Impacts of Gluten Imports on U.S. Food Wheat Use

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Introduction

Wheat gluten often is used as a substitute for high-protein wheat flour during the baking process (Holcomb, 2000). Bakers and processors' demand for imported wheat gluten increases in periods when the proportion of protein in the wheat crop is low, because of weather or other reasons (Boland *et al*. 2000 and 2005). Recent increases of wheat gluten imported into the US have lead to trade disputes (Balzer et al, 1999; Stiegert et al , 2001).[1] As a result, understanding demand for wheat food use by class could play an important role in reconciling US agricultural trade disputes and future trade negotiations.

In 1998, based on a petition filed by the Wheat Gluten Industry Council (WGIC), the U.S. International Trade Commission (USITC) charged the EU with dumping wheat gluten on the U.S. market. According to the WGIC the price of the EU gluten during period 1993 to 1996 was about $0.04 per pound lower than domestic gluten, which produced negative impacts on the U.S. gluten industry. In addition, a report issued by the USITC (1998) indicated that the EU's share of U.S total gluten imports was only two percent in 1985, but it had increased to 51 per cent by 1997. As a result of the USITC ruling, a three-year quota was approved on wheat gluten imports from Australia, the EU,

[1] Imports of gluten into the U.S. have dramatically increased 5.23 million bushels by 2003 from 1.79 million bushels in 1990. Researchers have argued that this has occurred for two main reasons. First, American consumer preference for "healthy" flour-based products has increased (Holcomb, 2000). Second, according to Balzer et al. (1999) and Stiegert et al. (2001), gluten suppliers in EU have attained government support from subsidizing wheat starch or gluten processing, and thus they can supply gluten at a lower price in both U.S. and world markets (which suggest that the U.S. millers and bakers are more likely to combine import gluten with lower protein wheat to produce flour for baked goods that require higher levels of protein).

and all other non-excluded countries on June 1, 1998.

Various studies have examined domestic wheat demand. Chai (1972) estimated domestic demand for wheat by class over the period from 1929 to 1963. Linear equation-by-equation OLS demand models were estimated using wheat cash prices from major markets. Price elasticities were reported to be more elastic for hard classes than soft classes of wheat. Barnes and Shields (1998) estimated a double-log demand system for wheat by class. Annual data from 1981 to 1998 were used in a demand system analysis with regional prices at the farm level. Inelastic own-price elasticities were reported for each of the five wheat classes with soft white wheat being the most elastic and durum wheat being the least elastic. Barnes and Shields (1998) also estimated linear equation-by-equation OLS models that yielded results qualitatively consistent with Chai (1972). More recently, Marsh (2005) estimated wheat demand by class using a production theory approach from 1975 to 2002. He found price elasticies more elastic for hard as opposed soft what classes.

Two studies have examined empirical relationships between wheat gluten and wheat demand. Ortalo-Magne and Goodwin (1992) developed a demand model for gluten imports, reporting that the demand for wheat gluten in the U.S. is positively related to the price of high protein wheat. Stiegert and Balzer (2001) modeled demand and supply for wheat gluten and intrinsic wheat protein for HRW and HRS. Their main conclusions suggested a strong influence of the hard red winter wheat on wheat gluten market. Also wheat gluten markets were reported to have significant impact on protein premium for hard red spring wheat but not on protein premium for hard red winter wheat.

Our interest is to determine if gluten imports into the U.S. impacted markets (quantities demanded or prices) for wheat food use by class (i.e., hard red spring, hard red winter, soft red, soft white, and durum wheat) during the 1990s. To do so, we specify both a standard demand system, from which we examine associations between gluten imports and quantity of wheat demanded, and an inverse demand system, from which we investigate associations of gluten imports and price formation. To address issues of model specification, we apply a nonnested generalized likelihood ratio test procedure to determine whether the data are consistent with quantity formation or price formation.

The remainder of this paper is organized in the following manner. First, we present our conceptual methodology. Second, data description and estimation issues are presented. Third, we provide a discussion of results. Finally, we finish with some conclusion remarks.

Methodology

The methodology section proceeds in the following manner. An indirect profit function is specified from which to derive a factor demand system for wheat food use. Next, an input distance function is specified to derive inverse demand system, with which to examine price formation across wheat classes. Then we discuss an approach to perform a nonnested test for model selection, determining if prices are adjusting to quantities or quantities are adjusting to prices in the U.S. wheat food use markets.

Profit Function Approach:

Following Marsh (2005), consider an indirect industry profit function of the flour milling industry specified as,

$$\Pi^l(\mathbf{p},\mathbf{w}) = \Pi^l(\mathbf{y}^l(\mathbf{p},\mathbf{w}),\mathbf{x}^l(\mathbf{p},\mathbf{w})) = \max_{\mathbf{y}^l,\mathbf{x}^l}\{\mathbf{w}'\mathbf{y}^l - \mathbf{p}'\mathbf{x}^l : \mathbf{y}^l = f^l(\mathbf{x}^l)\} \qquad (1)$$

where $\Pi^l(\bullet)$ and $f^l(\bullet)$ represent the l^{th} firm's profit and production technology respectively, $\mathbf{p} = (p_1, ..., p_m)'$ and $\mathbf{w} = (w_1, ..., w_n)'$ are prices of m inputs and n outputs respectively, $\mathbf{y}^l = (y_1^l, ..., y_n^l)$ is a $n \times 1$ vector of the l^{th} firm output quantities, and $\mathbf{x}^l = (x_1^l, ..., x_m^l)$ represents a $m \times 1$ vector of its input quantities. Assume all existing milling firms are price-taker in both input and output markets, then industry profit function is

$$\Pi(\mathbf{p}, \mathbf{w}) = \sum_{l=1}^{L} \Pi^l(\mathbf{p}, \mathbf{w}) \tag{2}$$

where L is total number of flour milling firms in the industry. Similarly, the industry input and output quantities can be derived respectively as,

$$\mathbf{x} = (x_1, ..., x_m : x_i = \sum_{l=1}^{L} x_i^l, i = 1, ..., m) \text{ and } \mathbf{y} = (y_1, ..., y_n : y_j = \sum_{l=1}^{L} y_j^l, j = 1, ..., n)$$

Assuming weak separability we separate inputs into two subgroups of wheat and other inputs.[2] Hence, the industry profit function is

$$\Pi = \Pi(\mathbf{p}, \mathbf{w}) = \Pi(\pi^1(\mathbf{p}, \mathbf{w}^1), \pi^2(\mathbf{p}, \mathbf{w}^2), \mathbf{p}) \tag{3}$$

where π^1 and π^2 are micro-function, $\mathbf{p}^1 = (p_1, ..., p_k)'$ is a vector of input prices representing the different classes of wheat, and $\mathbf{p}^2 = (p_{k+1}, ..., p_m)'$ is a vector of prices for remaining inputs. By applying Hotelling's Lemma to function π^1, factor demand for wheat by class can be derived as,

$$-\mathbf{x}^1(\mathbf{p}, \mathbf{w}^1) = \frac{\partial \pi^1}{\partial \mathbf{p}^1} \tag{4}$$

[2] This assumption imposes symmetric factor demand elasticities between two groups of inputs (Chambers, 1998).

This system of factor demand equations represent the flour miller's demand for wheat by class from the producer supplier.

Empirically, we specify a normalized quadratic profit function

$$\begin{aligned}\Pi^*(\tilde{\mathbf{p}}, G) &= a_0 + \sum_{i=1}^{k} c_i \tilde{p}_i + .5\left(\sum_{i=1}^{k}\sum_{j=1}^{k} b_{ij}\tilde{p}_i\tilde{p}_j\right) + t_0 T + \sum_{i=1}^{k} t_i \tilde{p}_i T + .5 t_{ii} T^2 \\ &+ \sum_{j=1}^{3} d_{0j} D_j + \sum_{i=1}^{k}\sum_{j=1}^{3} d_{ij}\tilde{p}_i D_j + \sum_{i=1}^{k} g_i \tilde{p}_i G + .5 g_{ii} G^2\end{aligned} \quad (5)$$

where $\Pi^*(\bullet)$ represents normalized profit, which is obtained by dividing profit $\Pi(\bullet)$ by a weighted average flour price $\bar{w} = \sum_{i=1}^{n} s_{ij} w_i$, $\tilde{p}_i = p_i / \bar{w}$ is a vector of input prices normalized by output price, *G* represents gluten imports into the U.S., *T* is a time trend that is used to capture technology progress and other changes over time, and $D_j\ (j = 1,2,3)$ are quarterly dummy variables. Here $c_i, b_{ij}, t, d_{ij}, g_i$ are parameters to be estimated. By equation (4), the demand equation for each class of wheat is then

$$-x_i = c_i + \sum_{j=1}^{k} b_{ij}\tilde{p}_j + t_i T + \sum_{j=1}^{3} d_{ij} D_j + g_i G \quad (6)$$

Distance Function Approach

A direct input distance function for the flour milling sector is defined from which we derive an inverse factor demand system. Classical duality theory suggests that the distance function approach is consistent to the cost minimization assumption or profit function approach. The standard properties of a distance function are that it is homogenous of degree one, non-decreasing, and concave in input quantities and non-increasing in outputs (Shepherd 1970). Define the distance function as

$$\begin{aligned} D(\mathbf{x}, \mathbf{y}) &= \max_{\delta} \delta \\ s.t.\ & f(\mathbf{x}/\delta) = \mathbf{y} \end{aligned} \quad (7)$$

where $\mathbf{y}$ and $\mathbf{x}$ are defined above, and $\delta \geq 1$ is the distance function representing a rescaling of all the input levels consistent with a target output level. Intuitively, δ is the maximum value by which one could divide $\mathbf{x}$ and still produce $\mathbf{y}$. Normalizing the price vector of inputs by total cost yields $p_i^* = p_i / \sum_{j=1}^{n} p_j x_j$. Applying Gorman's Lemma, inverse factor demand functions are

$$p_i^*(\mathbf{x}, \mathbf{y}) = \frac{\partial D(\mathbf{x}, \mathbf{y})}{\partial x_i} \tag{8}$$

Following Marsh and Featherstone (2003), the specified input distance function in (7) can be specified as a normalized quadratic

$$\begin{aligned} &D(\mathrm{x,y}) = \\ &b_0 + \sum_{i=1}^{n} b_i x_i + \sum_{i=n+1}^{n+m} b_i y_i + .5((\sum_{k=1}^{n} \alpha_k x_k)^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n} b_{ij} x_i x_j + \sum_{i=n+1}^{n+m} \sum_{j=n+1}^{n+m} b_{ij} y_i y_j) + \sum_{i=1}^{n} \sum_{j=n+1}^{n+m} b_{ij} x_i y_j \\ &+ g_0 G + .5 g_{00} G^2 + \sum_{i=1}^{n} g_i x_i G \\ &+ \sum_{j=1}^{3} d_j D_j + .5 \sum_{j=1}^{3} d_j D_j^2 + \sum_{i=1}^{n} \sum_{j=1}^{3} d_{ij}^* x_i D_j \\ &+ t_0 T + .5 t_{00} T^2 + \sum_{i=1}^{n} t_i x_i T \end{aligned} \tag{9}$$

with n inputs and m outputs; G, D_j (j=1,2,3), and T have the same definitions as above; the $b's$, $g's$, $d's$, and $t's$ are parameters to be estimated; and e the α_i are predetermined positive constants that dictate the form of normalization. Symmetry is imposed by restriction $b_{ij} = b_{ji}$. Using the Gorman's Lemma, the conditional inverse factor demand functions can be given by

$$p_i^* = b_i + (\sum_{k=1}^{n} \alpha_k x_k)^{-1} \sum_{j=1}^{n} b_{ij} x_j + \alpha_i (\sum_{k=1}^{n} \alpha_k x_k)^{-2} \sum_{i=1}^{n} \sum_{j=1}^{n} b_{ij} x_i x_j) + \sum_{j=n+1}^{n+m} b_{ij} y_j + g_i G + \sum_{j=1}^{3} d_{ij} D_j + t_i T \qquad (10)$$

Homogeneity of degree zero in inputs in the inverse factor demand equation implies that $\sum_{j=1}^{n} b_{ij} = 0$, while the normalization restriction requires that $\sum_{k=1}^{n} \alpha_k x_k = 1$ at a reference vector. Normalizing quantities by their mean values yields $x^* = (1,...,1)' = l_n$, which can be used as a reference bundle. At a reference vector x^*, the demand restrictions become

$$\sum_{k=1}^{n} \alpha_k x_k^* = \sum_{k=1}^{n} \alpha_k = 1, \alpha_k \geq 0, \forall k, \text{ and } \sum_{j=1}^{n} x_j^* b_{ij} = \sum_{j=1}^{n} b_{ij} = 0 \qquad (11)$$

To impose these restrictions, we normalized factor demand quantities by the k^{th} input as $x_s^* = x_s / x_k \ \forall s = 1,...,n$, and predetermined constants as $\alpha = (0,...,0,\alpha_k,0,...,0) \ni \alpha_k = 1$ such that $\sum_{s=1}^{n} \alpha_s x_s^* = 1$. Hence, the input demand functions in (10) become

$$p_i^* = b_i^* + \sum_{j=1}^{n-1} b_{ij}^* x_j^* + \sum_{j=n+1}^{n+m} b_{ij}^* y_j + g_i^* G + \sum_{j=1}^{3} d_{ij}^* D_j + t_i^* T + \varepsilon_i \quad \text{for } i = 1,...,n-1 \qquad (12)$$

with stochastic error terms ε_i.

Model Selection and Nonnested Test

To select between the two dual competing models discussed above, we apply nonnested test proposed by Vuong (1989, 1992). In specifying the nonnested test we derive share equations as alternatives to the demand system in (6) and to the inverse demand system in (12).[3] First, we construct share equations for the factor demand system. Multiplying both sides of the equation (6) by normalized price $p_i^* = p_i / \sum_{i=1}^{k} p_i x_i$, we get the share equations

$$-w_i = -p_i^* x_i = p_i^* c_i + p_i^* \sum_{j=1}^{k} b_{ij} \tilde{p}_j + p_i^* t_i T + p_i^* \sum_{j=1}^{3} d_{ij} D_j + p_i^* g_i G \qquad (13)$$

Second, we derive the share equations for the inverse factor demand system from equation (12) by multiplying both sides of the equation by the corresponding input quantity x_i

$$w_i = x_i p_i^* = x_i b_i^* + x_i \sum_{j=1}^{n-1} b_{ij}^* x_j^* + x_i \sum_{j=n+1}^{n+m} b_{ij}^* y_j + x_i g_i^* G + x_i \sum_{j=1}^{3} d_{ij}^* D_j + x_i t_i^* T + \varepsilon_i \qquad (14)$$

Using the two systems of share equations, we followed the proposed nonnested normalized likelihood ratio (LR) test by Vuong (1989, 1992) to determine a preferred model in pairwise evaluation. The test is based on the generalized likelihood ratio principle and is designed to test the null hypothesis that two dual models adjust to the data equally well versus the alternative hypothesis that one model fits better. We calculate the likelihood ratio statistic adjusted for the difference in the number of the estimated parameters and normalized by

$$n^{\frac{1}{2}} \hat{w}_n = \frac{1}{2} \left[\sum_{t=1}^{n} (\hat{\boldsymbol{\mu}}_{it}^{'} \sum_i^{-1} \hat{\boldsymbol{\mu}}_{it} - \hat{\boldsymbol{\mu}}_{dt}^{'} \sum_d^{-1} \hat{\boldsymbol{\mu}}_{dt})^2 \right]^{\frac{1}{2}} \qquad (15)$$

where $\boldsymbol{\mu}_s$ are the estimated residuals and $\sum_s$ are the estimated covariance matrix for

[3] Part of the motivation for specifying share equations is that in preliminary analysis we applied standard single equation nonnested tests between the competing models. However, these tests were inconclusive and we turned the generalized likelihood ratio test.

model M_s, $s= i, d$. i stands for inverse demand system, d stands for factor demand system. The resulting normalized statistic is asymptotically normally distributed under the null hypothesis of equal fit. When the absolute value of the normalized LR statistic is smaller than the critical value, then the data cannot identify a superior model. If the normalized LR statistic is smaller than the negative critical value, then we conclude that factor demand model is preferred; and if it is greater than the critical value, then we conclude that inverse model is preferred.

Data Description

There are five major classes of wheat grown in the U.S. for food consumption, including hard red winter (*HRW*), hard red spring (*HRS*), soft red winter (*SRW*), soft white (*SWW*), and durum (*DUR*) wheat. Quarterly prices and quantities from USDA-ERS are used for empirical analysis. Data period ranges from the first quarter of 1990 to the fourth quarter of 2003. Table 1 reports the descriptive statistics for quarterly prices and quantities. Total flour production reached to near 400 million cwt in 2003 from 343 million cwt in1990, averaging 392 million cwt. In same period, total wheat for food use has increased from 749 million bushels to 918 million bushels. Of five classes of wheat, the food uses of HRW and HRS have rapidly increased but other three classes SRW, SWW and DUR only have slightly shifted up (Figure 1). Averagely, the hard wheat (HRW and HRS) accounts for 76 per cent out of total food wheat use, while the percentages for SRW, SWW and DUR are 18.1 per cent, 7.69 per cent and 7.92 per cent, respectively.

Wheat quantity and price data used in this paper were from Economic Research Service, the U.S. Department of Agriculture (ERS, USDA). The original price data for classes of wheat were from four major markets. HRW price is represented by Kansas City, No.1 (13% protein); HRS price and DUR price are represented by Minneapolis, dark No.1 spring (14% protein) and No.1 hard amber durum respectively; SRW price by Chicago, No.2; SWW price by Portland No.1. Figure 2 shows these prices trend over our studying period. At the beginning of 1990's, they were quiet close, but gradually, year after year, these prices were far away from each other. By the end of our studying period, the price of DUR was about one and half time of SRW.

Quarterly imports of wheat gluten from World Trade Atlas increase over the studying period (Figure 3). Total imports reached 4.13 million bushels in 2003 from 1.64 million bushels in 1990. Meanwhile, it is apparent that they varied significantly across quarters, in particular during the three-years quota policy started June 1, 1998.

Empirical Results

Applying the nonnested test for model selection (Vuong 1989, 1992), the statistical test value based on the normalized LR statistics from formula (15) is 390.51. This is greater than any relevant critical values from the standard normal distribution. In all this provides strong evidence that the inverse system is preferred to the factor demand system, given the data and model specifications. Based on these results, we present results for the inverse demand system to investigate further the associations between gluten imports and price formation by wheat class.

The inverse demand system in (12) was estimated using iterative seemingly unrelated regression (SUR) estimator with AR(1) autocorrelation correction, concavity condition, and symmetry imposed. A bootstrap resampling procedure was used to form confidence intervals. The dummy variable for the fourth marketing quarter (March to May) was dropped in regression to prevent perfect multicollinearity. The model with the last two-quarter cumulative quantity for the gluten import was chosen because it offered the highest likelihood value than other cases. The estimated results are presented in Table 2. More than half of estimated coefficients are statistically significant at 10% level. The R-square values for equations HRS, HRW, SRW, and DUR are 0.91, 0.81, 0.90, and 0.64 respectively.

As expected, the coefficients for the own-quantity demanded are negative and statistically significant at 0.10-level for each wheat class. In terms of quarter dummy variables, four out of twelve are significantly different from the fourth quarter, but the estimated results are not consistent across both wheat classes and quarters. At the bottom of Table 2, we observe that the estimated time trend coefficient are significantly negative for SRW, meaning the price of SRW has decreased over the study period. This trend was not found for other three classes.

The estimated coefficients and confidence intervals also show that the gluten import quantity is significantly related to prices in three out of four wheat classes. Basically, gluten imports are associated negatively with prices of the high-protein wheat classes and positively with prices of lower protein wheat classes. This result is consistent with discussion in previous studies (e.g. Boland et al., 2000; Holcomb, 2000), meaning that the gluten may be a substitute relationship with higher-protein wheat classes.

Based on the estimated parameters for the four classes of wheat and the parameters for SWW equation that were recovered by the imposed restrictions, the price flexibilities at the mean for each inverse demand equation were calculated and reported in Table 3.[4] The confidence intervals for price flexibilities estimated by bootstrap resampling procedure are also reported in this table. As expected that all five of own-flexibilities are significantly negative as required with the imposition of concavity condition for each wheat class, but only the cross-price flexibilities between SWW and two high-protein wheat classes HRS and DUR are significant. Also, from the owned-price flexibilities, we notice that prices of two soft wheat classes are more responsive to owned-quantity changes than are hard wheat classes. Compared to Marsh (2003) where annual data were used, and without gluten imports included, all estimated price flexibilities in the current study show the same signs but slightly different magnitudes.

The price flexibilities for gluten imports at the mean for each inverse demand equation were calculated and reported in Table 4.[5] All three of the higher-protein wheat classes (HRW, HRS, DUR) have negative price-flexibilities with respect to gluten imports. HRW and DUR have significantly negative price-flexibilities while two soft wheat classes have significantly positive signs. Meanwhile, it is noticeable that SRW, DUR, and SWW exhibit price flexible with significant larger magnitudes (0.078, 0.094, and 0.082 respectively) than two hard wheat classes, HRW (0.018) and HRS (0.006), implying that the first three wheat classes had more sensitive responses to gluten imports than hard wheat classes.

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[4] The compensated price flexibilities are given by $f_{ij}^{*} = \dfrac{\partial \ln p_i}{\partial \ln x_j} = \dfrac{b_{ij} x_j}{\hat{p}_i}$ for $i, j = 1, \ldots, n$, using the estimated b_{ij} and the predicted $\hat{p}_i$. The price flexibilities with respect to the gluten import quantity are given by $\xi_i = \dfrac{\partial \ln p_i}{\partial \ln G} = \dfrac{g_i G}{\hat{p}_i}$ for $i = 1, \ldots, n$.

[5] Formulas for price flexibilities are presented in table footnotes.

Conclusion and Discussion

We examined the impact of wheat gluten on the markets for wheat food use in the U.S. Using quarterly wheat food use by class price and quantity data, we conceptualized and specified an inverse demand system and factor demand system for five classes of wheat. Applying a nonnested generalized likelihood ratio test, we rejected the factor demand system in favor of the inverse demand system. This suggests that prices were adjusting to quantities over the sample period.

Results from the inverse demand system using quarterly data over the 1990s indicate that: (1) durum (DUR), soft red winter (SRW), and soft white wheat (SWW) are more own-price flexible than hard red winter (HRW) and hard spring (HRS) wheat; (2) own-price and cross-quantity effects among five class of wheat are inflexible; (3) all three higher-protein wheat classes (HRW, HRS, DUR) were negatively associated with gluten imports, (4) the lower protein wheat classes (SWW and SRW) were positively associated with gluten imports, and (5) DUR, SRW and SWW exhibit more price flexiblity to gluten imports than do the other two classes.

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Table 1. Descriptive Statistics for Price and Quantity Data from 1990.1-2003.4

Variables	Mean	St. Dev.	Min	Max
Annual Quantity of Flour (1000 cwt)	392,040	18,490	354,350	421,270
Quarterly Quantity of Flour (1000 cwt)	98,164	5,956	85,692	112,240
Price of Flour ($/cwt)	10.38	1.56	7.53	15.18
Price of Hard Red Winter ($/bu)	3.96	0.75	2.86	6.51
Price of Hard Red Spring ($/bu)	4.17	0.76	2.83	6.49
Price of Soft Red Wheat ($/bu)	3.21	0.74	2.02	5.40
Price of Soft White Wheat ($/bu)	3.79	0.70	2.77	5.68
Price of Durum ($/bu)	4.96	1.11	3.10	7.08
Quantity of Hard Red Winter (million bu)	90.29	7.79	75.65	105.39
Quantity of Hard Red Spring (million bu)	56.11	7.19	40.00	70.00
Quantity of Soft Red Wheat (million bu)	37.84	2.53	33.00	44.93
Quantity of Soft White Wheat (million bu)	17.93	2.59	12.00	23.60
Quantity of Durum (million bu)	18.80	1.68	14.83	21.70
Quantity of Imported Gluten (million bu)	0.74	0.33	0.41	1.64

Table 2. Estimated Results from Bootstrap Resampling Procedure

	HRW	HRS	SRW	DUR
Constant	0.009053*	0.008079*	0.012530*	0.004853*
	(0.000669)	(0.000957)	(0.001396)	(0.002832)
HRW	-0.000021*	0.000036	0.000005	0.000047
	(2.763298)	(19.486031)	(21.452480)	(39.558581)
HRS		-0.000149*	0.000087	-0.000176
		(1.870224)	(3.212234)	(5.345885)
SRW			-0.000231*	0.000216
			(1.970346)	(6.180815)
DUR				-0.000706*
				(28.484673)
Flour	-0.004524*	-0.002832*	-0.009090*	0.001828
	(0.000683)	(0.000976)	(0.001404)	(0.002854)
Gluten	-0.000059*	-0.000019	0.000179*	-0.000334*
	(0.000047)	(0.000068)	(0.000093)	(0.000197)
1st quarter	-0.000013	0.000017	0.000042	-0.000231*
	(0.000041)	(0.000058)	(0.000083)	(0.000178)
2nd quarter	-0.00001	-0.000187*	0.000384*	-0.000485*
	(0.000059)	(0.000085)	(0.000123)	(0.000249)

(next)

Table 2. Estimated Results from Bootstrap Resampling Procedure *(cont.)*

	HRW	HRS	SRW	DUR
3rd quarter	0.000013	-0.000036	-0.000089	0.000219
	(0.00005)	(0.000072)	(0.000099)	(0.000209)
Time trend	-0.000002	-0.000005	-0.00001*	0.000008
	(0.000003)	(0.000004)	(0.000005)	(0.000012)
ρ	0.268788*			
	(0.065298)			

* Means statistically significant at 10% significant level.

Table 3. Price Flexibilities for Wheat Food Use by Class and Confidence Intervals[a, b]

Price	HRW	HRS	SRW	DUR	SWW
HRW	-0.02348*	0.02486	0.00239	0.01097	-0.01486
HRS	0.03857	-0.09729*	0.03863	-0.03894	0.04231*
SRW	0.00710	0.07361	-0.13350*	0.06200	-0.02101
DUR	0.04263	-0.09760	0.08159	-0.13179*	0.10941*
SWW	-0.07898	0.14501*	-0.03780	0.14959*	-0.15486*
	90% confidence Intervals-Lower				
HRW	-0.05805	-0.00234	-0.02314	-0.00923	-0.03851
HRS	-0.00361	-0.17466	-0.01279	-0.07980	0.00958
SRW	-0.06762	-0.02489	-0.30700	-0.02957	-0.08838
DUR	-0.03540	-0.20007	-0.03881	-0.25436	0.01390
SWW	-0.20455	0.03282	-0.15930	0.01906	-0.29515
	90% confidence Intervals-Upper				
HRW	-0.00236	0.04936	0.03200	0.03291	0.00371
HRS	0.07697	-0.03044	0.09949	0.00113	0.08037
SRW	0.09324	0.18986	-0.02097	0.15031	0.05277
DUR	0.12753	0.00281	0.19514	-0.03123	0.20467
SWW	0.01974	0.27594	0.09565	0.27945	-0.03694

[a] * means statistically significant at 10% significant level.

[b] Price flexibilities are given by $f_{ij}^{*} = \frac{\partial \ln p_i}{\partial \ln x_j} = \frac{b_{ij} x_j}{\hat{p}_i}$ for $i, j = 1,...,n$, using the estimated b_{ij} and the predicted $\hat{p}_i$.

Table 4. Price Flexibilities for Gluten Imports[a, b]

Prices	Gluten Imports	90% Confidence Intervals	
HRW	-0.01837*	-0.03721	-0.00062
HRS	-0.00640	-0.03768	0.01672
SRW	0.07772*	0.02396	0.12181
DUR	-0.09423*	-0.16409	-0.02366
SWW	0.08154*	0.00322	0.14504

[a] * means statistically significant at 10% significant level.

[b] price flexibilities with respect to gluten import quantity are given by $\xi_i = \frac{\partial \ln p_i}{\partial \ln G} = \frac{g_i G}{\hat{p}_i}$ for $i = 1,\ldots,n$.

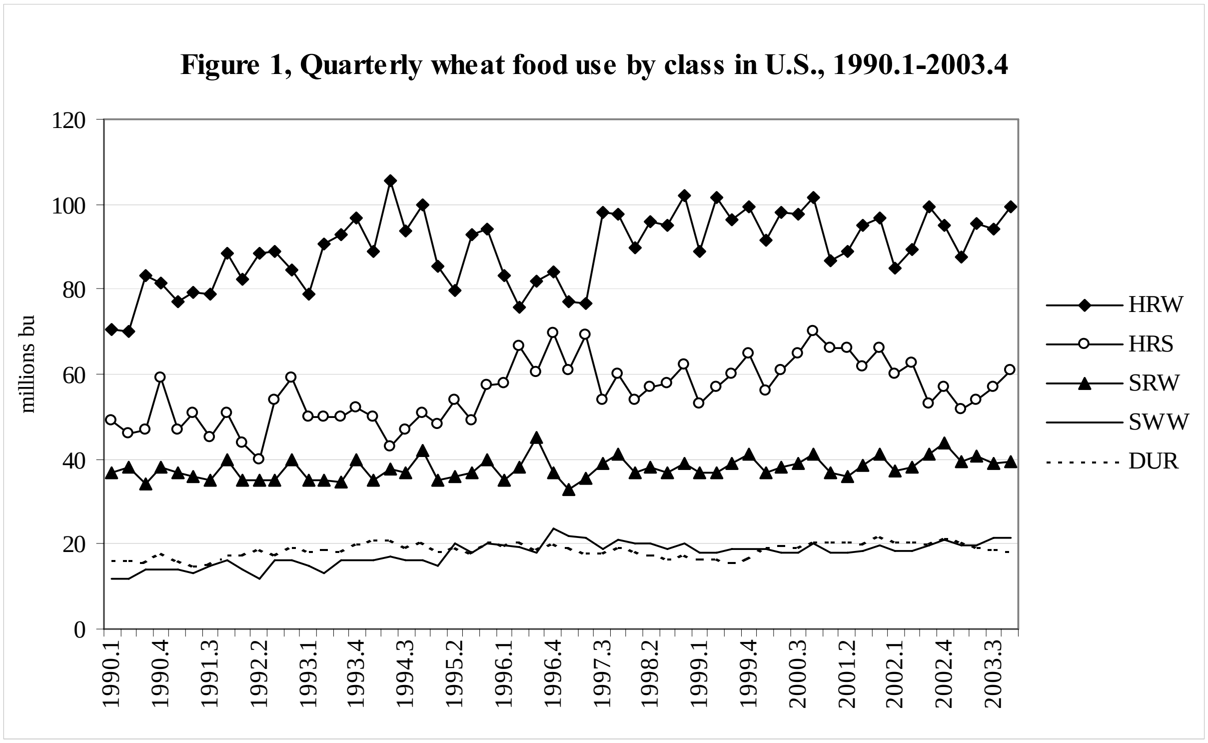
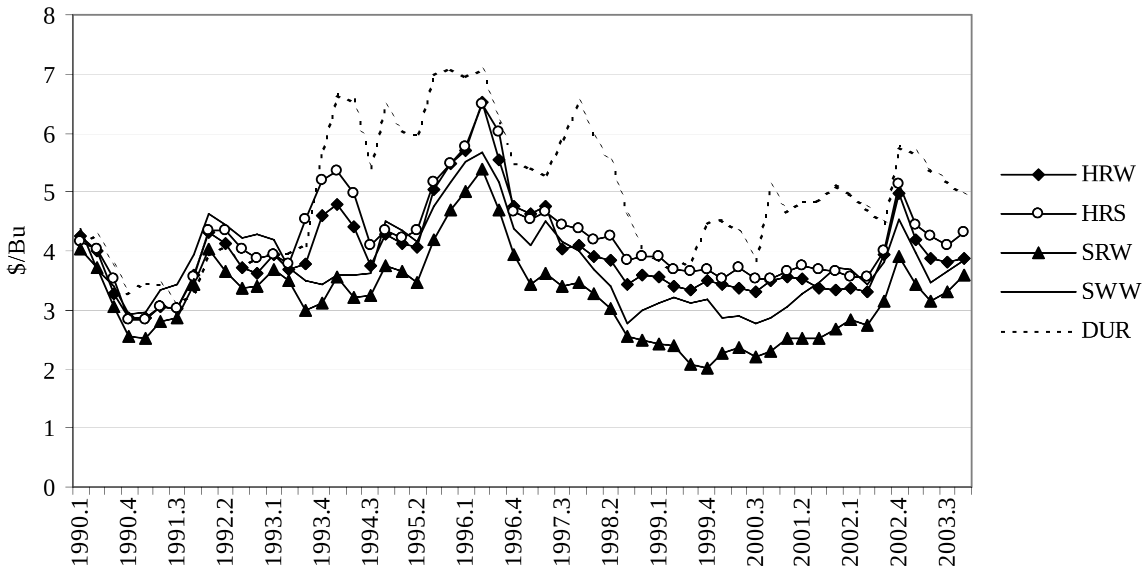


Figure 1, Quarterly wheat food use by class in U.S., 1990.1-2003.4
millions bu
120
100
80
60
40
20
0
1990.1
1990.4
1991.3
1992.2
1993.1
1993.4
1994.3
1995.2
1996.1
1996.4
1997.3
1998.2
1999.1
1999.4
2000.3
2001.2
2002.1
2002.4
2003.3
HRW
HRS
SRW
SWW
DUR

Figure 2, Quarterly prices of wheat food use in U.S. by class 1990.1-2003.4



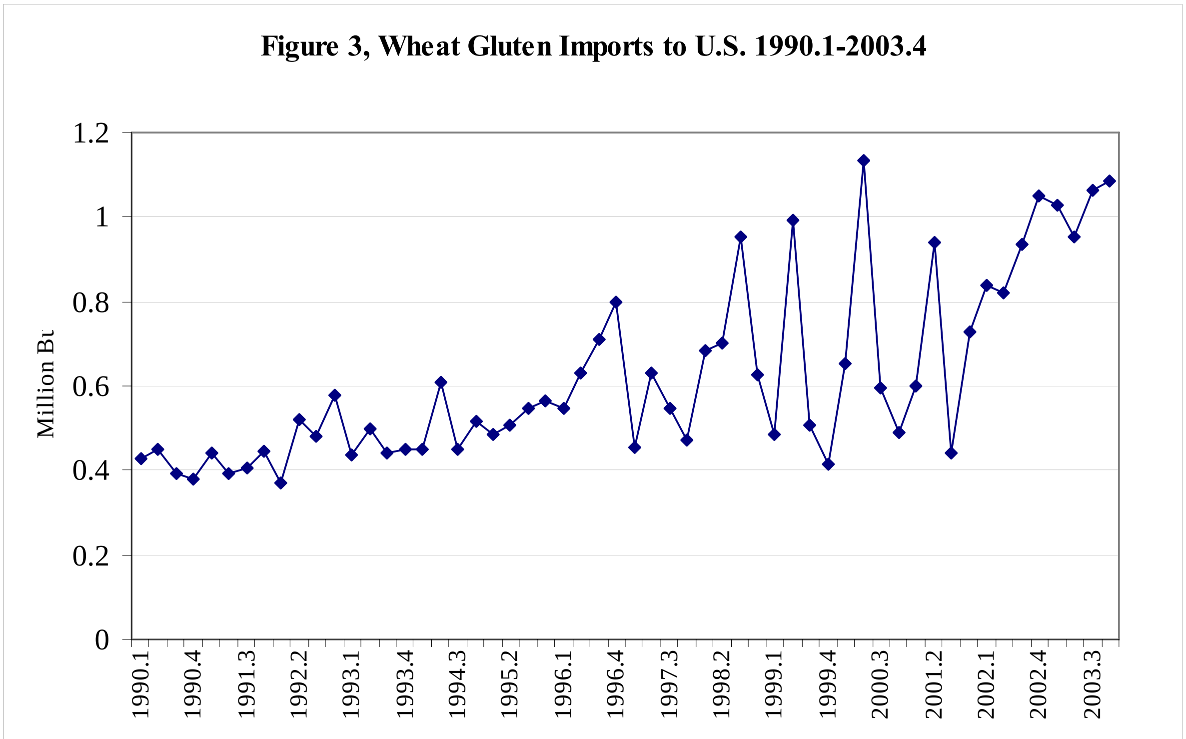


Figure 3, Wheat Gluten Imports to U.S. 1990.1-2003.4
Million Bu
1.2
1
0.8
0.6
0.4
0.2
0
1990.1
1990.4
1991.3
1992.2
1993.1
1993.4
1994.3
1995.2
1996.1
1996.4
1997.3
1998.2
1999.1
1999.4
2000.3
2001.2
2002.1
2002.4
2003.3