548	
LCEREAL (1=yes,0=no)	11183.041	14451.690	0.4390	
LFLOUR (1=yes,0=no)	-24785.883	20128.938	0.2182	
LSUPO (1=yes,0=no)	17564.879	16202.727	0.2783	
LTORT (1=yes,0=no)	-9253.962	11191.947	0.4083	
Constant	20454.126	61438.317	0.7392	

*** Indicates statistically significant at the 1% level
** Indicates statistically significant at the 5% level
* Indicates statistically significant at the 10% level
Log Likelihood = -808.8748

Yellow Corn

Total imports have a negative effect on the import demand for yellow corn. This would indicate that as imports from the U.S. increase there is a decrease in the import demand for yellow corn (i.e. less imports allowed). Trefler (1993) suggested that increased import penetration increases import protection and causes a decrease in import demand.

Table 4. Mexican Import Demand for Yellow Corn with Endogenous Protection

Variable	Coefficient	Standard Error	P[|Z|>z]	Marginal Effects
WCORN (metric ton)	1.3513***	0.341	0.0001	1.344
SORG (metric ton)	-0.0201	0.200	0.9165	-0.021
TTLEXP (metric ton)	-66.187**	29.511	0.0249	-35.839
MXMZ (metric ton)	-0.005	0.012	0.6842	-0.005
MXSG (metric ton)	1.020***	0.336	0.0024	1.015
NEWS (# of articles)	5074.616	8278.395	0.5399	
FLOUR (1=yes,0=no)	-28513.553	48134.476	0.6627	
TORT (1=yes,0=no)	-27912.178	35491.217	0.4337	
STCH (1=yes,0=no)	85809.544*	49670.745	0.0841	
LIVE (1=yes,0=no)	31460.933	44396.237	0.6024	
SUPO (1=yes,0=no)	76210.957	47892.429	0.1115	
CEREAL (1=yes,0=no)	-126192.041***	43130.226	0.0034	
CPWC (Peso/MT)	59314.064**	29323.940	0.0431	59002.223
CPYC (Peso/MT)	-129399.213***	42282.223	0.0022	-128718.904
CPSG (Peso/MT)	74316.667***	17430.680	0.0000	73925.952
LNEWS(1=yes,0=no)	-7856.594	6899.675	0.2548	
LSTCH (1=yes,0=no)	2873.388	44975.220	0.9491	
LLIVE (1=yes,0=no)	35365.919	48386.186	0.4648	
LCEREAL (1=yes,0=no)	-44859.687	41464.514	0.2793	
LFLOUR (1=yes,0=no)	99919.014*	53863.498	0.0636	
LSUPO (1=yes,0=no)	-79063.345*	43416.274	0.0686	
LTORT (1=yes,0=no)	-6232.428	32108.764	0.8461	
Constant	-268253.875	165100.310	0.1042	

*** Indicates statistically significant at the 1% level
** Indicates statistically significant at the 5% level
* Indicates statistically significant at the 10% level
Log Likelihood = -1077.225

White corn imports have a positive and statistically significant effect on yellow corn import demand. Increased Mexican sorghum production is associated with increases in the import demand for yellow corn. Increased sorghum production will mean less yellow corn acreage which in turn results in an increase in demand for yellow corn. White corn and sorghum prices have a positive and statistically significant effect on the import demand for yellow corn. This is intuitive, since they often are substitutes for yellow corn. Yellow corn price has a negative and significant impact on yellow corn imports.

Import permit allocations to the starch industry seem to have a positive effect on the import demand for yellow corn. This suggests when the starch industry lobbies for import permits, it increases the import demand for yellow corn. However, the cereal industry receiving import permits is negatively correlated with import demand for yellow corn. The cereal industry would apply for a white corn import permit, and if they were successful, it would cause the Tariff Commission to issue less import permits for yellow corn.

The lagged variables for permit allocations to the corn flour industry and CONASUPO (the government agency) were also statistically significant. Import permit allocation to the corn flour industry in the previous period has a positive effect on the import demand for yellow corn. In contrast, import permit allocation in the previous period to CONASUPO has a negative effect on import demand in the current period. A plausible explanation is that the Tariff Commission is staggering allocations for white corn and yellow corn. In other words, when they issue white corn import permits for an industry in one period, they are less likely to issue import permits to the same industry for white corn in the next period.

Table 5. Mexican Import Demand for Sorghum with Endogenous Protection

Variable	Coefficient	Standard Error	P[|Z|>z]	Marginal Effects
WCORN (metric ton)	0.548***	0.198	0.0056	0.547
YCORN (metric ton)	-0.007	0.061	0.9053	-0.007
TTLEXP (metric ton)	6.631	16.886	0.6946	6.619
MXMZ (metric ton)	-0.006	0.007	0.3506	-0.006
MXSG (metric ton)	-0.102	0.197	0.6036	-0.102
NEWS (# of articles)	-5607.844	4583.230	0.2211	
FLOUR (1=yes,0=no)	4145.666	26819.035	0.8772	
TORT (1=yes,0=no)	-6085.181	19840.324	0.7591	
STCH (1=yes,0=no)	-12127.955	28117.271	0.6662	
LIVE (1=yes,0=no)	-8509.352	24746.856	0.7310	
SUPO (1=yes,0=no)	-23041.569	26981.482	0.3931	
CEREAL (1=yes,0=no)	27937.803	25017.267	0.2641	
CPWC (Peso/MT)	-17776.379	16588.474	0.2839	-17744.485
CPYC (Peso/MT)	51877.440**	24138.856	0.0316	51784.362
CPSG (Peso/MT)	-26667.981***	10277.994	0.0095	-26620.134
LNEWS (# of articles)	-277.601	3869.568	0.9428	
LSTCH (1=yes,0=no)	-7445.326	25014.264	0.7660	
LLIVE (1=yes,0=no)	-47349.023*	26489.830	0.0739	
LCEREAL (1=yes,0=no)	8676.628	23201.398	0.7084	
LFLOUR (1=yes,0=no)	36305.043	30295.749	0.2308	
LSUPO (1=yes,0=no)	-7228.303	24622.505	0.7691	
LTORT (1=yes,0=no)	17298.580	17758.355	0.3300	
Constant	-31408.510	93842.842	0.7379	

*** Indicates statistically significant at the 1% level
**Indicates statistically significant at 5% level
* Indicates statistically significant at the 10% level
Log Likelihood = -1017.397

Sorghum

Table 5 shows the positive association of white corn imports with sorghum demand. Therefore, as white corn imports increase so does the demand for sorghum imports. Yellow corn prices are also positive and statistically significant. Again, yellow corn and sorghum are substitutes. Sorghum price has a negative and significant impact on sorghum imports. Import permits issued to the livestock industry in the previous period also has a negative effect on the import demand for sorghum. This could mean that when the livestock industry is issued corn import permits in one period, it will not import as much sorghum in the next period.

Discussion

The results from the three import demand models predominately are consistent with the hypothesized effects. When a large industry with political influence such as the tortilla industry lobbies the Mexican government, import demand tends to increase for the particular commodity it has a vested interest in. We hypothesized that if the tortilla industry increased its lobby, it would be associated with an increase in imports of white corn and a decrease in the imports of yellow corn (and vice versa for the starch industry). We can also see from tables 3 and 4 that the starch industry may be more political active than the tortilla industry. This is because when the starch industry lobbies it has a negative and statistically significant effect on white corn imports, but the tortilla industry was not statistically significant (while having a negative effect on yellow corn imports).

We see a positive association between corn and sorghum when there is a large quantity demanded via consumers or industry. In tables 3-5, white corn has a positive association with both yellow corn and sorghum. Of the three commodities, white corn is the most protected in Mexico. We believe that when import protection is lowered for white corn then yellow corn and sorghum also have an increased probability of being imported. Hence, an increased amount of yellow corn and sorghum are imported when white corn imports increase.

PRICE SIMULATIONS AND MARKETING MARGINS

The uncertain nature of the allocation process for corn import permits influences expected import prices and, hence, expected marketing margins. The marketing margin is the difference between the import price and the export price excluding transportation costs. The marketing margin captures transactions costs such as the price of documentation, phytosanitary restrictions, and/or other technical trade barriers that may be present during the transaction period. We conduct import price simulations for white corn, yellow corn and sorghum to quantify expected marketing margins.

Gallagher (1998) presents a stylized model of exporters that motivates the impact of trade uncertainty on expected marketing margins. Under risk neutrality, the expected marketing margin can be expressed as:

$$P^e = \frac{P_f + c}{\rho} + \frac{q^d (1-\rho)}{\beta_r} \tag{10}$$

where expected import price (P^e) is a function of a parameter value from the import demand model (β_r), export price (P_f), transportation costs (c), quantity imported per transaction (q^d), and the probability of a successful transaction (ρ).[23] The difference between the expected import price and export price consists of two terms. The first term is the sum of the export price plus transportation costs divided by ρ. The second term represents a wedge driven into the expected marketing margin due to trade uncertainty. For instance, if trade is certain ($\rho = 1$), then the expected import price would just equal the export price plus transportation costs. If trade is uncertain ($0 < \rho < 1$), then the expected marketing margin would be nonzero and positive.

Expected import prices are simulated for white corn, yellow corn and sorghum imports using the empirical models reported above. Tobit models provide the probability of a successful transaction per month from the observed data (Greene 1998, 2003). On average over the study period the estimated probabilities were 0.76, 0.95 and 0.96 for white corn, yellow corn and sorghum, exhibiting magnitudes consistent with the observed data. Given the uncertainty of import permits to Mexico for corn and the results of (10), we anticipated and observed that annual marketing margins were positive.

For illustrative purposes the resulting marketing margins for yellow corn and sorghum are shown in figure 1.[24] The difference between the expected import prices and actual export prices are the expected marketing margins associated with the import permit. The expected import prices for yellow corn and sorghum fell between the actual import and export prices. This result is expected as it captures the impact of the uncertain import permit on the marketing margin. The expected marketing margin for the period between 1994 and 2000 stays below $6.50 except in 1995 and 1996 where the expected marketing margins are $9.98 and $9.65, respectively. This increase in the marketing margin coincides with the decreased probability of a successful transaction in 1995 and the shock to the corn market in 1996 because of a lack of supply and an increase in world demand.

The sorghum market has its highest predicted marketing margin of $29.91 in 1996. The lowest expected margin is $10.38 in 2000. Interestingly, although sorghum has no import permit process, the price simulations suggest that there is a higher expected marketing margin relative to yellow corn. This is consistent with the actual marketing margin for sorghum, which is higher than the actual marketing margin for yellow corn (based on FAO prices).

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[23] The import demand equation is specified as $Q_r = \alpha_r - \beta_r P_r$. Here ρ is the probability of a successful import transaction (i.e., an import permit has been issued), which is derived from a Bernolli random variable. See Gallagher (1998) for further details.

[24] Initially we also simulated marketing margins for white corn, which exhibited qualitative results consistent with yellow corn and sorghum results. However, given the white corn price variable in the white corn import model was insignificant, this yield magnitudes of the price simulations and marketing margins that were much less reliable. Hence, while white corn marketing margin were positive as expected, they are not reported.

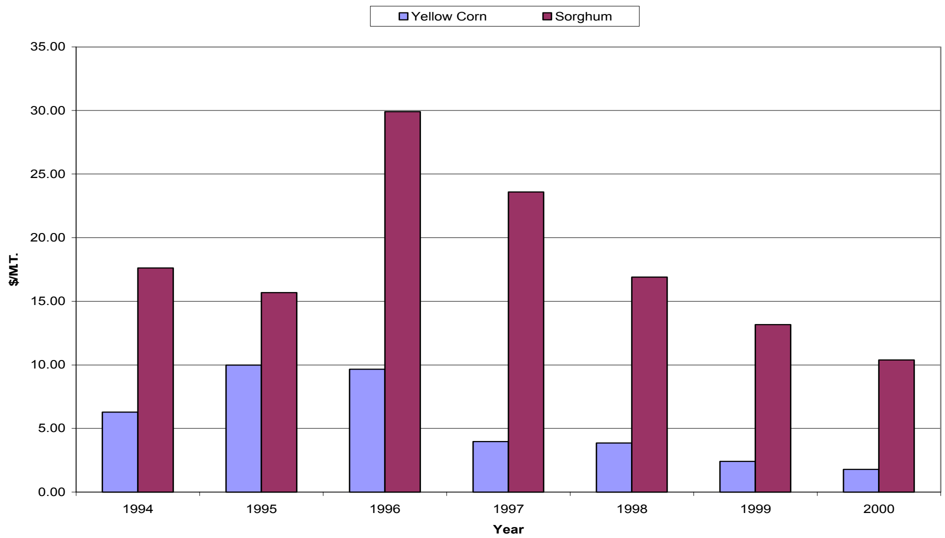

Figure 1. Expected marketing margins for yellow corn and sorghum from 1994 to 2000

Given that sorghum is imported when corn import permits are not available or when the corn premium is higher than the protein differential merits, then a plausible explanation is that the sorghum model may also be capturing indirect impacts of trade uncertainty in the corn market.[25] Results presented in table 5 indicate a significant and positive relationship between the price of yellow corn and the amount of sorghum imported, supporting that actual exports of yellow corn and sorghum into Mexico are substitutes. As a result, there is potential for an implicit administrative barrier in the sorghum market due to a spillover impact from corn import permits. This is consistent with Dong, Marsh, and Stiegert (2006), whom demonstrate that product substitution can influence the strategic nature of international trade.

CONCLUSIONS

A conceptual structural model of international marketing margins and trade uncertainty that links the private market to political factors influencing administrative trade barriers was specified. The intent was to provide a trade model with endogenous protection that can help explain uncertainty in the trade of agricultural commodities due to technical trade barriers. This provided a framework to identify factors that determine the likelihood of issuing import permits for corn exported from the U.S. to Mexico and to determine its impact on international import demand and marketing margins. Mexican import demand models are estimated simultaneously with the demand for protection using an instrumental variables approach. In this study import demand equations are estimated using a simultaneous tobit model with monthly data from January 1994 to December 2000.

Some interesting insights are provided about trade uncertainty, political markets, and product substitution. We find empirical support for the hypothesis that trade uncertainty

[25] Of course, alternative interpretations could include other administrative trade barriers, potential model misspecifications, or combinations of each.

increases expected import prices. Also our results are consistent with public choice theory, which indicates that increased import penetration will increase lobbying efforts from domestic firms to decrease imports. When lobbying efforts of corn and sugar producers increase, the import quantity of corn and sorghum decreases, meaning that they have succeeded in their efforts to reduce import competition. Marketing margins associated with the corn import permit range form an average of \$6 for yellow corn and \$18 for sorghum. An interesting insight is that although the import permit restricts corn imports, sorghum import prices are also affected and the uncertainty caused by spillover effects from the technical trade barrier.

A conclusion drawn from the results is that trade uncertainty caused higher prices in the import market. Without the import permits, importers would be able to contract for corn in the U.S. on a long-term basis and prices in Mexico would be lower. Mexican importers would be able to take advantage of low spot prices. The removal of duty-free quotas in Mexico would increase U.S. exports to Mexico. This leads us to believe that the license permit regime in Mexico is having a detrimental affect on U.S. exports of corn to Mexico. If the import permit regime ended, the U.S. should expect a higher demand for corn from Mexico since prices would no long be kept artificially high by the import quota. Mexican importers would also come into the market throughout the year and not only during the "open" season that their government allows. Producers in the U.S. would be better able to predict not only when Mexican importers would come to the market but also demand for corn.

In all, this suggests that trade uncertainty caused by administrative or technical trade barriers alters the trade flow between countries. Our results provide evidence of higher prices for the importing country and lower quantities for the exporting country. It is apparent that this research has practical implications. One could forecast the effect a possible change in policy, be it political or macroeconomic, might have on import prices. Underlying the model is the fundamental idea that one is linking political and private market processes to better understand the implications of administrative trade barriers.

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EXPLOITATION AND ANALYSIS OF LONG RUN COINTEGRATION PROPERTIES OF A QUARTERLY SYSTEM OF U.S. WHEAT-RELATED PRODUCT MARKETS

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ABSTRACT

We apply methods of the cointegrated vector autoregression (VAR) model to quarterly U.S. markets for wheat and for the wheat-using products of wheat flour, mixes and doughs, bread, wheat-based breakfast cereals, and cookies and crackers. This study extends recent reduced-form VAR econometric work done on the same markets, by dichotomizing the system into a long run error-correction space of economic relationships and a short run component, with some structural relationships having emerged from the long run error correction space. Results include an array of empirical estimates of the parameters (some structural) and relationships that drive and with which the wheat-related markets interact. An array of empirical estimates of market impacts on policy, institutional, and trade events is also provided.

Key words: *Cointegration, wheat-based U.S. markets, vector autoregression and error correction models.*

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INTRODUCTION: ANALYTICAL PURPOSE AND PRIOR RESEARCH

Our primary purpose is to provide a procedure of specific cointegrated VAR methods inspired by Johansen (1988) and Johansen and Juselius (1990, 1992), and recently refined by Juselius (2004) and Juselius and Toro (2005). Such a procedure combines well-known and widely applied cointegration methods, along with some refinements and advancements introduced below that are relatively new and that have not yet appeared in the agricultural economics literature. Our second purpose is, for perhaps the first time, to use this combined set of established and newly refined cointegration procedures to extend some recent quarterly econometric research on U.S. wheat-based markets by Rich, Babula, and Romain (2002) and Babula, Bessler, and Payne (2004).

RBR (2002) applied methods of vector autoregression or VAR modeling to a quarterly system of six U.S. markets for wheat, wheat flour, mixes and doughs (mixes/doughs), bread, wheat-based breakfast cereals (cereals), and cookies and crackers (cookies/crackers). They then provided detailed analyses of impulse response simulations and of other econometric procedures and estimates to illuminate the dynamic quarterly patterns with which this reduced form system interacts and some empirical estimates of principal market parameters which propel these markets. Babula, Bessler, and Payne (hereafter, BBP 2004) updated RBR's (2002) sample and methods. BBP (2002) extended RBR's findings by having combined Bernanke (1986) structural VAR methods with directed acyclic graph (DAG) analysis, to the same set of six quarterly markets. One of BBP's contributions was a more evidentially sound treatment of contemporaneously causal orderings than inherent with the more traditionally used Choleski-type orderings used in Sims (1980) and Bessler (1984a, b).

It is well-known that Johansen (1988) and Johansen and Juselius (1990, 1992) extended the non-structural reduced form levels-VAR methods that Sims (1980) developed, that Bessler (1984 a, b) and Chambers (1984) applied early-on to U.S. agriculture, and that RBR (2002) and BBP (2004) applied to the above six U.S. wheat-related markets. Such extensions are detailed below and involved dichotomizing a levels VAR model into short run and long run components, and then exploiting long run component's cointegration properties so as to permit the emergence of identified and structurally interpreted economic relationships from the once exclusively reduced form framework of Sims (1980) and Bessler (1984a, b). Our results identify a set of structural, along with some reduced form, insights that enhance the precision of the findings of the RBR (2002) and BBP (2004) work. We include estimates of an own-price elasticity of supply, cross-price transmission or response parameters, and a rich set of empirically estimated effects of policy, trade, economic, and political events (hereafter, important market/institutional events).

This paper has seven ensuing sections. The first presents cointegrated VAR methods as a way to empirically model the quarterly U.S. system of wheat-related markets. The second section provides a summary of specification implications of the modeled data's stationarity/nonstationarity properties. It is well-known that such utilization of information inherent in nonstationarity properties is required for non-compromised inference and to avoid spurious regressions (Granger and Newbold 1986, pp. 1-5; Hendry 1986). The third section summarizes efforts to achieve an adequately specified levels VAR model (and its unrestricted VEC equivalent) to exploit what are later revealed to be the system's substantial cointegration properties. We provide a rigorous analysis of the model's statistical adequacy based on results from a battery of diagnostic mis-specification tests suggested by Juselius (2004, pp. 72-82) and Juselius and Toro (2005). In a fourth section, evidence from Johansen and Juselius' (1990, 1992) well-known trace tests and from other sources is used to determine the number of cointegrating vectors or relationships (hereafter, CVs). The cointegration space is then restricted for reduced rank. The fifth section employs Johansen and Juselius' (1990, 1992) hypothesis test procedures on the rank-restricted cointegration space to illuminate the long run economic relationships that drive and tie-together the upstream and downstream wheat-based markets. This stage injects economic and statistical theory, along with market expertise, into the analysis so as to ultimately restrict the error correction space into structurally interpreted economic relationships. A sixth section provides economic interpretations of the CVs that are fully restricted for rank and for statistically supported restrictions from the hypothesis tests. A summary and conclusions follow.

Time Series Econometrics, Modeled Markets, and Data Resources

It is well-known that economic time series often fail to meet the conditions of weak stationarity (i.e., stationarity and ergodicity) required of valid inference, and in some cases, unbiased estimates, from regressions using time-ordered data (Engle and Granger 1987; Granger and Newbold 1986, pp. 1-5). And while data series are often individually nonstationary, they can form vectors with stationary linear combinations, such that the vector of inter-related series are "cointegrated" and move in tandem as an error-correcting system (Johansen and Juselius 1990, 1992).

We updated BBP's (2004) quarterly sample of the following endogenous variables (denoted throughout interchangeably by the parenthetical labels) defined and sourced as follows:

1. U.S. price of wheat (PWHEAT): reflected by the U.S. all-wheat price published by the U.S. Department of Agriculture, Economic Research Service (USDA, ERS 2004, 2005).
2. U.S. market-clearing quantity of wheat (QWHEAT): defined as the sum of beginning stocks, production, and imports, published by the USDA, ERS (2004, 2005).
3. U.S. wholesale price of wheat flour (PFLOUR): represented by the U.S. producer price index (hereafter, PPI) for wheat flour, series no. PCU3112113112111, published by the U.S. Department of Labor, Bureau of Labor Statistics (Labor, BLS 2005).
4. U.S. wholesale price of mixes and doughs (PMIXES): reflected by the U.S. PPI for flour mixes and refrigerated and frozen doughs and batters, series no. PCU3118223118226, published by Labor, BLS(2005).
5. U.S. wholesale bread price (PBREAD): defined as the U.S. PPI for bread, series PCU3118123118121, published by Labor, BLS (2005).

6. U.S. wholesale price of wheat-based breakfast cereals (PCEREAL): represented by the U.S. PPI for wheat flakes and other wheat breakfast foods, series no. PCU311230311230112, published by Labor, BLS (2005).
7. U.S. wholesale price of cookies and crackers (PCOOKIES): reflected as the U.S. PPI for cookie and cracker manufactured products, series PCU311821311821of Labor, BLS (2005).

Data are quarterly, seasonally unadjusted, and were placed into natural logarithms. Data were available for the 1985/86:01–2004/05:04 sample period.[26] Since data were not available for all variables prior to 1985/86, our analysis may have potential problems from small samples.

Analysis of Wheat-Based Time Series Data

Following Juselius (2004, chapters 3-4), we carefully examined the modeled data's logged levels and differences to discern stationarity (and nonstationarity) properties. In seminal work, Granger and Newbold (1986, pp. 1-5) and Hendry (1986) established that failure to utilize such inherent information can lead to compromised inference and spurious regressions. Examining the data's nonstationarity properties provides levels-VAR specification implications that utilize such nonstationarity-based information needed to achieve a statistically adequate model with which cointegration can be exploited.

A weakly stationary series has a constant and finite mean and variance, has time-independent observations, and generates regression coefficient estimates that are time-invariant (Juselius 2004, chapters 3 and 4). Weakly stationary data frequently cycle and mean-revert. Due to page length considerations, we do not include the plotted levels and differences of the seven wheat-based variables, and refer readers to Babula, Rogowsky, and Romain's (2006) plots of the same data. The following are resulting highlights of the specification implications needed to capture information inherent in the nonstationary elements of the modeled data:

- PWHEAT and PFLOUR follow time-enduring cycles; seldom mean-revert; display periods of slope changes possibly from policy and market events discussed later; trend through substantial subsamples; and display apparently non-constant variation levels (heteroscedasticity or ARCH effects). Differences suggest possible extraordinary effects of observation-specific events (hereafter "outlier" effects and observations). Model specification considerations include: various demand shift dummy (i.e. binary) variables, outlier variables, and a linear trend.
- QWHEAT: Wheat quantity is clearly saddled with seasonal effects and subperiods of trending. Specification considerations include centered seasonal variables and a linear trend.

[26] The U.S. wheat market year or MY extends from June 1 through May 31 of the ensuing calendar year. Hence, 1985/86:01 represents the first MY quarter extending from June through August, 1985; 1985/86:02 represents the second MY quarter extending from September through November of 1985; 1985/86:03 reflects the third MY quarter extending from December 1985 through February of 1986; and 1985/86:04 reflects the fourth MY quarter extending from March through May of 1986.

- PMIXES: The price follows a clear upward trend; displays no cycling or mean-reverting behavior; and displays marked changes in slope, particularly during 1989/90-91/92 and after 2000/01, from market/institutional events discussed below. Differences suggest ARCH effects with more volatile behavior in early subsamples, and periodic instances of outlier effects during 1986/87– 1990/91, and towards the sample's end. Specification considerations include permanent shift and outlier binary variables, and a linear trend.
- PBREAD and PCOOKIES clearly trend; do not cycle or mean-revert; and display a number of slope changes (e.g. 1996/97 for PBREAD, late-2002/03 for PCOOKIES). Differences of both variables exhibit potentially extraordinary outlier and ARCH effects throughout, particularly during the early subsamples. Specification considerations include a number of permanent shift and outlier binary variables, along with a linear trend for both prices.
- PCEREAL: The index follows different trend patterns during several subsamples; displays little or no cycling or mean-reversion; and appears to experience changes in behavior and slope from possibly market and policy events during 1995/96–1996/97, just as the 1996 U.S. farm bill and Uruguay Round were implemented. Differences reflect possible non-constant levels of variation, especially in early 1993/94 and early 1996/97. Specification considerations include a number of permanent shift and outlier binaries, and a linear trend.

THE STATISTICAL MODEL: THE UNRESTRICTED LEVELS VAR AND VEC EQUIVALENT [27]

To avoid confusing several different but overlapping definitions in the literature, we choose a number of definitions here for use throughout: (1) the *unrestricted levels VAR* denotes a VAR model in logged levels; (2) the *unrestricted VEC* denotes the algebraic equivalent of the unrestricted levels VAR in error correction form, before the cointegration space is restricted for rank or for statistically supported hypothesis test results; (3) the *cointegrated VEC* is the unrestricted VEC where the cointegration space has been restricted for reduced rank; and (4) the *fully restricted cointegrated VEC* or Juselius' (2004*) cointegrated VAR* is the unrestricted VEC after the cointegration space's restriction for reduced rank and for the statistically supported restrictions from the hypothesis tests. The "p" denotes the number (seven) of endogenous variables; "p1" denotes the number of variables in the cointegration space (seven endogenous, and other deterministic and trend variables introduced later); and "r" represents the cointegration space's reduced rank and the number of cointegrating vectors or CVs. Chosen methods specify and estimate an adequately specified levels VAR (and unrestricted VEC), where residual behavior approximates well-known assumptions of multivariate normality (Juselius 2004, chapter 5).

[27]This section draws heavily on the work of Johansen and Juselius (1990, 1992) and Juselius (2004).

THE LEVELS VAR AND UNRESTRICTED VEC OF THE WHEAT-BASED MODEL SYSTEM

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 remaining endogenous variables in the system. The above wheat-related variables render the following unrestricted, seven-equation VAR model in logged levels:

X(t) = a(1,2)*PWHEAT(t-1) +. . . . . + a(1,k)*PWHEAT(t-k)+
a(2,1)*QWHEAT(t-1)+ . . . . + a(2,k)*QWHEAT(t-k)+
a(3,1)*PFLOUR(t-1) +. . . . . +a(3,k)*PFLOUR(t-k)+
a(4,1)*PMIXES(t-1)+ .. . . +a(4,k)*PMIXES(t-k)+
a(5,1)*PBREAD(t-1)+. . . . . +a(5,k)*PBREAD(t-k) +
a(6,1)*PCEREAL(t-1)+ .. . . +a(6,k)*PCEREAL(t-k)+
a(7,1)*PCOOKIES(t-1)+. . . . . +a(7,k)*PCOOKIES(t-k)+
a(c)*CONSTANT +. . . . . +a(S)* SEASONALS + ε(t)

(1)

The asterisk denotes the multiplication operator throughout. The ε(t) is distributed as white noise. X(t) = PWHEAT(t), QWHEAT(t), PFLOUR(t), PMIXES(t), PBREAD(t), PCEREAL(t), and PCOOKIES(t). The a-coefficients are ordinary least squares (OLS) estimates with the first parenthetical digit denoting the seven endogenous variables as ordered above, and with the second referring to the lags 1, 2, ... , k. The a(c) denotes an intercept. The parenthetical terms on the endogenous variables refer to the lag: t to the current period-t, and t-k to the kth lag. Equation 1 also includes three quarterly centered seasonal variables and other potential permanent shift and outlier binaries not shown notationally. We applied Tiao and Box's (1978) lag selection procedure that uses a likelihood ratio test corrected for small samples, and results suggested a two-order lag structure (k=2).

Johansen and Juselius (1990) and Juselius (2004, p. 66) demonstrated that the VAR model in equation 1 is rewritten more compactly as an unrestricted VEC:

$$x(t) = \Gamma(1) *\Delta x(t-1) + \Pi^* x(t-1) + \Phi^* D(t) + \varepsilon(t) \quad (2)$$

The ε (t) are residuals distributed as white noise. The x(t) and x(t-1) are p by 1 vectors of the above seven wheat-based variables in current and lagged levels, Γ(1) is a p by p matrix of short run regression coefficients on the lagged differences, and Π is a p by p long run error correction term to account for endogenous variable levels. The Φ*D(t) is a set of deterministic variables: three seasonals and a host of other trend and dummy variables which will be added to address the data issues identified above as the analysis unfolds. The rank-unrestricted Π or error correction term is decomposed as follows:

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

where α is a p by r matrix of adjustment speed coefficients and β is a p by r vector of error-correction coefficients.

The $\Pi = \alpha*\beta'$ term is interchangeably denoted as the levels-based long run component, error correction term, or cointegration space of the model. The Π-term retains the levels-based information and includes long run arguments: non-differenced linear combinations of non-differenced and individually I(1) endogenous variables (under cointegration); "permanent shift" binaries with enduring effects (presented below); and a linear trend. The $[\Gamma(1)*\Delta x(t), \Phi*D(t)]$is collectively considered the short run model component and is reserved for short run arguments: the permanent shift binaries in differenced form; observation-specific outlier binaries (introduced below); and seasonal binaries.[28]

It is well-known that the unrestricted VAR model framework developed by Sims (1980), and introduced early-on to U.S. agriculture by Bessler (1984a, b) and Chambers (1984), is a reduced form one, where estimated relations reflect a mix of demand- and supply-side elements, typically without clear structural interpretations (see Hamilton 1994, chapter 11). Yet it is the very dichotomization of Sims' (1980) and Bessler's (1984) models by Johansen (1988) and Johansen and Juselius (1990, 1992) into the above long run and short run components in equations 2-3 that extended the original levels VAR methods. More specifically, one is able to sometimes identify structural error-correction relationships from what was once exclusively reduced-form approaches of Sims (1980) and Bessler (1984a, b) by separating-out the long run error-correction term from the short run component; by injecting economic theory and statistical inference through well-known Johansen-Juselius hypothesis test methods; and through reduced-rank estimation with statistically-supported restrictions from such hypothesis tests (Johansen 1988; Johansen and Juselius 1990, 1992).

Our analysis and previous research by BBP (2004), RBR (2002), and Babula and Rich (2001) suggested that we consider inclusion of a linear trend and nine permanent shift binaries discussed below in equation 2's long run levels-based cointegration space. These same variables in differenced form and a set of three centered seasonals were considered for equation 2's short run component. Analysis also led to consideration of various outlier binaries (introduced below) in the short run component. The following permanent shifters, denoted by the upper case labels, were considered for the following reasons:

- URUGUAY: valued at 1.0 for 1994/95:02–2004/05:04 MY period and 0.0 otherwise, to capture the effects of the Uruguay Round's January 1995 implementation.
- NAFTA: valued at 1.0 for the 1993/94:02–2004/05:04 MY period, and 0.0 otherwise, to capture the effects of the North American Free Trade Agreement's implementation in January 1994.
- CUSTA: valued at 1.0 for the 1988/89:02–2004/05:04 MY period, and 0.0 otherwise, to capture the effects of the January 1989 implementation of the Canadian/U.S. Free Trade Agreement.
- QUOTA: valued at 1.0 for the 1994/95:02–1995/96:02 MY period, and 0.0 otherwise, to capture the effects of the two temporary U.S. tariff rate quotas (hereafter, TRQs) placed on certain imports of Canadian durum and non-durum wheat for the year ending September 11, 1995.

[28] We used the term "short run component" here rather than the term "short run/deterministic component" used in current literature (Juselius and Toro 2005) because an anonymous reviewer questioned the longer term's appropriateness insofar as both the questioned term and long run component indeed include deterministic variables.

- FBILLS: valued at 1.0 for the 1996/97:01-2004/05:04 MY period , and 0.0 otherwise, to capture the effects of the last and current U.S. farm bills.
- TITLE7: valued at 1.0 for the 2002/03:02–2004/05:04 MY period, and zero otherwise, to account for effects of the U.S. implementation of preliminary and final antidumping and countervailing duties on certain imports of Canadian durum and/or hard red spring wheat, as a result of U.S. International Trade Commission Investigation Nos. 701-TA-430A and 430B, and 731-TA-1019A and 1019B (Final). See USITC (2003).
- DROU88: valued at 1.0 for the 1987/88:02–1989/90:04 MY period, and 0.0 otherwise, to account for the effects of the U.S. Midwest drought.
- HIDD9396: valued at 1.0 for the 1993/94:01–1996/97:01 MY period, and 0.0 otherwise, to account for the effects of the period of high levels of world grain/oilseed demands and prices.
- CONFECT: valued at 1.0 for 2001/02:01–2004/05:04 to account for sustained increases in confectionary and bakery production costs from (1) a marked 2001 increase in world cocoa prices as a result of the Ivory Coast Civil War, and (2) a steep, late-2002 incline in prices of non-cocoa confectionary inputs.[29]

The starting point for the unrestricted VEC was equation 2 with no deterministic trend or binary variables. A well-specified unrestricted VEC was ultimately achieved in a series of sequential estimations. These estimations added the seasonal variables and then a linear trend, various permanent shift binaries, and a number of quarter-specific outlier binaries – generally one variable for each estimation. An added variable was retained if the diagnostic test values moved in favorable patterns indicative of improved specification. Juselius (2004, chapters 4, 7, and 9) recommends the following battery of diagnostics: (a) trace correlation as an overall goodness-of-fit indicator, (b) likelihood ratio test of autocorrelation, (c) Doornik-Hansen (D-H) tests for equation residual normality, (d) and indicators of skewness and kurtosis. The estimations were stopped when the array of diagnostic values failed to further improve with inclusion of additional variables. After achievement of an adequately specified levels VAR and unrestricted VEC, tests for parameter constancy and for the presence of I(2) trends were performed.

A statistically adequate VAR model emerged from two sets of sequential estimations. The first focused on each of the above-mentioned permanent shift binary variables (and a trend), and all 10 were retained. The second set further improved specification of the unrestricted VEC that included the ten just-mentioned variables. When a potential outlier was identified as extraordinarily influential based on a "large" standardized residual, an appropriately specified variable was included in equation 2's short run component, and retained if the battery of diagnostic values indicated improved specification.[30] Five quarter-specific transitory outlier binaries were included.[31]

[29] PCOOKIES and PMIXES are influenced by movements in confectionary and cocoa input costs. Analysis and information leading to the justification and formulation of this binary variable was received in private communications with market analysts of the U.S. International Trade Commission, Office of Industries, and with analysts from Labor, BLS during August, 2004. Also, see Babula and Newman (2005).

[30] We followed Juselius' (2004, chapter 6) analysis of potential outlier events. An observation-specific event was judged as a potentially "extraordinary" one if its standardized residual was 3.0 or more in absolute value. Such a rule for outliers was designed based on the 76-observation sample size using the Bonferoni criterion: $\text{INVNORMAL}(1\text{-}1.025)^{T}$, where T=76, INVNORMAL is a function for the inverse of the normal distribution

An adequately specified model should generate statistically normal residuals. Table 1 provides a battery of diagnostic test values for two estimations: the initially estimated unrestricted VEC before sequential estimations aimed at improved specification and with no deterministic variables, and for the unrestricted VEC judged as adequately specified after inclusion of centered seasonals, nine permanent shift and five outlier binaries, and a linear trend.[32] Table 1's results reveal clear benefits from efforts to improve specification: the model's ability to explain data variation increased 70 percent, as the trace correlation, a goodness of fit indicator for the 7-equation model, rose from 0.50 to 0.85.

A Doornik-Hansen (D-H) value tests the null hypothesis that the relevant equation's residuals are normal, which is rejected at the 1-percent level when the D-H value exceeds 9.2. In all cases but the ΔPMIXES equation, residuals follow normal behaviour for the unrestricted VEC after efforts at improved specification.[33] D-H values improved noticeably for the ΔPFLOUR and ΔPCEREAL equations.

Table 1 provides indications on skewness and kurtosis of each equation's residuals. Results suggest both sets of values generated by the model that benefited from specification efforts fell within ranges considered acceptably indicative of approximately normal residual behaviour.

function that returns the variable for the c-density function of a standard normal distribution (Doan 1996: Estima 2004). The Bonferoni variate had a 3.4 absolute value. Given that a number of seemingly influential quarter-specific events generated absolute standardized residual values within the 3.0–3.3 range, we chose a conservative Bonferoni criterion of absolute standardized residuals valued at 3.0 or more. Observations with absolute standardized residuals of 3.0 or more were considered potential outliers, and we specified an appropriately defined binary variable for the relevant observation for the sequential estimation procedure.

[31] To conserve space, we do not include extensive variable-by-variable analyses and estimation results. All included outlier binary variables were of the transitory "blip" form following formulation procedures in Juselius (2004, chapter 6), and were placed in the short run component of the model. The following variables are named numerically for the quarter during which the outlier event's influences were likely manifest: DTR8701, DTR8801, DTR9003, DTR9201, and DTR0201. For example, DTR8701 is defined as unity for 1987/88:01, and 0.0 otherwise. DTR8701 and DTR8801 likely captured quarter-specific expectationary influences of CUSTA's implementation and of the 1987-1990 drought on the U.S. wheat market not captured by the relevant permanent shift binary variables. DTR9003 likely captured influences on the cookies/crackers and mixes/doughs markets from wheat-related input cost effects from implementation of CUSTA and the 1990 U.S. farm bill that the relevant permanent shift binaries did not manage to capture. The effects captured by DTR9201 on PCEREALS likely arose from escalating prices of wheat, a major input cost for wheat-based breakfast cereals. DTR0201 likely captured effects on PWHEAT as the commodity boom of 2002-2004 unfolded.

[32] Each equation for the levels VAR and its unrestricted VEC equivalent was estimated over the 1986/87:03–2004/05:04 period. Four quarterly observations for the 1985/86 MY were set-aside for the Tiao and Box (1978) lag search. Given two lags, the full sample was 76 observations, with 74 observations in the estimation period.

[33] The ΔQWHEAT equation residuals generated a 9.3 D-H value that nearly equals the 9.2 critical value, and the test value's margin of excess over the critical value was too marginal to use as a sole criterion for conclusions of non-normal residual behavior. We opted to consider the equation's residuals as approximately normally behaving ones, given the generally favorable battery of other diagnostics generated by this equation.

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 to efforts at specification adequacy	After efforts at specification adequacy
Trace correlation	System-wide goodness of fit: large proportion desirable	0.5	0.85
ARCH tests for heteroscedasticity	Ho: no heteroscedasticity by 1st and 4th lags. Reject for $p<0.05$	lag 1: 101.2 (p=0.000) lag 4: 101.1 (p=0.000)	lag 1: 61.1 (p=0.12) lag 4: 55.8 (p=0.23)
Doornik-Hansen tests for normal residuals	Ho: residuals are normal. Reject for values above 9.2 critical value		
ΔPWHEAT		8.3	4.9
ΔQWHEAT		13.4	9.3
ΔPFLOUR		15.9	7.9
ΔPMIXES		9.4	13.4
ΔPBREAD		1.3	1.2
ΔPCEREAL		36.3	2.5
ΔPCOOKIES		1.8	2.2
Skewness (kurtosis) values	skewness: ideal is 0.0; small absolute values ≤1.0 acceptable. kurtosis: ideal is 3.0; values 3-5 acceptable.		
ΔPWHEAT		0.83 (4.46)	0.05 (3.9)
ΔQWHEAT		-0.80 (3.04)	-0.34 (4.5)
ΔPFLOUR		1.3 (6.0)	0.10 (4.2)
ΔPMIXES		0.38 (4.56)	0.15 (4.8)
ΔPBREAD		0.28 (3.1)	0.28 (3.0)
ΔPCEREAL		-1.36 (10.5)	0.39 (3.4)
ΔPCOOKIES		0.14 (3.35)	-0.12 (3.4)

Specification efforts did not render estimated residuals for the ΔPMIXES equation that behaved with approximate normality. However, table 1 shows clear, substantial progress from specification efforts, and that the entire system generates approximately normal residuals as a system.[34] We followed Juselius (2004, chapter 4) and concluded that overall evidence suggested that the VAR system achieved reasonable adequacy of specification despite ΔPMIXES' weak evidence of normality.

COINTEGRATION: CHOOSING AND IMPOSING REDUCED RANK ON THE ERROR CORRECTION SPACE

The endogenous variables are shown below to be I(1), and their differences are I(0). Cointegrated variables are driven by common trends, and stationary linear combinations (Juselius 2004, p. 86). The Π-matrix in equations 2 or 3 is a 7 by 7 matrix equal to the product of a p by r matrix β of long run error correction coefficients that under cointegration, combine into $r \leq p$ stationary linear combinations of the seven wheat-related variables and a p by r matrix α of adjustment speed coefficients (Johansen and Juselius 1990, 1992). As a result of a reduced rank for Π, β'*x(t) is I(0), even though x(t)'s seven variables are nonstationary.

Determination of cointegration rank is a three-tiered process. First, one conducts trace tests of Johansen and Juselius (1990, 1992). Second, one examines patterns of characteristic roots of companion matrices generated under relevant assumptions of reduced rank. And third, one examines plotted cointegrating relationships for elements of stationary behaviour.

NESTED TRACE TESTS AND OTHER EVIDENCE FOR CHOOSING THE REDUCED RANK OF II

Table 2 provides nested trace test evidence for rank determination. Evidence at the five percent significance level is sufficient to soundly reject the first five nested hypotheses, suggesting that $r \leq 4$. Evidence is marginally sufficient to reject the sixth that $r \leq 5$ as the test value approaches the 42 critical fractile, and is insufficient to reject that $r \leq 6$. Trace tests alone suggest that r=6 and that there are six CVs, although evidence marginally rejects that $r \leq 5$, suggesting that r may be less than 6. We follow Juselius' (2004, chapter 8) suggestion against sole reliance on trace test evidence to determine rank.

If r is an appropriate choice, then one expects p-r characteristic roots that are unity or near-unity in the companion matrix. When r is imposed, and there are p-r+1 roots that are unity or near unity, then one could consider reducing rank to r-1. Patterns of characteristic roots under alternative r-assumptions suggest that r is likely between 1 and 3 with evidence pointing especially to r=3.[35]

[34] An anonymous Reviewer pointed out that readers should cautiously interpret inference statistics for the ΔPMIXES equation.

[35] This analysis follows Juselius (2004, chapter 8). We examined patterns of characteristic roots for companion matrices generated under all possible reduced-rank levels: 1, 2, 3, 4, 5, 6. Because of space considerations, we do not report all six matrices and full analyses, that are available from the authors on request. Summarily, patterns of characteristic roots under r=1 through 6 suggests that reduced rank is less than 6, and most likely within the range of 1-3, with evidence pointing particularly to a reduced rank of 3. If r=3 is appropriate, then one expects p-r or 4 characteristic roots of unity, with the fifth and subsequent roots less than unity. When r=3 was imposed, the first five of the characteristic roots were as follows, suggesting that r=3: 1.0, 1.0, 1.0, 1.0, 0.87, with the fifth, 0.87, below unity.

Table 2. Trace test statistics and related information for nested tests for rank determination

Null hypothesis	Trace value	95% fractile (critical value)	Result
Rank or r ≤ 0	370.6	166.5	Reject null that rank is zero
Rank or r ≤ 1	279.8	133.7	Reject null that rank ≤ 1
Rank or r ≤ 2	184.8	104.8	Reject null that rank ≤ 1
Rank or r ≤ 3	120.6	79.9	Reject null that rank ≤ 1
Rank or r ≤ 4	89.6	59.0	Reject null that rank ≤ 1
Rank or r ≤ 5	45.7	41.9	(Marginally) Reject null that rank ≤ 1
Rank or r ≤ 6	15.7	28.7	Fail to reject that rank ≤ 6

Notes. - As recommended by Juselius (2004, p. 171), CATS2-generated fractiles are increased by 9*1.8 or 16.2 to account for the 9 permanent shift binary variables restricted to lie in the cointegration space. Trace values are corrected with Bartlett's adjustment for small samples.

The plots of the three CVs are in figures 1, 2, and 3. The BETA*x(t) plots are for the model with short run effects, and the BETA*R1(t) plots are for the model corrected for short run effects, with Juselius (2004, chapter 8) favouring the latter as more reliable.

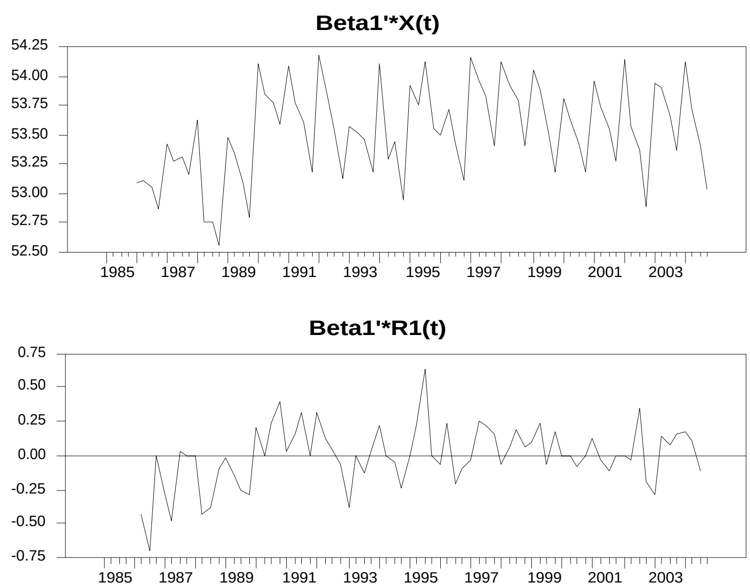

Figure 1. Cointegrating relation 1 with and without correction for short run effects

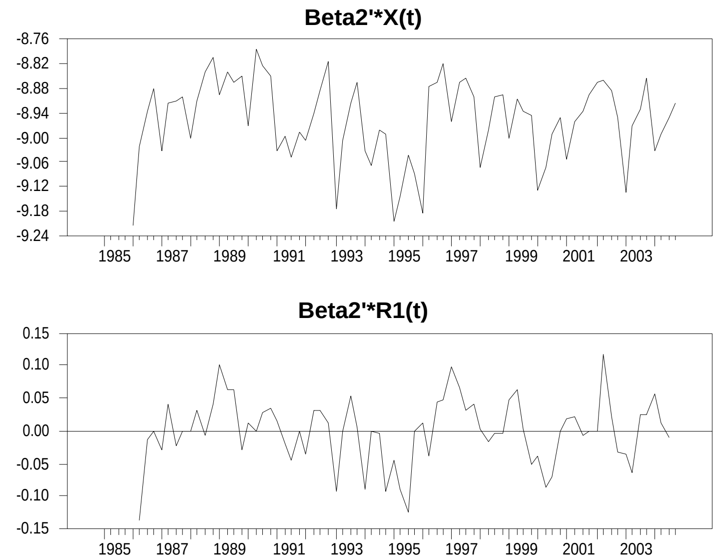

Figure 2. Cointegrating relation 2 with and without correction for short run effects

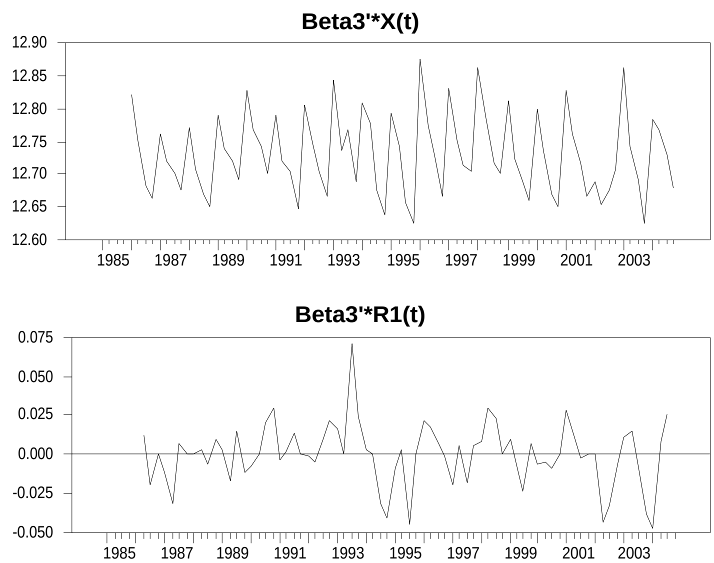

Figure 3.Cointegrating relation 3 with and without correction for short run effects

Figures 1-3 suggest that all three CVs are for the most part stationary, and hence that r=3 rather than one or two: plots cycle and mean-revert frequently, and variation levels appear constant (Juselius 2004, chapter 8).[36] All three evidence sources above suggest that the reduced rank of equation 2's Π-matrix is likely three, with three stationary linear combinations of the seven I(1) wheat-based variables error-correcting the system.

Further Diagnostic Tests for Parameter Constancy and I(2) Trends.

Two final diagnostic tests were applied to the cointegrated VEC and results suggested that the model achieved statistical adequacy: tests are for constancy of error correction parameter estimates and for the presence of trends that are integrated of order-2 or I(2). The "known" beta test detailed in Juselius (2004, pp. 186-190) tests if there is constancy or time-invariance of cointegration parameter estimates. This method tests if the full sample "baseline" model's cointegration relations could have been accepted as those of each recursively estimated model over the 2002/03:02-2004/05:04 period. Values are indexed by the 95 percent fractile, and should ideally be unity or less to indicate parameter estimate constancy. All values were below unity and suggested time-invariant estimates. Known beta plots were not included to conserve space, and are available from the authors.

Nielsen (2002) and Juselius (2004) noted that imposing reduced rank restrictions on an unrestricted VEC's error correction space when there are I(2) trends encounters well-known and potentially serious econometric problems, including compromised inference, because the data still have unit roots. Evidence from a series of tests for I(2) trends was sufficient to reject the null hypotheses of I(2) trends in all cases.[37]

Equations 4-6 are the three cointegrating relationships that emerged after imposing rank and re-estimation with Johansen and Juselius' (1990, 1992) reduced-rank estimator. Estimates are not yet restricted for statistically supported restrictions that emerge from the next section's hypothesis tests.

$$\begin{aligned}QWHEAT = &-1.77*PWHEAT + 2.30*PFLOUR - 4.94*PMIXES +\\ &4.51*PBREAD - 0.28*PCEREAL -10.80*PCOOKIES-0.49*URUGUAY +\\ &0.59*CUSTA -0.11*NAFTA + 0.52*QUOTA +1.25*FBILLS + 0.65*TITLE7\\ &+0.07*DROU88 + 0.40*HIDD9396 - 0.26*CONFECT + 0.01*TREND\end{aligned} \tag{4}$$

$$\begin{aligned}PWHEAT = &\ 0.10*QWHEAT + 2.24*PFLOUR -3.18*PMIXES +\\ &1.78*PBREAD -2.51*PCEREAL + 3.60*PCOOKIES\\ &-0.12*URUGUAY - 0.02*CUSTA -0.04*NAFTA + 0.05*QUOTA -\\ &0.26*FBILLS + 0.17*TITLE7-0.09*DROU88 + 0.19*HIDD9396 -\\ &0.04*CONFECT-0.01*TREND\end{aligned} \tag{5}$$

[36] Juselius (2004, chapter 8) notes that a CV will not likely follow perfectly stationary behavioral paths. Each CV generally behaves with stationarity, with a couple of short term deviations notwithstanding: some short-lived cycling in 1995 for CV1 and CV2, and some volatile behavior in 1993-1994 for CV3.

[37] Nielsen's (2002) chosen test for I(2) trends compares the I(2) model of H(r,s): p variables, r I(0) directions, s I(1) directions, and p-r-s I(2) directions, against the unrestricted model of H(p). In our case, p=7 and r = 3. The null hypotheses are H(r,s)|H(p) and one rejects the null when models are too restricted. Rejecting all models where (p-r-s) > 0 implies evidence that is sufficient to reject I(2) trends. To conserve space, we do not report results and analysis of the 28 tests where (p-r-s) > 0. In all cases, however, evidence at both the one- and five-percent significance levels was sufficient to reject the null hypothesis of I(2) trends. Also see Juselius (2004, chapter 16).

PFLOUR = 0.53*PWHEAT - 0.12*QWHEAT + 0.09*PMIXES - 1.11*PBREAD +0.14*PCEREAL - 0.76*PCOOKIES
+ 0.06*URUGUAY + 0.03*CUSTA - 0.08*NAFTA -0.03*QUOTA
+0.08*FBILLS - 0.08*TITLE7-0.03*DROU88 - 0.02*HIDD9396 -
0.05*CONFECT+ 0.02*TREND

(6)

HYPOTHESIS TESTS AND INFERENCE ON THE ECONOMIC CONTENT OF THE THREE COINTEGRATING RELATIONS[38]

We begin with equations 4, 5, and 6, the unrestricted CVs, conduct a series of hypothesis tests on the $\Pi = \alpha'*\beta$, and then re-estimate the system with the statistically-supported restrictions imposed. Hypothesis tests on the beta coefficients take the form:

$$\beta = H^* \varphi \tag{7}$$

Above is a pl by pl vector of coefficients on variables included in the cointegration space; H is a pl by s design matrix, with ‘s” being the number of unrestricted or free beta coefficients; and φ is an s by r matrix of the unrestricted beta coefficients. The hypothesis test value or statistic is:

$$-2\ln(Q) = T^*\sum[(1-\lambda_i^{\,*})/(1-\lambda_i)] \text{ for } i = 1, 2, \text{ and } 3\ (=r). \tag{8}$$

The asterisked (non-asterisked) eigenvalues (λ_i , i = 1-3) are generated by the model estimated with (without) the tested restriction(s) imposed.

Likewise, the hypothesis tests concerning the α or adjustment speed coefficients permit a characterization of relative speeds of error-correcting adjustment with which the system responds to a given shock. The null hypothesis or H(0) is:

$$H(0)\colon \alpha = A^*\psi \tag{9}$$

Above, A is a p by s design matrix; s is the number of unrestricted coefficients in each of the r=3 columns of the α matrix; and ψ is the s by r matrix of the non-restricted or “free” adjustment speed coefficients. Equation 8's test statistic also applies here, and is distributed asymptotically as a chi-squared distribution with degrees of freedom equal to the number of imposed coefficient restrictions. Hypothesis tests on the betas, followed by tests on the alphas, are provided below.

There are three sets of hypothesis tests on the beta coefficients. The first set of six examines if each endogenous variable is stationary under the imposed rank of three. Second, there are 17 “exclusion” hypothesis tests of whether each of the variables included in the CVs

[38] This methods section closely follows the those developed and/or refined in Johansen and Juselius (1990, 1992) and Juselius (2004, chapter 11).

have zero-valued β-estimates. A third set is performed on individual β-estimates in equations 11-13, with any statistically supported stationarity and/or exclusion restrictions imposed.

Tests of Stationarity. Juselius (2006; 2004, pp. 220-222) and Juselius and Toro (2005) contend that univariate one-dimensional unit root tests are not appropriate for such a p-dimensional VAR as our 7-dimensional model. They instead recommend a likelihood ratio test of each endogenous variable's stationarity within a system setting and given the imposed rank (here r=3). They argue that using univariate (say Dickey-Fuller) critical values for a one-dimensional or univariate VAR model should not be used to appropriately test for unit roots in a p-dimensional (here 7-dimensional) VAR setting (Juselius 2006; Juselius 2004, pp. 220-222; and Juselius and Toro 2005). The recommended likelihood ratio tests examine if each endogenous variable itself constitutes a separate stationary cointegrating relation, with a unity value for the tested variable's betas. Equation 7 is rewritten as follows:[39]

$$\beta^c = [b,\varphi] \tag{10}$$

With a rank of r=3, equation 8's test value is distributed under the null hypothesis of stationarity as a chi-squared variable with three degrees of freedom. Evidence was sufficient to reject that all seven endogenous variables were stationary, leading to our conclusion that they are I(1).[40]

Tests of Beta Exclusions. There are p1=17 variables in equation 2's cointegration space, and so in turn, as many exclusion tests are performed. Failure to reject the null that a variable's betas are zero-valued suggests that the variable should be excluded from the cointegration space.[41] On balance, evidence suggested that all variables should be, at least initially, retained in the cointegration space.[42]

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[39] This test can be conducted in CATS2 (beta version) in two settings: with and without inclusion of the nine deterministic variables and trend restricted to the cointegration space. We chose to include these deterministic variables in the tests, due to the institutional importance of events for which the variables were defined, as discussed in earlier research (BBP, 2004; RBR 2002). Note that results from both settings were similar. In equation 10, β^c is the p1 by r (17 by 3) beta matrix with one of the variable's levels restricted to a unit vector; b is a p1 (or 17) by 1 vector with a unity value corresponding to the relevant variable whose stationarity is being tested; and φ is a p1 by (r-1) or 17 by 2 matrix of the remaining two unrestricted cointegrating vectors.

[40] Equation 8's test value is distributed under the null hypothesis as a chi-squared variable with, here, 3 degrees of freedom, and calculated values were as follows (with parenthetical p-values): 32.5 (0.000) for PWHEAT; 31.88 (0.000) for QWHEAT; 25.54 (0.000) for PFLOUR; 17.6 (0.000) for PMIXES; 23.35 (0.000) for PBREAD; 20.4 (0.000) for PCEREAL; and 11.07 (0.03) for PCOOKIES. The null hypothesis was rejected for p-values below 0.05, corresponding to the five-percent significance level.

[41] The hypothesis test value in equation 7 would include a 17 by 3 β-vector; a 17 by 16 design matrix, H, with 16 being the number of unrestricted beta coefficients in each relation; and a 16 by 3 matrix φ of 16 unrestricted coefficients in each of the three cointegrating relationships (Juselius 2004, chapter 10). Basically, the φ matrix is the β-matrix without the beta coefficients for the variable being tested for exclusion.

[42] The exclusion test values (and parenthetical p-values) for the following 15 variables reflected evidence at the 5-percent significance level that was sufficient to reject the null hypotheses of zero-valued beta coefficients: 28.03 (0.000) for PWHEAT; 29.1 (0.0000) for PFLOUR; 23.1 (0.0000) for PMIXES; 17.1 (0.001) for PBREAD; 22.65 (0.0000) for PCEREAL; 27.28 (0.0000) for PCOOKIES; 9.75 (0.03) for URUGUAY; 9.12 (0.03) for NAFTA; 8.3 (0.04) for QUOTA; 30.12 (0.000) for FBILLS; 17.76 (0.000) for TITLE7; 8.9 (0.03) for DROU88; 20.15 (0.000) for HIDD9396; 9.95 (0.04) for CONFECT, and 20.5 (0.000) for TREND. Evidence at the five percent significance level was not sufficient to reject the null hypothesis of zero-valued beta coefficients for QWHEAT, with a test value of 5.7 and p-value of 0.12 and for CUSTA with a test value of 7.0 and a p-value of 0.07. We decided to include CUSTA, given the marginal test value that suggested evidence that was sufficient to reject the null of zero beta's at the seven percent significance level. We also chose to retain QWHEAT in the error-correction space, despite the test value's p-value of 0.12, because of substantial evidence from BBP (2004, pp. 12-18). Their results suggested that QWHEAT has rich and bi-

Set of Sequential Hypothesis Tests on Individual Beta Coefficients. Since no variables were excluded or stationary, one must now meet the rank condition of identification by imposing at least r-1 identifying restrictions directly on each of equations 4-6. (Juselius 2004, pp. 245-246). These added hypotheses arose from theory, market knowledge, prior research, and/or suggestions implied by coefficient estimates and are tested using equations 7-8 (Juselius 2004, pp. 245-246). A restriction to be tested is imposed, the model re-estimated with the well-known Johansen-Juselius reduced rank estimator, and the test value for the hypothesized restriction calculated. If statistically supported at the 5-percent level of statistical significance (hereinafter, the 5-percent level), the restriction is retained. We repeated this process on the three CVs. Space limitations preclude reporting results for all sequential estimations, although table 3 summarizes this multi-iterative process. Here, we briefly direct the reader through the array of testable hypotheses to provide a systematic method for choosing testable hypotheses. Economic and policy analysis/implications of these restrictions in the finally restricted CVs are reserved for an ensuing section below.

Test set 1 (TS-1) provides the first set of zero restrictions on selected β-estimates: three on PCOOKIES, PMIXES, and PCEREAL in CV1 normalized on QWHEAT; two on QWHEAT and TREND in CV2 normalized on PWHEAT; and three on QWHEAT, PCOOKIES, and PMIXES in CV3, normalized on PFLOUR. These restrictions imply low levels of influence on the normalized variables, and were chosen based on previous research's analyses of forecast error variance (FEV) decompositions and/or impulse response simulations generated by VAR models of the same markets (BBP 2004, pp. 14-18; RBR 2002, pp. 109-111). TS-1 restrictions meet the rank condition of identification, although as is often the case, the test value fails to initially accept the restrictions, suggesting the need for added economic content through other restrictions to generate a statistically accepted set at the chosen 5-percent significance level (when p-values > 0.05).

Test sets 2 through 8 postulate, impose, re-estimate, and then test a series of zero restrictions on beta coefficients in CV1, CV2, and CV3 that arise from theory, counsel from recognized market experts, and patterns of the estimated beta t-values (statistically insignificant ones, generally). More economic content is yet required for acceptance at the chosen 5-percent level.

Test set 8's beta estimates suggested the following testable hypothesis in CV1 that was added to render test set 9: β(TITLE7) = β(NAFTA) that suggests that the AD/CVD investigations and NAFTA's 1994 implementation (and other concurrent events) had, on average, equal quarterly effects on QWHEAT, CV1's dependent variable.[43] Test set 9 with this restriction imposed generated evidence that supported this equality restriction: t-values of -5.9 for both restricted CV1 coefficients, and TS-9 restrictions were statistically supported at the 2-percent level.

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directional causal interplay at long run and short run horizons in the system. QWHEAT's causal importance to the system appeared to escalate at the longer run horizons in the BBP analysis of FEV decompositions. To exclude QWHEAT from our long run space based solely on exclusion test evidence would seem overly simplistic and would ignore the strong findings of BBP's recent and related study. As well, this test value for QWHEAT approaches rejection at the 10 percent level and may indicate that the variable could be included in some and excluded from other CVs. Results from the third set of hypothesis tests presented below indeed verified that QWHEAT should remain in CV1 and be excluded from CV2 and CV3.

[43] The sequential estimation under test set 8's restrictions yielded the following CV1 results: coefficient value of -0.881 on β(TITLE7) with a t-value of -4.5, and of -0.883 on β(NAFTA) with a t-value of -5.5. Clearly, these significant beta coefficients in CV1 should be tested for equality.

Table 3. Sequential Hypothesis Tests on Beta Estimates in the Error-Correction Space of the U.S. System of Wheat-Based Products

Tested restrictions restriction numbers (Marginally added restriction(s) in bold)	Explanation, reasoning	Test values, test results, and interpretation of coefficient estimates
Test set 1(TS-1): Various restrictions suggested by previous research and needed to meet rank condition of identification		
3 in CV1: β(PCEREAL)=β(PMIXES)= β(PCOOKIES = 0 *2 in CV2*: β(QWHEAT)=β(TREND)=0 *3 in CV3:* β(QWHEAT)=β(PCOOKIES)= β(PMIXES) = 0	Suggested by BBP(2004) FEV analysis. Suggested by data analysis and BBP(2004), RBR(2001). BBP(2004), RBR(2001)	Test value of 19.2 (df=2) with p=0.000 suggests more restrictions need to be found for a statistically supported set at the five percent significance level. *Estimate interpretation*: t[β(CUSTA)]=0.1 in CV1; add zero restriction for CUSTA.
Test set 2: previous TS-1 restrictions plus β(CUSTA)=0 in CV1.		
4 in CV1: 3 TS-1 restrictions retained, plus **β(CUSTA)=0** *2 in CV2:* 2 TS-1 restrictions retained. *3 in CV3:* 2 TS-1 restrictions retained.	Weak t-value on β(CUSTA), prior estim'n.	Test value of 19.1 (df=3) with a p=0.003 suggests some progress towards statistical acceptance. More restrictions needed for a statistically acceptable set. *Estimate interpretation*: t[β(HIDD9396)]= 1.8, in CV1; add zero restriction on HIDD9396.
Test set 3: TS-2 restrictions plus β(HIDD9396)=0 in CV1.		
5 in CV1: 4 prior TS-2 restrictions retained plus **β(HIDD9396)=0** *2 in CV2:* 2 TS-2 restrictions retained. *3 in CV2*: 3 TS-2 restrictions retained.	Weak t[β(HIDD9396)] prior estimation	Test value of 19.9 (df=4) with p-value of 0.001 suggests some progress; more restrictions needed for statistically supported set. *Estimate interpretation:* t[β(CUSTA)]=-0.6 in CV2; add zero restriction on CUSTA.
Test set 4: TS-3 restrictions plus β(CUSTA) in CV2.		
5 in CV1: 4 TS-3 restrictions retained, *3 in CV2:* TS-3 restrictions retained plus **β(CUSTA) = 0** *3 in CV3*: TS-3 restrictions retained	Weak t[β(CUSTA)], prior estimation	Test value of 19.9 (df=5) with p-value of 0.0013 suggests some progress in statistical support; more restrictions needed for statistically supported set. *Estimate interpretation:* t[β(QUOTA)] = -2.2 in CV2.; add as zero restriction.

Table 3. Continued

Tested restrictions restriction numbers (Marginally added restriction(s) in bold)	Explanation, reasoning	Test values, test results, and interpretation of coefficient estimates
Test set 5: TS-4 restrictions plus β(QUOTA) = 0		
5 in CV1: 5 TS-4 restrictions retained. *4 in CV2:* 3 TS-4 restrictions retained, plus **β(QUOTA) = 0** *3 in CV3:* 3 TS-4 restrictions retained.	Weak t[β(QUOTA)] in prior estimation.	Test value of 21.1 (df=6) with p-value of 0.02 suggests progress: statistical acceptance at 2% level. More restrictions needed for acceptable set at 5% level. *Estimate interpretation*: t[β(URUGUAY)] = 1.0; add as zero restriction in CV2.
Test set 6: TS-5 restrictions plus β(URUGUAY) = 0 in CV2		
5 in CV1: 5 TS-5 restrictions retained. *5 in CV2:* 4 TS-5 restrictions retained, plus **β(URUGUAY) = 0** *3 in CV3*: 3 TS-5 restrictions retained.	Weak t[β(URUGUAY)], prior estimation	Test value of 21.3 (df=7), p-value = 0.0033. More restrictions needed for statistically supported set. *Estimate interpretation*: t[β(DROU88)] = -0.02 in CV3; add as zero restriction.
Test set 7: TS-6 restrictions plus β(DROU88) = 0 in CV3.		
5 in CV1: 5 TS-6 restrictions retained. *5 in CV2*: 5 TS-6 restrictions retained *4 in CV3:* 3 TS-6 restrictions retained, plus **β(DROU88) = 0**	Weak t[β(DROU88)], prior estimation	Test value of 21.3 (df=8), p-value = 0.007. More restrictions needed for statistically supported set. *Estimate interpretation:* t[β(FBILLS)] = 1.3 in CV3; add as zero restriction.
Test set 8: TS-7 restrictions plus β(FBILLS) = 0 in CV3		
5 in CV1: 5 TS-7 restrictions retained. *5 in CV2:* 5 TS-7 restrictions retained *5 in CV3:* 4 TS-7 restrictions retained, plus **β(FBILLS)** = 0	Weak t[β(FBILLS)], prior estimation	Test value of 21.4(df=8), p-value = 0.01. More restrictions needed for statistically supported set at 1% level. *Estimate interpretation:* β(TITLE7) = β(NAFTA) in CV1; add as equality restriction.
Test set 9: TS-8 restrictions plus β(TITLE7) = β(NAFTA) in CV1.		
6 in CV1: 5 TS-8 restrictions retained, plus **β(TITLE7) = β(NAFTA)** *5 in CV2:* 5 TS-8 restrictions retained *5 in CV3*: 5 TS-8 restrictions retained	Examination of last estimates: average market impacts of TITLE7 and NAFTA events about equal on QWHEAT.	Test value of 21.6 (df=10) with p-value of 0.02 suggests progress in statistical acceptance at 2%; more restrictions needed for acceptance at 5% level. *Estimate interpretation:* β(PCEREAL) = -β(PCOOKIES) in CV2; add as inequality restriction.

Table 3. Continued

Tested restrictions restriction numbers (Marginally added restriction(s) in bold)	Explanation, reasoning	Test values, test results, and interpretation of coefficient estimates
Test set 10: TS-9 restrictions plus β(PCEREAL) = -β(PCOOKIES) in CV2.		
6 in CV1: 5 TS-9 restrictions retained. *6 in CV2:* 5 TS-9 restrictions retained, plus **β(PCEREAL) = -β(PCOOKIES)** *5 in CV3:* 5 TS-9 restrictions retained.	Examination of last estimates. Inequality restriction suggests that PWHEAT is dependent on the difference in prices of wheat cereal and cookies/crackers.	Test value of 21.6 (df=11) with p-value of 0.03 suggests progress in statistical acceptance at 3% level; more needed for acceptance at 5% level. *Estimate interpretation*: β(QUOTA) =β(TITLE7) in CV3; add as equality restriction.
Test set ll: TS-10 restrictions plus β(QUOTA) =β(TITLE7) in CV3.		
6 in CV1: 5 TS-10 restrictions retained. *6 in CV2:* 6 TS-10 restrictions retained *6 in CV3***:** 5 TS-10 restrictions retained, plus **β(QUOTA) =β(TITLE7)**	Examination of last estimates: average market impacts were about the same from QUOTA and TITLE7 events.	Test value of 21.6 (df=12) with p-value of 0.042 suggests progress: statistical acceptance at 4% level. More restrictions needed for acceptance at 5%. *Estimate interpretation:* : β(QUOTA) =β(TITLE7) = -β(NAFTA) in CV3.
Test set 12: TS-11 restrictions plus β(QUOTA) =β(TITLE7) = -β(NAFTA) in CV3.		
6 in CV1: 6 TS-11 restrictions retained. *6 in CV2:* 6 TS-11 restrictions retained *7 in CV3:* 6 TS-11 restrictions retained, plus **β(QUOTA) =β(TITLE7) = -β(NAFTA)**	Examination of last estimates: average market impacts of QUOTA events or of TITLE7 events were about negated by NAFTA events.	Test value of 21.6 (df=13) with p-value of 0.063 suggests statistical acceptance at more than the desired 5% level: at 6% level. *Estimate interpretation:* β(PBREAD) = -β(PCEREAL) in CV3
Test set 13: TS-12 restrictions plus β(PBREAD) = -β(PCEREAL) in CV3		
6 in CV1: 6 TS-11 restrictions retained. *6 in CV2:* 6 TS-11 restrictions retained *8 in CV3*: 7 TS-11 restrictions retained, plus **β(PBREAD) =-β(PCEREAL)**	Examination of last estimates: average market impacts depend on difference between bread and wheat cereal product prices.	Test value of 22.5 (df=14) and p-value of 0.07 reflects that we have achieved evidence of a statistically supported set of restrictions at 7% level (above desired 5% level).

Test set 9's coefficient estimates in CV2 normalized on PWHEAT suggest that the following is clearly a testable hypothesis: β(PCEREAL) = -β(PCOOKIES), which implies that PWHEAT is influenced by the difference in PCOOKIES and PCEREAL.[44] Adding this

[44] The sequential estimation under test set 9 generated for CV2 the following: β(PCOOKIES) = 15.8 (t = 7.1) and β(PCEREAL) = -16.8 (t=-12.8). This places the following as a testable hypothesis: β(PCOOKIES) = -β(PCEREAL), which suggests that CV2's dependent variable, PWHEAT, is a function of the difference of these two prices. The economic and/or market importance of this restriction is discussed below when the finally-restricted CVs are examined.

latter condition to TS-9 renders test set 10. The test value improves, as evidence accepts the restrictions at the 3-percent level.

Test set 10's coefficient estimates suggested that β(QUOTA) = β(TITLE7) in CV3, suggesting that the set of two temporary U.S. TRQs on certain imports of Canadian wheat, and the array of AD/CVD duties during 2002-2004 had (collectively with other concurrent events) approximately equal market impacts.[45] The addition of this equality restriction in CV3 to TS-10 rendered test set 11, which generated restricted coefficients which were statistically significant (t-values of -6.2), and a test value which accepted restrictions at an increased 4 percent significance level (p-value = 0.042).

Test set 11's coefficient estimates suggested further that β(QUOTA) = β(TITLE7) = -β(NAFTA) in CV3. [46] The interpretation of this multi-parameter restriction is left to the next subsection on economic content. Adding this restriction in CV3 to TS-11 rendered test set 12 in table 3. The re-estimation restricted for this equality condition generated statistically significant coefficients, and strong support for, this restriction: t = -7.3 for betas on QUOTA and TITLE 7; t = -7.3 on NAFTA. The test value's p-value (0.063) reflected evidence that accepted the restrictions at the 6-percent level.

Test set 12's coefficient estimates suggested that in CV3, perhaps β(PCEREAL = -β(PBREAD), which suggests that market effects through PFLOUR hinge on the difference between PBREAD and PCEREAL.[47] We provide an interpretation of this restriction in the ensuing subsection on economic content. This restriction was added to TS-12's to render test set 13. Evidence suggested that the last CV3 restriction was statistically significant (t-values of ± 10.3), and that TS-13's restrictions were accepted at the 7-percent significance level (p-value of 0.07), which exceeds our decision rule of 5-percent.

HYPOTHESIS TESTS ON THE ADJUSTMENT SPEED OR A COEFFICIENTS

A principal hypothesis on the estimated adjustment speed coefficients is if each of the variables is weakly exogenous. A variable is weakly exogenous if it influences the error-correction process without itself adjusting or responding to the process, thereby implying a one-way causal relation to the equilibrating relation. Equivalently, one tests if, given the statistical significance of at least some of a variable's β-estimates, the variable's r=3 α-coefficients are all zero (Juselius 2004, pp. 231-232). Evidence in all cases was sufficient to reject the null of weak exogeneity.[48]

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[45] Test set 10's estimates generated the following in CV3: β(TITLE7) of -0.20 (t = -3.9) and β(QUOTA) = -0.19 (t =-3.3). This suggests a testable hypothesis of β-equality that, when added to TS-10, rendered test set 11.

[46] Test set 11's restrictions and sequential estimation generated the following CV3 results: β(QUOTA) =β(TITLE7) = -0.195 (t= -6.2), and β(NAFTA) = 0.196 (t=3.3).

[47] Test set 12's coefficient estimates generated the following in CV3: β(PCEREAL) = 2.92 (t=11.1) and β(PBREAD) = -3.4 (t=-7.3).

[48] The weak exogeneity test values and (parenthetical) p-values were as follows: 32.9 (0.000) for PWHEAT, 16.2 (0.001) for QWHEAT, 26.5 (0.000) for PFLOUR, 7.4 (0.06) for PMIXES, 10.1 (0.02) for PBREAD, 15.1 (0.002) for PCEREAL, and 36.6 (0.000) for PCOOKIES. Evidence was sufficient at the 5-percent level or less to reject the null of zero-valued α-coefficients for all endogenous variables except PMIXES. Evidence was sufficient at the 6-percent level to reject PMIXES weak exogeneity – a very marginal result. BBP's (2004, pp. 16-19) analysis of FEV decomposition patterns generated by a Bernanke (1986) structural VAR (with directed acyclic graph analysis) of the same markets revealed evidence of endogenous participation of PMIXES,

ECONOMIC ANALYSIS OF THE THREE COINTEGRATING RELATIONSHIPS

The fully restricted CVs are equations 11-13. To conserve space, we present CV1, CV2, and CV3 in abbreviated form, with DCV1, DCV2, and DCV3 reflecting vectors of econometric estimates for permanent shift binaries that we deemed to be of lesser relevance and/or interest, but whose inclusion was required to achieve a statistically adequate model.[49] The CVs are followed by the α-estimates. Parenthetical t-values reflect that most estimates have achieved clear statistical strength.

Limitations of imprecision in interpreting the coefficient estimates on binary (dummy) variables are well known (USITC 1995). Typically, partial effects cannot be *solely* attributed to an event for which a binary is defined, but must be collectively attributed to that event and all other relevant events that concurrently occurred during the period (USITC 1995; Babula 1997). For ease of exposition, we provide uni-event attribution with multi-event attribution implied.

There are three CV's: the first appears to be a U.S. wheat supply, and the other two, long run price transmission relationships. We first discuss the supply and price transmission relationships, and then collectively analyze the information from the coefficient estimates on the permanent shift binary variables from all three CVs.

$$\begin{aligned}
QWHEAT = {} & \underset{(10.0)}{5.76}{*}PWHEAT - \underset{(-10.24)}{10.95}{*}PFLOUR - \underset{(-9.2)}{13.04}{*}PBREAD \\
& - \underset{(-5.98)}{0.91}{*}NAFTA - \underset{(-5.98)}{0.91}{*}TITLE7 - \underset{(-4.47)}{0.72}{*}QUOTA - \underset{(-2.7)}{0.30}{*}CONFECT \\
& + \underset{(9.89)}{0.18}{*}TREND + DCV1
\end{aligned} \tag{11}$$

$$\begin{aligned}
PWHEAT = {} & \underset{(5.2)}{3.19}{*}PFLOUR - \underset{(-5.8)}{10.9}{*}PMIXES + \underset{(7.86)}{6.84}{*}PBREAD + \\
& \underset{(\pm 13.4)}{18.03}{*}(PCEREAL - PCOOKIES) - \underset{(-5.78)}{1.15}{*}NAFTA + \underset{(8.1)}{0.56}{*}TITLE7 \\
& - \underset{(-2.96)}{0.55}{*}CONFECT + DCV2
\end{aligned} \tag{12}$$

particularly among other wheat-using value-added product prices. Given the marginal test value and this added BBP evidence of PMIXES' endogenous participation, we chose to treat PMIXES as endogenous and not weakly exogenous.

[49] In the vector definitions that follow, t-values are included parenthetically. DCV1 is a vector of the following CV1permanent shift binary variable coefficient estimates: 1.0*URUGUAY (t= 6.1); -1.6*FBILLS (t = -8.95); -0.34*DROU88 (t=-2.8). DCV2 is a vector of the following CV2 permanent shift binary variable coefficient estimates: -2.45*FBILLS (t= -9.1); -0.53*DROU88 (t= -3.98); 1.63*HIDD9396 (t= 8.1). DCV3 is a vector of the following CV3 permanent shift binary variable coefficient estimates: -0.21*CUSTA (-3.0); -0.57*HIDD9396 (t= -8.4); 0.31*URUGUAY (t=4.0).

PFLOUR = 0.49*PWHEAT + 5.0*(PBREAD - PCEREAL) - 0.33*QUOTA
(4.6) (±10.3) (-7.1)
–0.33*TITLE7 + 0.33* NAFTA + 0.19*CONFECT + 0.02*TREND
(-7.1) (7.1) (2.2) (8.9)
+ DCV3

(13)

ALPHA	Alpha1	Alpha2	Alpha3
ΔPWHEAT	0.0997	-0.1407	-0.1031
	(3.8849)	(-6.5953)	(-2.1826)
ΔQWHEAT	-0.0356	0.0471	0.3062
	(-0.7805)	(1.2400)	(3.6431)
ΔPFLOUR	-0.0123	-0.0501	-0.0293
	(-0.9774)	(-4.8001)	(-1.2669)
ΔPMIXES	-0.0040	0.0028	0.0102
	(-1.0472)	(0.8823)	(1.4613)
ΔPBREAD	-0.0119	0.0060	0.0078
	(-4.0674)	(2.4786)	(1.4557)
ΔPCEREAL	-0.0230	0.0199	0.0647
	(-4.6858)	(4.8955)	(7.1729)
ΔPCOOKIES	-0.0072	0.0125	0.0280
	(-3.6892)	(7.6537)	(7.7581)

The first CV focused on the upstream wheat market and has a more precise structural interpretation (as a supply curve) than CV2 or CV3. This is probably because the available information set for the wheat market was relatively more adequate (particularly from inclusion of quantities) than available information sets for the downstream markets that were the focus of CV2 and CV3. Earlier research noted that such quarterly data are generally considered business proprietary and are not in the public domain (BBP, 2004; RBR, 2002; Babula and Rich 2001). As a result, we followed prior research and modeled the downstream markets solely as reduced form price relations. The CV2 and CV3 (equations 12, 13) that focused on these downstream markets characterized by less adequate information sets have less precise, non-structural and reduced form interpretations.

Juselius (2004, p. 175) noted the frequently encountered difficulty in attributing structural economic interpretations to such CVs as our equations 12 and 13:

> "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 and only exceptionally can be given a direct interpretation as an economic equilibrium relation. The reason for this is that a theoretically meaningful relation can be (an often is) a linear combination of several 'irreducible' cointegration relations."

As such, a relation may reflect both the demand and supply elements of a market. Apparently, our CV2 and CV3 that emerged from incomplete downstream market information sets devoid of quantities are two such reduced form relations that include several "irreducible" economic relations that are likely to remain so until more complete information sets are generated (with quantities) and larger samples emerge over time. Nonetheless, our economic analysis below makes progress in illuminating long run structural and reduced form price transmission relationships that error-correct the six U.S. wheat-based markets. And while a structural supply curve emerges from CV1, we leave more complete economic structural interpretations for CV2 and CV3 to future research when more comprehensive information sets and certainly larger samples will be available.

CV1: A U.S. Wheat Supply

Equation 11 or CV1 appears to be a U.S. wheat supply of notable statistical strength. There is a positive and very statistically significant own price elasticity of 5.8. Wheat is one of North America's most researched and litigated farm commodities, and consequently, the literature provides a wide array of empirically estimated or assumed values for North American (U.S. and Canadian) price elasticities of supply. Moreover, the following trade investigations, litigations, events, and/or studies since 1990 related to the U.S./Canadian wheat trade has provided exhaustive reviews of such literature: the U.S. International Trade Commission's or USITC's 2005 reversal of a final 2003 affirmative injury determination that resulted in antidumping and countervailing duties (ADs/CVDs) on selected U.S. imports of Canadian wheat; a 2005 NAFTA Panel remand of the USITC's 2003 injury determination; a set of AD/CVD cases on certain U.S. imports of Canadian wheat that resulted in final AD/CVDs in 2003; two USITC section-332 fact-finding investigations on the U.S./Canadian wheat trade and Canadian Wheat Board trading practices; a 2000 section-301 trade remedy investigation on certain U.S. imports of Canadian wheat; two temporary U.S. tariff rate quotas on selected Canadian durum and non-durum wheat for the year ending September 11, 1995; a special U.S./Canadian Wheat Commission's 1995 study on U.S. imports of Canadian wheat; and a widely-watched USITC section-22 investigation on certain U.S. imports of wheat and wheat products for President Clinton in 1994. These cases and events are summarized collectively by the Canadian Wheat Board (2005), BBP (2004), USITC (2003), U.S. Trade Representative (2000), Glickman and Kantor (1995), and the USITC (1994). Own-price elasticities of North American wheat supply are classified into domestic and trade (export) supply estimates, with the USITC (2003, p. II.21 and II.22) and Alston, Gray, and Sumner (1994) having noted that the export supplies are generally far more price-elastic. Domestic elasticities generally range from 0.3 to 1.0, with selected estimates such as that of Burt and Worthington (1988) reaching as high as 1.9 for longer-run parameters. Well-known estimates within this range include values of 0.6 of Gardiner and Dixit (1987) and 0.7-0.8 of Mielke and Weersink (1990). Estimates for own-price elasticities of North American export supplies range from 5 to 10 (USITC 2003, p. II.23).

We followed BBP (2004), RBR (2002), and Babula and Rich (2001) and defined QWHEAT as a market-clearing quantity based on total usage that includes domestic and traded quantities. And as a result, our U.S. price elasticity of supply in equation 11 should be taken as an average of domestic and export-related price elasticities of supply, and as such, our estimate of 5.8 falls well within the overall range of 0.3-10.0 noted above. Our price elasticity of supply is consequently an average of the less price-elastic domestic and more price-elastic export-related supplies, and is within the literature's general range. Since our estimate emerged from equation 2's error correction component, we follow Juselius (2004) and interpret it as a long run elasticity. Our extensive specification efforts carefully attributed market influences to concurrent events of specific subsamples through a complex array of binary variables restricted to both the long run and short run components of equation 2. After having attributed the market influences to a large array of specific events through Juselius' (2004) and Juselius and Toro's (2005) recommended specification efforts, U.S. policy makers, agribusiness agents, and researchers should note that what remains is an average U.S. long run supply to domestic and world markets that is very responsive to PWHEAT changes but still within the range of estimates in the literature.

The quantity of wheat in CV1 appears negatively but sensitively related to movements in major wheat-based product prices, PFLOUR and PBREAD. Interested agents and researchers should note that the wheat supply also responds to price-affecting events in downstream wheat-based markets, and these events must be factored into any accurate estimation of wheat market impacts of a policy change or other market event. When price levels of wheat flour and bread products fall, demand for flour and bread may rise, and ultimately elicit augmented wheat volumes for use as inputs. These statistically strong influences of flour and bread price behavior on QWHEAT are consistent with BBP's (2004, pp. 16-18) analysis of FEV decompositions that suggested high levels of influence by PFLOUR and PBREAD variation on QWHEAT behavior.

CV2: A Reduced-Form Price Transmission Relationship Normalized on PWHEAT

Equation 12 or CV2 appears to be a reduced form relationship among wheat-related prices reflecting elements of both demand and supply. U.S. wheat price appears positively related to PFLOUR and PBREAD. This positive price transmission coincides with recent analyses of FEV decompositions that suggested that flour and bread price movements are prime determinants of wheat price behavior, and from analysis of impulse response simulations that suggest a positive PWHEAT/PFLOUR relationship (RBR 2002 and BBP 2004, pp. 14-16).

PWHEAT appears negatively related to the value-added manufacturing product prices further downstream for mixes/doughs. As production costs raise PMIXES, perhaps mixes/doughs supply shifts negatively, leading to less wheat ultimately delivered as an input and a fall in PWHEAT. PWHEAT appears positively related to the difference between PCEREAL and PCOOKIES. The price of wheat-based cereals reflects prices of products with generally a lower degree of value-added processing and a higher wheat-related proportion of production costs than the more processed product array represented by PCOOKIES. As demand tightens in the wheat-intensive cereal market relative cookies/crackers market, more wheat is demanded as an input and a higher wheat price may result. We acknowledge that these CV2 price relationships lack straightforward structural interpretations, and likely arose from inadequate information sets. They appear to be what Juselius (2004, p. 175) would consider reduced form combinations of two or more irreducible economic relations that are likely to remain so until downstream market information sets expand to include quantities and samples become larger over time. We relegate more precise interpretations to future research. Nonetheless, our findings correspond with those of BBP (2004, pp. 16-19) that PCEREAL and PCOOKIES importantly influenced PWHEAT.

CV3: A Reduced-Form Price Transmission Relationship Normalized on PFLOUR

CV3 posits flour price as positively related to PWHEAT, with each percentage change in PWHEAT eliciting, on average historically, a 0.49 percent, similarly-directed change in PFLOUR. This reduced form response elasticity coincides closely with prior comparable estimates generated by reduced form VAR impulse response simulations of 0.40 by BBP (2004, p. 14) and RBR (2002, p. 110). U.S. policy makers have an empirical estimate of the pass-through effect of wheat price changes to flour millers downstream. Farmer benefits from rising PWHEAT may burden millers with increased wheat input costs, while PWHEAT declines that burden farmers may benefit flour millers with falling wheat input costs, and by a factor of about half of the percentage change in PWHEAT.

CV3's significant β(PBREAD) and low-valued and insignificant α-estimate suggests that PBREAD influences but is not influenced by this relation’s error correction process. Such results are consistent with BBP (2004, pp. 16-19) findings for these markets: that PBREAD influences the other downstream prices, with little or no feedback from to PBREAD, which may serve as a widely-watched “bell weather” indicator of general bakery market conditions. As well, flour price appears positively related to the difference between PBREAD and prices of wheat-using value added products reflected by PCEREAL. As demand conditions in the widely-watched bread market tighten, PBREAD rises, the PBREAD/PCEREAL wedge widens, and PFLOUR may rise as more flour is demanded. The importance of PWHEAT, PBREAD, and PCEREAL movements in determining PFLOUR behavior coincide with findings by BBP (2004, pp, 16-17).

Analysis of Error-Correction Estimates on the Deterministic Components

Most binary β-estimates in equations 11-13 achieved strong statistical significance. Given that non-binary variables were modeled in logarithms, we used Halvorsen and Palmquist’s (1980) well-known method to convert the binary β-estimates into average percent change effects on the dependent variable from concurrent events associated with the sub-sample for which the binary was defined (hereafter, HP calculated effects.).[50] For space considerations, we focus on the implied effects associated with the AD/CVD case (TITLE7 binary), the NAFTA agreement (NAFTA binary), the sustained increases in confectionary production costs that began in early-2001 (CONFECT binary), and the two temporary U.S. tariff rate quotas imposed on certain imports of Canadian wheat (QUOTA binary).

Effects of the antidumping/countervailing duty case. The filing of the AD/CVD case against certain U.S. imports of Canadian wheat ultimately led to a series of preliminary or final AD/CVD duties on certain imports of Canadian durum and/or hard red spring wheat from 2002/03:02 through the end of the sample (see USITC 2003 for a case summary). The

[50]Considering the β(CONFECT)-estimate of -0.30 in CV1, Halvorsen and Palmquist’s (1980) method takes “e,” the base of the natural logarithm; raises it to the power of the value of the coefficient (here the power of -0.30); subtracts 1.0; and multiplies the result by 100. What results is an average percentage change effect, here –25.9 percent, on the dependent variable (QWHEAT in equation 11). This suggests that the increased confectionary input costs that began in 2001 resulted in about 26 percent less wheat supplied than without CONFECT’s events.

HP calculated effects from TITLE7's β-estimates in CV1-CV3 were rather pronounced: on average, the AD/CVD and related concurrent events resulted in quarterly QWHEAT levels that were 59.7 percent lower; quarterly PWHEAT levels that were 75 percent higher; while quarterly U.S. flour prices were 28.2 percent lower. The preliminary and final tariffs[51] were modest and likely insufficient to have generated such large AD/CVD effects. USITC (2003) and USDA (2005a, b) analyses suggest that other concurrent events such as the tight world grain supplies and high world levels of wheat prices and wheat demand during 2002/03 - 2004/05 could have magnified these HP calculated effects for TITLE7 on QWHEAT and PWHEAT.

An interesting result is the negative TITLE7 β-estimate and -28.2 percent HP calculated effect on CV3's PFLOUR. After U.S. AD/CVD tariffs were imposed on certain U.S. imports of Canadian wheat after 2002/03:02, official trade data from the USITC (2005) clearly reflected a shift from imports of Canadian wheat to imports of wheat flour, which were not covered by the AD/CVD orders. More specifically, average annual U.S. imports of wheat flour quantity for 2002-2004 coinciding with the imposed AD/CVD duties were 50 percent above the 1996-2001 pre-duty average annual imports (USITC 2005). Such sustained increases in U.S. wheat imports associated with the AD/CVD orders' implementation resulted in the perhaps unexpected negative PFLOUR effects. An implication for U.S. implementers of trade remedy cases arises: while statistical evidence suggests that the 2003 AD/CVD cases likely benefited U.S. wheat farmers with higher PWHEAT (CV2), the cases may have unexpectedly resulted in lower flour prices from added wheat flour imports not covered by the dumping and CVD orders.

Effects of the January 1994 implementation of NAFTA. This binary was defined to capture the period from NAFTA's January, 1994 implementation to the end of the sample, a truly "crowded" period when many concurrent events other than NAFTA likely influenced the modeled markets. The U.S. imported steadily increasing volumes of Canadian wheat, while the U.S. government also tapered down levels of farm price supports with implementation of the U.S. farm bill of 1996. There was also a short period during 1994-1995 when world wheat demand, export, and price levels were elevated, followed by a sustained commodity boom during from 2002 through late-2004. NAFTA period events appeared on balance to decrease QWHEAT. As U.S. purchases of Canadian wheat escalated, perhaps the drop in supply price to farmers from concurrent declining levels of wheat program price support and increasing imports led to a drop in U.S. production that was disproportionally larger than the import increase. USDA, ERS (2005a, b) data clearly demonstrate that PWHEAT fell during a substantial period after NAFTA's 1994 implementation, likely reflected in CV2's negative PWHEAT effect.

Effects of the 2001-2002 sustained increases in confectionary production costs. During early 2001, a marked and sustained rise in cocoa prices began in response to disruptions from the Ivory Coast civil war, and in late-2002, there was another run-up in non-cocoa confectionary input costs.[52] CONFECT was defined to capture the effects of these

[51] For the rather complicated array of preliminary and final AD/CVD tariffs imposed on certain U.S. imports of durum and/or hard red spring wheat, see USITC (2003). The final tariffs imposed were on imports of Canadian hard red spring wheat and amounted to just over 14 percent.

[52] This analysis and information was compiled by a U.S. International Trade Commission industry analyst responsible for monitoring markets for sugar and confectionary products, in two emails received by an author on August 18 and 19, 2004. A more in-depth analysis on the effects on U.S. sugar-based product markets of

confectionary input cost increases (and other relevant concurrent events) for wheat-based confectionary products included in PMIXES and PCOOKIES. The HP calculated effects for CONFECT suggested that the input cost increases and other concurrent events resulted in 26-percent lower QWHEAT levels and a 42-percent lower PWHEAT. The calculated HP effect of CV3's β(CONFECT) estimate suggests that flour price was on average 21 percent above would-be levels without the sustained increases in confectionary input costs. Confectioners may have substituted flour for other increasingly expensive inputs, thereby augmenting PFLOUR. The implication is that U.S. policy makers should factor-in cross-market effects to avoid inadvertent augmenting or offsetting estimated effects of chosen policy changes, no matter how far-off or indirect such events as the Ivory Coast war may initially appear. Our statistical evidence strongly suggests that this war and other events' influences on confectionary costs had important cross-market influences on the U.S. wheat market.

Effects associated with two temporary U.S. tariff rate quotas on Canadian-sourced wheat. Two U.S. TRQs were imposed on certain U.S. imports of Canadian wheat for the year ending September 11, 1995 (Glickman and Kantor 1995). As expected, events associated with QUOTA suggested negative impacts on QWHEAT in CV1 as imports were restricted, and on PFLOUR as importers shifted to importing more wheat flour not covered by the tariff rate quotas. The HP calculated effects were -51 percent for QWHEAT in CV1 and -28 percent on PFLOUR in CV3. The following other concurrent events may account for the magnified effect estimates: the 1995 start of a commodity boom with high world demand levels for wheat, and possible expectationary market influences of the then-anticipated 1996 U.S. farm bill which noticeably lessened the U.S. wheat program support levels (among other events).

CV3 incorporates the statistically supported restriction rewritten as: -0.33*(QUOTA + TITLE7 - NAFTA). This restriction suggests to policy makers that events associated (i) with the TRQs and the set of AD/CVD cases had similar and decreasing effects on PFLOUR as imports shifted from wheat to wheat flour, and (ii) with NAFTA raised PFLOUR by about as much as the events associated with each of the two trade remedies decreased it.

SUMMARY AND CONCLUSIONS

RBR (2002) and BBP (2004) applied the largely reduced form and nonstructural VAR econometric modelling methods collectively developed and introduced to U.S. agriculture by Sims (1980) and Bessler (1984) to a quarterly set of U.S. markets for wheat, flour, mixes/doughs, bread, wheat-based breakfast cereals, and cookies/crackers. Johansen (1988) and Johansen and Juselius (1990, 1992) developed a well-known set of extensions to such reduced form methods to permit structural relationships to emerge through exploitation of the system's cointegration properties. These extensions have been further refined by Juselius (2004) and Juselius and Toro (2005). For perhaps the first time, we applied such cointegrated VAR methods to these same markets. We exploited the modelled system's cointegration properties, and incorporated a wide array of binary variables to capture empirically estimated effects of important market/institutional events. Results illuminated a rich long run error

these two run-ups in confectionary input costs is provided in Babula and Newman (2005). Given that many confectionary products, both cocoa-based and non-cocoa, use wheat, we included CONFECT in our analysis, and with clear statistical support as seen from our results.

correction space which provided structural and reduced form estimates on how these markets run and interact. Parameter estimates from three CV s emerged – a U.S. wheat supply curve and two transmission relationships among U.S. wheat-based prices, along with estimated effects from important market/institutional events.

The first CV, a wheat supply curve, suggested that in the very long run, market-clearing wheat quantities are highly and positively related to changes in PWHEAT. With QWHEAT defined to include domestic and traded quantities, our 5.8 price elasticity of supply reflects an average of less elastic domestic and more elastic export supplies, and falls within the general range of the literature's estimates. Perhaps when adequate specification effort is made to fully capture market effects of important economic and institutional changes, average U.S. domestic and export wheat supplies is more price-elastic than initially thought. Future research can assess whether after such specification effort, the elasticity estimate's central location within a very wide literature range of estimates reflects a more accurate estimate, with influences of other events appropriately attributed through binaries. Wheat supply was also found highly responsive to price-impacting events in downstream wheat-related markets.

CV2 and CV3 estimates provided statistically strong signals to policy makers that wheat prices are highly influenced by market conditions downstream. When U.S. policy makers, agribusiness agents, or researchers estimate PWHEAT impacts of a policy change (altered loan rate, e.g.) or another event, such estimates need to concurrently consider clearly important cross market impacts in order to arrive at accurate estimates. In CV3, flour price appears positively related to PWHEAT, with each percentage change in PWHEAT eliciting, on average historically, similarly directed changes of 0.49 percent in PFLOUR – a response that closely corresponds to previous estimates by closely relevant prior research (BBP 2004, p. 14; RBR 2002, p. 110). This appears to be an important empirical estimate of downstream flour price effects for U.S. policy makers, agribusiness agents, and/or researchers when pondering implications for a policy change or event impact for the wheat market.

We provided a number of empirically estimated market effects associated with the AD/CVD case (TITLE7 binary), NAFTA agreement implementation (NAFTA binary), 2001 sustained rises in confectionary production costs (CONFECT binary), and the two TRQs imposed on U.S. certain imports of Canadian wheat (QUOTA binary). While the AD/CVD case and concurrent events resulted, as perhaps expected, in QWHEAT declines and in PWHEAT increases, effects on PFLOUR were (perhaps unexpectedly) negative as importers shifted towards imports of wheat flour not covered by the AD/CVD orders. When assessing overall impacts of an AD/CVD remedy on behalf of farmers, such downstream cost on flour millers should be considered. NAFTA's coefficients suggested negative effects on QWHEAT and PWHEAT as the U.S. wheat markets were opened to the Canadian exporters [primarily the Canadian Wheat Board] (see USITC 1994, chapter II). The 2001-2002 sustained rises in confectionary input prices appeared to have had statistically strong cross-market impacts on U.S. wheat and flour markets – indications on the importance of seemingly far-off events such as an Ivory Coast war to U.S. policy makers, agribusiness agents, and researchers. The temporary U.S. tariff rate quotas on U.S. imports of Canadian wheat appeared to restrict QWHEAT through impeded imports, while PFLOUR fell as importers switched to imports of wheat flour not covered by the quotas.

U.S. policy makers, agribusiness agents, and researchers may find the following empirical implications of our results of interest: (1) that NAFTA and the AD/CVD cases clearly reduced QWHEAT in U.S. markets; (ii) that NAFTA decreased and the AD/CVD cases raised PWHEAT for U.S. farmers; and (iii) that tariff rate quotas and AD/CVD cases successfully raised wheat prices for farmers, but had equally depressing effects on PFLOUR for millers downstream.

Our goal was to provide, for agricultural economists, a cogent application of a procedural methodology for exploitation of cointegration properties that is currently evolving in the econometrics literature and summarized by Juselius (2004). To our knowledge, our procedure provides early or first-time applications in the agricultural economics literature of a systems-based and rank-dependent stationarity test, a method for testing for and choosing cointegration space rank using several sources of evidence, a new test for structural change, and a recently developed test for the presence of I(2) trends in modeled data.

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