# A Time-Series Analysis of the U.S. Durum Wheat and Pasta Markets

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A quarterly, partial-equilibrium vector-autoregression model of the U.S. durum wheat and pasta markets was estimated and simulated under three trade-barrier changes that are of potential relevance for the current round of WTO agricultural negotiations: a rise in the U.S. market-clearing durum wheat quantity from increased imports; a policy- or tariff-reduction-induced decline in U.S. durum wheat price; and a tariff-induced rise in U.S. pasta product prices. In response to each shock, an array of quarterly dynamic response characteristics are examined: response reaction times, direction and pattern of quarterly responses, response durations, response multipliers, and strength of durum/pasta market interrelationships.

The value-added side of the food industry has often been neglected as a topic for empirical research. One of the main reasons for this omission is a lack of published data on these industries. Unlike commodities, such as corn, wheat, and soybeans, the United States Department of Agriculture (USDA) and other agencies often do not publish highly periodic (monthly or quarterly) data on quantities (demanded or supplied) or stocks of value-added products. Moreover, food industries typically keep information on their own prices, production, and distribution proprietary and thus out of the public purview. Any existing data on food products typically concerns the retail side of the industry. As a result there are few studies offering empirical econometric parameter estimates that allow policymakers, business strategists, and researchers to quantify and determine the monthly or quarterly patterns of the impacts of market and policy changes on value-added products and the interaction of value-added product markets with those markets which produce commodities that serve as inputs.

This paper addresses this gap in the literature by examining the nature of the interactions that exist within the durum wheat and pasta markets. The market-parameter and elasticity estimates from such analysis are then applied in the context of trade policy—for example, reduced tariffs or import restrictions. These applications attribute a change in price or quantity to a change in a trade barrier and then map out the effects of the commodity market for durum wheat and the related downstream markets (semolina and pasta). As with many components of the food industry, little or no published econometric research exists to illuminate market parameters of the U.S. durum wheat and pasta (dry, uncooked, non-egg)[1] markets, or on the dynamic nature of the interactions of these two markets. The primary goals of this paper are to demonstrate how one can work around serious data deficiencies and apply vector-autoregression (VAR) econometrics to appropriately and effectively model U.S. durum and wheat and pasta markets, and to obtain market parameters and market effects of proposed trade barrier changes. In the following sections, this paper:

- introduces the VAR methodology as an appropriate way of capturing dynamic interactions between a commodity market and data-deficient commodity-using markets downstream generally and between the U.S. durum and pasta markets specifically. Focus is placed on how the VAR methods' reduced form properties are invoked to successfully work around severe data deficien-

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[1] Herein, *dry pasta* refers to the products included in HTS 4-digit classification of 1902, in Chapter 19 of United States International Trade Commission (USITC 1999). Throughout, *dry pasta* refers to egg and non-egg dry, uncooked pasta, and excludes fresh and frozen pasta as well as dry pasta incorporated into manufactured products.

cies for downstream processed food markets.

- specifies and estimates a reduced-form VAR model of the U.S. durum wheat and pasta markets. Emphasis is placed on how the model meets rigorous diagnostic standards of specification adequacy in order to show that the reduced-form VAR model is appropriate and well-specified even in the face of severe data deficiencies and what some may deem abbreviated specifications by structural model standards.
- uses the VAR model to simulate three shocks likened to import-induced quantity increases and price changes (perhaps from policy or tariff changes) in pasta and durum wheat prices. Focus is placed on the dynamic simulation results and market parameter estimates which illuminate how such changes in policy or trade barriers affect upstream and downstream markets for durum-dependent products, and how useful such results would be to trade negotiators and policymakers. The results also capture the empirical nature of causal market relationships driving the U.S. durum wheat and pasta markets
- provides a summary, conclusions, and recommendations for future research on the nature of U.S. durum wheat and pasta market reactions to proposed policy and trade barrier changes implied by the imposed shocks that were simulated.

## Motivation for the Study

A number of recent trade issues concerning durum wheat and pasta demonstrate the need for tools that analyze the relationships between upstream and downstream products in the context of trade policy. One such issue concerns the imposition of tariffs on imports of Italian and Turkish pasta. In 1995, the United States International Trade Commission (USITC) and the U.S. Department of Commerce (USDOC) ruled affirmatively that certain U.S. imports of Italian and Turkish dry pasta were materially injuring or threatening to injure the U.S. industry (USITC 1996). The USITC imposed countervailing and antidumping duties (CVDs, ADs) on such imports in 1995 and 1996 (USITC 1996; Rich 1999, pp. 4.16–4.19).[2] Remedial tariff policy is clearly a current issue for both the U.S. pasta market and the closely related U.S. durum wheat market. U.S. trade policy analysts would currently benefit—as they would during the CVD/AD investigations—from econometric estimates and other empirical guidance regarding the market parameters and dynamic relationships linking these two markets. Such estimates and empirical guidance would permit analysts to better understand the market effects of implementing and formulating the remedial pasta duties and the impact of proposed measures of future trade liberalization.

A second related issue revolves around the U.S./Canada wheat trade, which has been a longstanding, visible, and contentious issue for U.S. and Canadian Federal trade and agricultural authorities and policymakers (see Babula and Jabara 1999; Alston, Gray, and Sumner 1994 and 1999; Babula, Jabara, and Reeder 1996). There is currently increasing concern by U.S. growers of wheat (particularly durum wheat) over such issues. In September 2000 the North Dakota Wheat Commission filed a petition under Section 301 of the Trade Act of 1974 against the Canadian Wheat Board (CWB) alleging that the CWB is engaging in "unreasonable" trade practices which have resulted in economic harm to U.S. wheat growers; the U.S. Trade Representative initiated this investigation in October 2000 (USTR 2000; Inside Washington Publishers 2000a, b). The alleged practices were not specified

[2] Imposed in 1996, these duties ranged as follows: AD rates from 0.67 percent to 46.67 percent and CVD rates from 0 percent to 11.23 percent on subject imports of Italian pasta and AD rates from 56.87 percent to 63.29 percent and CVD rates from 3.87 percent to 15.82 percent on subject imports of Turkish pasta (USITC 1996; Rich 1999, p. 4.16). Note that the U.S. International Trade Commission voted affirmatively in the preliminary stages of the countervailing and antidumping duty (CVD, AD) cases, imposing preliminary duties: CVDs on October 17, 1995 and antidumping duties on January 19, 1996. The Commission ultimately voted affirmatively in the final stages of the cases and imposed final duties in July and August of 1996. Duties are herein considered to have been imposed from the onset of the preliminary duties. Hence both the preliminary duties were imposed in the 1995/1996 June/May wheat market year: the second quarter for CVDs and the third quarter for AD duties. Dates of imposition were provided by the cases' supervisory investigator in a USITC staff telephone conversation with the authors on January 5, 2000. For an explanation of the workings of CVD and AD cases at the U.S. International Trade Commission, see USITC (1998).

but reportedly concern how the CWB markets wheat, particularly durum and hard red spring wheat, in the United States and third country markets (USTR 2000; Inside Washington Publishers 2000a, b).[3] These issues have also resulted in other events and investigations since the early 1990s: a USITC competitive conditions study of the U.S. and Canadian durum wheat industries (USITC 1990); a USITC Section-22 investigation of whether U.S. imports of largely Canadian wheat materially injured the U.S. wheat program (USITC (1994)); a study by a U.S./Canadian Joint Commission on Grains of the U.S./Canadian wheat and grain trade (Canada/U.S. Joint Commission on Grains 1995; Reeder 1995); and separate temporary U.S. import quotas for durum and non-durum (mainly Canadian) wheat (Glickman and Kantor 1995; Reeder 1995).

Seidband (1999, pp. 1–2) recently noted a growing concern over increasing U.S. imports of Canadian durum wheat, as market year (MY) 1998/99 (*i.e.*, June 1998–May 1999) imports surged by about one- third from previous MY levels at a time when U.S. production was at a five-year high (Seidband 1999, pp. 1–2). Seidband also notes that while U.S. pasta exports to Canada are rising, Canadian durum wheat sales to the United States far exceed U.S. pasta sales in Canada in value; in fact, the value of Canada's pasta sales in the U.S. market exceeds the value of the U.S. pasta exports to Canada, with Canadian durum wheat sales to the U.S. notwithstanding (Seidband 1999, pp. 1–2).[4]

Osorio (1999) noted an increased interest in modeling the trade-induced impacts not only on the markets for raw commodities (e.g. grains) but on downstream markets that use these raw commodities. He contended that this is because downstream value-added products are often subject to higher tariff and nontariff barriers than raw commodities; indeed, the current round of WTO agricultural negotiations may well focus on such commodity-based value-added markets (Osorio 1999). Trade in wheat-derived products such as pasta may thus be as important an issue as trade in durum wheat during the current WTO agricultural round. Trade reforms in processed products such as pasta and its component inputs (in this case durum wheat and semolina) will have some level of impact along the production chain. Yet for trade negotiators to negotiate optimally they must also be able to identify the empirical and dynamic nature of the direct or primary market effects of a proposed barrier reduction and should know the magnitude and nature of any upstream or downstream effects in the economy.

Tools and modeling methods that provide policy makers and negotiators with empirical estimates of and empirical guidance concerning the effects of proposed changes in trade barriers on U.S. durum wheat and pasta markets are urgently needed, but data limitations have precluded the construction of such models and have thus precluded the provision of such empirical estimates and guidance. The next sections will introduce the VAR methodology as an appropriate means to work around these data deficiencies and apply the available data to the study of the dynamics in the durum wheat and pasta markets in the context of trade policy.

## VAR Econometrics

Under general conditions an m-component vector, indexed by time period t, admits an autoregressive representation expressed as

$$\text{(1)} \qquad \underline{x}(t) = \sum_{s=1}^{\infty} [\underline{b}(s)*\underline{x}(t-s) + \underline{e}(t)$$

where the $\Sigma$ with subscripts and superscripts denotes the summation operator from lag period 1 through infinity ($\infty$). Underlined characters represent matrices or vectors. The $\underline{b}(s)$ term is an m by m matrix of autoregressive (AR) regression coefficients, and $\underline{e}(t)$ is an m-element vector of white noise residuals or innovations (Bessler 1984, Sims 1980). The white noise nature of $\underline{e}(t)$ satisfies

[3] According to the U.S. Trade Representative (USTR 2000), the Section 301 investigation provides a means for U.S. business, farmers, and workers to solicit Federal Government aid in obtaining relief from burdensome and U.S. commerce-restricting trade practices. After an investigation that can last as long as one year, USTR has discretion on whether or not to take retaliatory action (Inside Washington Publishers 2000b; USITC 1998, p. 29).

[4] It should be noted that the majority of pasta imports from Canada in 1998 were produced by an American multinational that had moved a portion of its capacity from the United States to Canada. Before 1998 most pasta imports from Canada were produced by Canadian firms. See Rich (1999, p. 4-18).

(2) $E(\underline{e}(t)) = 0$ for all t, and

(3) $E(\underline{e}(t)*\underline{e}(t)') = 0$ for $t \neq s$;
$= \underline{S}$, a positive-definite, m*m covariance matrix, if t=s.

"E" denotes the expected value multiplier. For applied work, the infinite lag sequence of equation 1 must be small enough to be operational but large enough for the residuals to approximate white noise (Bessler 1984, p. 112; Hamilton 1994, ch. 11). A lag structure was chosen with methods presented below.

Compared to more conventional "structural" econometric models, VAR econometrics is an approach that reveals empirical regularities from time-ordered data. The approach imposes as few *a priori* theoretical restrictions as possible on the data. Rather, VAR models loosely utilize theory to suggest which variables constitute a dynamic system in equation 1 and permits the regularities embedded in the time-ordered data to reveal themselves (Bessler 1984). All variables in the system are initially considered endogenous and each variable influences itself and all other variables in the system with lags.

Strength of relationships among a model's endogenous variables may be revealed through a second tool or operation of VAR econometrics—the analysis of forecast error variance (FEV) decompositions. By *strength of relationships* we mean whether or not one variable's movements cause changes in another variable's movements, the degree of such influence, and the timing of such effects (Bessler 1984, p. 111). FEV decompositions are, at alternative horizons, attributed to shocks in each of the dynamic system's series such that the desired measurement of strength of relationships emerges (Bessler 1984, p. 117). Bessler (1984, p. 111) established that FEV decomposition analysis is closely related to Granger causality testing. Yet unlike Granger causality testing, which simply discerns whether a causal link between variables exists, analysis of FEV decompositions provides further information on the timing and strength of such causal relationships (Bessler 1984, pp. 111-117).

## VAR Model Specification, Econometric Estimation, and Goals of Estimation and Simulation

This paper's purpose is to reveal the dynamic and empirical nature of the market parameters and relationships that drive and interrelate the U.S. durum wheat and pasta markets. This is done by estimating a quarterly VAR model of relevant U.S. durum wheat and pasta market relationships, simulating the model's impulse-response function with shocks attributable to notable current trends or proposed trade barrier changes, and analyzing the model's FEV decompositions.

### *VAR Model Specification*

To capture U.S. durum wheat and pasta market relationships, a VAR model of the following variables (hereafter denoted interchangeably by the parenthetical labels) was specified and estimated:

1. U.S. durum wheat price (PDURUM): This is reflected by the per-bushel dollar price of Minneapolis no. 1 hard amber durum wheat published by the USDA's Economic Research Service (USDA, ERS 1999b, p. 73). Quarterly MY values were calculated from published monthly values.
2. Durum wheat quantity supplied to and consumed in the U.S. market (QDURUM): This is reflected by the sum of the unpublished quarterly values of beginning stocks, imports, and production, obtained from the Economic Research Service (USDA, ERS, 1999a).
3. U.S. semolina price (PSEMOL): This price was calculated into quarterly MY prices using the monthly prices in dollars per hundredweight of semolina in Minneapolis obtained from *Milling and Baking News* (M&BN 1999).
4. U.S. price of pasta products (PPASTA): This price is reflected by the producer price index of pasta products, Series Number WPU021402, from the U.S. Department of Labor, Bureau of Labor Statistics (Labor, BLS 1999). Quarterly MY values were calculated from published monthly values.

Detailed derivations and summaries of VAR econometric methods are provided by Sims (1980), Bessler (1984), and Hamilton (1994, ch. 11). A VAR model posits each of the above four endogenous variables as a function of a specified number of lags of all four variables (Sims 1980; Bessler 1984; and Hamilton 1994, ch. 11). Schwarz's (1978) and Tiao and Box's (1978) lag selection criteria applied to the above data both suggested a one-order lag structure. The four-equation VAR is specified as:

(4) $X(t) = a_0 + a_{x,1}*PDURUM(1) + a_{x,2}*QDURUM(1) + a_{x,3}*PSEMOL(1) + a_{x,4}*PPASTA(1) + a_{x,T}*TREND + R_x(t)$

The parenthetical numbers refer to a value's time period: t for the current period and 1 for the one-order quarterly lagged value. The X(t) = PDURUM(t), QDURUM(t), PSEMOL(t), and PPASTA(t). The 0-subscript refers to the intercept. The a-coefficients reflect regression parameter estimates. The x-subscript denotes the x-th equation, the T-subscript refers to the time trend (TREND), and the second numeric subscript refers to each of the four lagged regressors. $R_x(t)$ refers to the current period t-estimates of the x-th equation's white noise residuals.

Quarterly data on the four endogenous variables were available from the first quarter of market year 1985 (hereafter 1985:1) through the fourth quarter of market year 1998 (or 1998:4).[5] The model was estimated over MY 1986:1-1998:4, because the four quarterly MY 1985 observations were "saved" for use in the Tiao-Box and Schwarz lag searches. For well-known reasons detailed by Bessler (1984) and Sims (1980) the four-equation VAR model was appropriately estimated with ordinary least squares. Doan's (1996) package, RATS, was used.

A number of considerations governed the VAR model's specification and estimation. First, the model was estimated in natural logarithms so shocks to and impulse responses in the logged variables reflect approximate proportional changes in the nonlogged variables. Second, dynamic results were desired, requiring quarterly (preferably monthly) data. Quarterly data were used because monthly QDURUM data are not available (USDA, ERS 1999b, pp. 42–47).

Third, unavailability of quarterly (or monthly) data on U.S. supply, consumption, shipments, or stock quantities of pasta precluded the inclusion of quantity variables for the pasta market downstream from the durum wheat market. While inclusion of pasta price and quantity variables would have been preferable, such pasta quantity data does not exist. In fact, such production data are no longer available even on an annual basis. Previous research and literature demonstrates how VAR econometric models, beset with data unavailability, may rely on a VAR model's reduced-form qualities and on the theory of the stochastic process to capture a market's forces of demand and supply through inclusion of a single price equation (Hamilton 1994, pp. 324–329; Babula 1996, p. 71). The diagnostics presented below show that the PPASTA equation appears to be an adequately specified reduced-form relation that likely reflects as much of the downstream pasta demand and supply conditions as limited data sources will allow. But further, and perhaps most importantly, this method of invoking VAR model reduced-form attributes is a viable econometric way to obtain empirical estimates of highly periodic market relationships given the unavailability of quarterly stock and quantity data for many downstream wheat-based products (Babula 2000).

Fourth, reliable quarterly data on U.S. supply, consumption, or stocks of semolina were not available.[6] The VAR model's reduced-form attributes

[5] The U.S. wheat market year runs from 1 June of a year through 31 May of the following year. Throughout, the numerals right of a quarterly date's colon refers to the MY quarter. For example, 1998:1 refers to the quarter spanning June, July and August; 1998:2 refers to the quarter spanning September, October, and November; 1998:3 to the quarter spanning December 1998 and January and February 1999; and 1998:4 to the quarter spanning March, April, and May of 1999. For quarterly market-year dates only the first year is listed; 1998/99 market year is taken as MY1998 and its quarters as 1998:1 through 1998:4.

[6] A quarterly market-year variable for U.S. semolina production was assembled from 1985–1996 monthly data and 1997–1999 quarterly data provided by the U.S. Department of Labor, Bureau of the Census (Labor, Census 1985–1999). However, the variable did not exhibit "rational" relationships with semolina price. In all simulations the quantity of U.S. semolina rose and fell (moved positively) with semolina price. According to reliable industry sources in a telephone conversation with the authors (18 January 2001), the reliability of this quantity variable is suspect because of alleged incompleteness of the surveys used to gather data. This conversation's contentions were confirmed by staff of *Milling and Baking News* (2000) in a front-page article concerning inaccuracies of the data. Sources for this data included Labor, Census (1985–1999). Calendar-year quarterly data were obtained from the quarterly reports for 1997–1999 (Commerce,

were again invoked to have a single reduced-form PSEMOL price equation capture the demand and supply elements of the U.S. semolina market. As with PPASTA, diagnostic evidence presented below suggests that PSEMOL is adequately specified, and that the reduced-form relationship likely captures as many of the market's demand and supply elements as limited data resources permit.

A number of binary (dummy) variables were considered. Following previous quarterly econometric research on U.S. wheat markets, centered seasonal binary variables were included to capture exogenous seasonal influences (Babula 2000; Babula, Jabara, and Reeder 1996; USITC 1994, pp. II.80–II.96 and appendix N). A binary variable, defined as 1.0 for 1995:2 and subsequent quarters and zero otherwise, was initially included in each VAR equation to account for the influences of the preliminary and final antidumping and/or countervailing duties imposed in 1995–1996 on certain U.S. imports of Turkish and Italian dry pasta (see footnote 2 and USITC 1996). We ultimately excluded this binary from all equations because evidence at the five-percent-significance level suggested that the coefficients were statistically zero in all four cases.

### *VAR Model Estimation and Diagnostics*

A number of important points concerning the adequacy of the estimated model's specification are made. Among these are the rationalization of the choice of a VAR model in logged levels over a vector error-correction model; adequacy of VAR model specification supported by an established battery of diagnostic test results; and evidence of time-invariance of parameters.

*Issue of cointegration.* Enough evidence emerged from the logged levels data to suggest that cointegration among the four logged endogenous variables may not be an issue. Consequently, the VAR model specified above was chosen over modeling the system as a vector error correction (VEC) model using Johansen and Juselius' (1990, 1992) maximum likelihood methods. When a vector system of individually nonstationary variables moves in tandem and in a stationary manner, the variables are said to be cointegrated (Johansen and Juselius 1990, 1992). With more than two cointegrated variables one should model the vector system as a VEC model with Johansen and Juselius' (1990, 1992) maximum-likelihood methods. Cointegration was not an issue here because evidence from several unit root tests suggested that the four endogenous variables are likely stationary in logged levels.

Two well-known unit root tests were applied to the logged levels of the VAR model's four endogenous variables: The Sargan-Bhargava (SB) test and the Dickey-Fuller (DF) test (or the augmented Dickey-Fuller [ADF] test where appropriate). The DF or ADF $T\mu$ tests were applied to all variables except PDURUM, and the DF or ADF $T\tau$ test was applied to PDURUM.[7] Harris (1995, pp. 27-29) and Kwiatowski, Phillips, Schmidt, and Shin (1992) discuss the well-known DF-type test problems of generating false conclusions of nonstationarity, particularly when samples are—as in this study—finite and when variables are stationary but have roots that are near-unity (*i.e.*, variables which are

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Census 1997–1999). Monthly data for 1990–1996 were obtained from the annual summary issues (Commerce, Census 1990–1996). These monthly data for 1990 –1996 were then converted to calendar-year quarterly values. Monthly 1985–1989 data were obtained from monthly issues of Commerce, Census (1985–1990). These 1985–1989 monthly data were also converted to calendar year quarterly values. The 1985–1999 quarterly data were then converted to a market year starting in July of one year and extending through June of the next year.

[7] Throughout, levels data and first differences are levels and first differences of the data in natural logarithms. The Sargan-Bhargava (SB) test is outlined in Sargan and Bhargava (1983); the Dickey-Fuller (DF) I: and IJ tests are detailed in Fuller (1976) and Dickey and Fuller (1979), while the "augmented" forms of the DF tests are outlined in Hamilton (1994, pp. 516, 528). The SB test entails running a regression of the variable against a constant—and a trend when appropriate—and concluding that evidence at the 5-percent-significance level is sufficient to reject the null hypothesis (null) of nonstationarity when the Durbin-Watson value equals or exceeds 0.26. The DF IJ or I: tests entail running a regression of a variable's first differences against a constant and a lag of the differenced dependent variable, with and without a trend, respectively. The augmented forms of these two tests, the ADF I: and IJ tests, are often used where at least one lagged dependent variable is included in the DF I: or DF IJ regression. In this paper the number of lagged dependent variables was determined using Akaike's (1973) information criterion program in Doan (1996, p. 5.18). Evidence is sufficient in the DF or ADF tests when the pseudo-t (I: or IJ) value on the non-differenced lagged regressor is both negative and of an absolute value exceeding the following critical values: -2.89 (5-percent level) and -2.58 (10-percent level) for the I: test, and -3.45 (5-percent level) and -3.15 (10-percent level) for the IJ tests (Hamilton (1994).

almost nonstationary).[8] In order to avoid erroneous treatment of such "almost nonstationary" variables as nonstationary, Kwiatowski et. al. recommend their KPSS test, which has versions with and without a time trend, as complementary evidence when variables generate marginal or ambiguous DF test evidence of nonstationarity. We placed increased reliance on KPSS test results in such marginal or ambiguous cases. Evidence from the SB, DF or ADF, and KPSS tests applied to the logged levels of the four endogenous variables suggests, on balance, that all series are stationary.[9]

*Estimation and Adequacy of Specification.* Following Babula et al. (1998) the model was judged adequately specified based on evidence generated by Ljung-Box portmanteau and DF unit root tests on the residual estimates of the four VAR equations. The Ljung-Box "Q" statistic is used to test the null hypothesis that the equation has been adequately specified with the null rejected for high Q-values (see Granger and Newbold 1986, pp. 99–101). Granger and Newbold (1986, pp. 99–101) caution against the exclusive reliance on the portmanteau tests for model adequacy. Consequently, DF $T\mu$ unit root tests were conducted on each VAR equation's residual estimates since stationary residual estimates provide evidence of adequate model specification (Babula et al. 1998, pp. 44–45). Evidence at the 1-percent-significance level from both tests suggests that all four equations have been adequately specified.[10]

*Time Invariance of Regression Parameter Estimates.* Two market developments may have induced structural change or time variance of parameters during the 1986:1–1998:4 quarterly sample: developments in U.S. consumer demand for pasta products during the mid-1990s and the previously

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[8] Harris (1995, pp. 27-29) notes that DF-type unit root tests often fail to reject the null hypothesis of nonstationarity when samples are small and/or when time series are stationary but have near-unity roots. Kwiatowski et. al. (1992) further contend that classical hypothesis testing usually requires strong evidence to reject a null hypothesis, which is nonstationary for the DF-type tests, and note that using their KPSS test with the opposite null, a null of *stationarity,* is useful as a complementary test for consideration with marginal or ambiguous DF test results. We followed this recommendation and complemented the DF and/or ADF tests with the KPSS test and placed increased reliance on the latter when the DF or ADF test results were marginally or ambiguously indicative of nonstationarity. The KPSS test was used to "break the tie" in such cases.

[9] SB and DF evidence at the 5-percent-significance level (5-percent level) is sufficient to reject the null that QDURUM is nonstationary: the SB test value of 1.39 exceeds the critical value of 0.26 and the DF I: test value of -5.7 exceeds in absolute value the critical value of -2.89.

PSEMOL's SB and ADF test evidence is marginally sufficient to reject the null of nonstationarity: while the SB test value of 0.30 exceeds the 0.26 critical value (5-percent level) the ADF I: value of -2.5 falls below in absolute value but nearly equals the critical value of -2.58 (10-percent level). Complementary KPSS test evidence at the 5-percent level is <u>insufficient</u> to reject the null that PSEMOL is <u>stationary</u>, as the test value of 0.254 falls below the critical value of 0.463. Coupled with the marginal SB- and DF-type evidence, this KPSS test evidence led to our conclusion that PSEMOL is likely stationary.

SB and ADF evidence concerning PPASTA's stationarity is ambiguous. While the ADF I: test value of -2.6 exceeds in absolute value the critical value of -2.58, suggesting that evidence at the 10-percent level is sufficient to reject the null of nonstationarity, the SB value of 0.09 falls below the critical 0.26 critical value, suggesting that evidence at the 5-percent level is insufficient to reject the null of nonstationarity. DF evidence suggests stationarity, but only at the 10-percent level, while SB evidence is insufficient to conclude stationarity. However, complementary KPSS test evidence at the 1-percent level <u>fails to reject</u> the null of PPASTA's <u>stationarity</u> since the test value of 0.64 falls below the critical value of 0.739. The ambiguous evidence from the DF and SB tests combined with the KPSS evidence of stationarity led to our conclusion that PPASTA is likely stationary.

PDURUM's SB and ADF test evidence is ambiguous concerning the acceptance or rejection of the null of nonstationarity: while SB's test value of 0.29 exceeds the critical value of 0.26 which rejects the null at the 5 percent level, the ADF IJ value of -1.8 fails to exceed the critical value of -3.15 and to reject the null of nonstationary at even the 10-percent level. However, complementary KPSS test evidence at the 5-percent level fails to reject the null of stationarity, since the test value of 0.075 is less than the critical value of 0.146 (note the KPSS test version inclusive of a trend was used for PDURUM). Taken with the ambiguous SB and DF-type evidence of stationarity the KPSS test evidence of PDURUM's stationarity led to the conclusion that the series is likely stationary.

[10] The VAR model's four equations generated Ljung - Box (Q) values ranging from 11.2 to 19.7, which are all less than the critical chi-square value of 27.7 (13 degrees of freedom; 1-percent significance), suggesting that evidence in all four cases is insufficient to reject the null hypothesis of adequate specification (Granger and Newbold 1986, pp. 99-101). The four equations' residuals generated DF I: values that ranged from -6.0 to -8.4, which are all negative and have absolute values in excess of those of the -3.51 critical value, suggesting that evidence at the 1-percent-significance level is sufficient to reject the null hypothesis of nonstationary residuals and to conclude that all four equations are likely adequately specified.

mentioned 1995 and 1996 imposition of antidumping and countervailing duties on certain U.S. imports of Italian and Turkish dry pasta. Existence of structural change generally signifies that market relationships embedded in the regression coefficients have changed such that the regression estimates themselves vary over time and that the coefficients estimated over the entire sample period are invalid. Existence of structural change usually requires division of the sample into subsamples at junctures of the change's occurrence and reestimation of the model separately for the subperiods (Larue and Babula 1995, pp. 163–164). If patterns of change—perhaps from the consumer developments and/or imposition of antidumping and countervailing duties—were not adequate to induce structural change and time variance of parameters, it is appropriate to estimate over the entire sample period and proceed as if parameter estimates are time invariant (Babula 1997).

Developments in consumer demand for pasta in the mid-to-late 1990s in the United States could have induced structural change or time-variance of parameters over the 1986:1–1998:4 quarterly sample. The growth of consumption in the pasta market, which accelerated during the early 1990s, slowed considerably in the mid-1990s. According to Leath (1999), per-capita consumption of semolina-based products (mainly pasta) peaked in 1994 at 13.9 pounds per person, and has since fallen to 12.5 pounds per person in 1997 and 11.6 pounds per person in 1998. These trends are reflected in the sluggish sales figures for pasta, with dry pasta sales down 4 percent over 1997-98 (Food Institute Report 1998: 5). Moreover, consumer tastes may be changing, as evidenced by the pasta industry's development of new niche markets such as flavored pasta, organic pasta, and dessert pastas (Sjerven, 1996, pp. 30-31), and by new applications of pasta products, including complete pasta dinners and non-Italian food applications of pasta (Bloom 1997, pp. 30 and 34).

A second potential source of structural change in the pasta industry concerns the CVD and AD duties petitioned by the U.S. industry and imposed on selected U.S. imports of Italian and Turkish pasta. In 1996, the USDOC assessed CVD rates ranging from 0 to 11.23 percent on certain Italian pasta manufacturers and from 3.87 to 15.82 percent on certain Turkish manufacturers (Rich, 1999: 4–16). Antidumping margins imposed on certain Italian pasta manufacturers (weighted-average less-than-fair-value margins) were much larger and ranged from 0.67 to 46.67 percent, while these AD margins for Turkish manufacturers ranged between 56.87 and 63.29 percent.

The antidumping and countervailing duties (ADs and CVDs) imposed on certain imports of Italian and Turkish pasta may have induced structural change by eliciting changes in the U.S. market shares captured by foreign suppliers and by changing the volumes of imported product. This is because, as is well known, the ADs and CVDs imposed on Italian and Turkish pasta were not time-enduring and equivalent duties levied on all supplies of each affected country (USITC 1993, pp. 8 –8; 1996). Instead, the orders reflected an array of firm-specific duties that varied across each affected nation's suppliers and were not necessarily imposed on all of each country's dry pasta exports to the United States. Further, these duties vary annually as the U.S. Department of Commerce reviews and adjusts the firm-specific duties in accordance with the remaining dumping or subsidy margin.[11]

Following established econometric research procedures, each estimated equation was subjected to a two-tiered structural-change test method that combines the CUSUM/CUSUM-squared and Chow test procedures (Larue and Babula 1994; USITC 1997, pp. 5.54–5.66 and appendix G; and Babula et. al. 1998 pp. 45–46). In the first tier the recursive residuals for the VAR equations were generated using Doan's (1996) RATS software and the data-analytic CUSUM/CUSUM-squared plot-test methods detailed in Harvey (1990, pp. 153–155) were applied to each equation's recursive residuals to discern potential points or junctures of structural change. Three junctures of potential structural change were prescribed by the CUSUM/CUSUM-squared tests: two breaks at 1991:1 and 1993:2 for the PSEMOL equation and one break at 1997:2 for the PPASTA equation. In the second tier a Chow test for structural change was conducted for the relevant equation at each potential juncture of change indicated by the CUSUM/CUSUM-squared tests (see USITC 1997, pp. 5.54–5.66 and appen-

[11] For details on how an AD and/or CVD order is implemented, see USITC (1998, pp. 8-9; 1995; 1996). These details relevant here are concisely summarized in Babula (1997, pp. 82-83).

dix G; Babula et al. 1998, pp. 45–46; Larue and Babula 1994, pp. 163–164). One concludes that no structural change occurred if the Chow F-test generates evidence at the one-percent level that is insufficient to reject the null of no structural change at each potential juncture (Larue and Babula 1994, pp. 163–164). Evidence at the one-percent-significance level was insufficient to reject the null hypothesis of no structural change for the semolina price equation at 1991:1 and 1993:2 or for the pasta price equation at 1997:2.[12] The failure of the latter period is especially noteworthy given that the period 1997:2 is conceivably a potential juncture at which market effects from the CVDs and ADs imposed on certain U.S. imports of Italian and Turkish pasta in 1995 and 1996 could have become manifest.[13]

Competing circumstances could have contributed to the lack of evidence supporting structural change, particularly with respect to pasta. On the consumption side the lack of evidence supporting structural change implies that the decline in the consumption of pasta may simply have resulted from lower demand for dry pasta rather than from a wholesale change in the structure of the market. From the standpoint of the CVD and AD duties, the failure to induce structural change is more interesting and likely due to a number of simultaneously occurring events that undermined any possibility for the market to be significantly altered. Results from the USITC (1996, pp. II.11 and II.15) suggest that there are moderate to high degrees of substitutability between both U.S.-produced and imported dry pasta and between dry pasta supplies from competing assessed and non-assessed foreign dry pasta suppliers. It is also noteworthy that not all Italian or Turkish manufacturers were assessed AD or CVD duties. Given this degree of substitutability it is possible that the share of the U.S. pasta market lost by assessed Italian and Turkish suppliers may simply have been offset by increased U.S. sales of close substitutes by non-assessed foreign suppliers from Italy, Turkey, and other countries such as Canada. Indeed, Rich (1999) noted that while U.S. imports of assessed Turkish pasta suppliers plummeted during the 1996–1998 period following the duties' imposition, imports of non-assessed Italian suppliers rose along with U.S. sales by foreign suppliers from other countries—most notably Canada—to the point that U.S. aggregate volumes of pasta imports were not reduced for sustained periods after imposition of the duties (Rich 1999, pp. 4.17–4.19).

## Three Model Simulations with the Impulse Response Function

One aspect of VAR econometrics useful in applied work is the impulse-response function. The impulse-response function simulates over time the effect of a one-time shock in one of the system's series on that series and on the other series in the system. This is done by converting the VAR model into its moving average (MA) representation (Bessler 1984; Hamilton 1994, ch. 11). The parameters of the MA representation are complex nonlinear combinations of the VAR regression coefficients. By imposing a one-time exogenous shock on one of the VAR variables on the system, one may examine the quarterly impulse responses of the other respondent endogenous variables and discern what the sample's long-run and historical trends would generate as the five dynamic and empirical attributes of U.S. durum wheat and pasta market relationships and market interactions (Bessler 1984; Babula et al. 1998; and Hamilton 1994, ch. 11). These include (a) reaction times of quarterly responses of affected or "respondent" variables, (b) directions and patterns of quarterly responses, (c) response durations, (d) magnitudes of overall responses, and (e) strength of relationships among U.S. durum wheat and pasta market variables (Bessler 1984).

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[12] The Chow F-tests at the 3 potential junctures of structural change were conducted at the 1-percent-significance level. For PSEMOL the test value of 1.71 fell below the critical F-value of 2.97 (9 and 34 degrees of freedom) for the 1991:1 break and the test value of 1.87 fell below the critical F-value of 2.97 (9 and 24 degrees of freedom) at the 1993:2 break. For PPASTA the test value of 0.67 fell below the critical F-value of 3.18 (7 and 36 degrees of freedom). In all 3 cases the test values fell below the critical F-values, suggesting insufficient evidence to reject the null hypothesis of no structural change.

[13] As noted earlier, binary variables were included to capture effects of the imposed ADs and CVDs imposed in 1995 and 1996 on certain U.S. imports of Turkish and Italian dry pasta. The statistical insignificance of this binary variable initially included in each VAR equation supports this evidence that suggests an absence of structural change and of time variance of parameters.

The model's four equations may have contemporaneously correlated current errors or innovations. Failure to correct for contemporaneously correlated current errors may produce impulse responses that do not reflect historical patterns (Sims 1980; Bessler 1984). Procedures from previous research using quarterly reduced-form U.S. wheat models were followed and a different Choleski decomposition was imposed on each of the model's three simulations described below (Babula 2000; USITC 1994, pp. II.80–II.96 and appendix N). As Sims (1980) and Bessler (1984) note, the Choleski decomposition resolves the issue of contemporaneously correlated current innovations. Each decomposition requires an arbitrary imposition of a theoretically-based Wold causal ordering among the current values of the model's four variables in each simulation. These three orderings are presented below.

Using literature-established methods, multipliers are calculated from each simulation's statistically nonzero impulse responses (see Babula, Colling and Gajewski 1994, p. 380). The multipliers are similar to elasticities and indicate history's long-run average percentage change in a response variable per percentage change in the shock variable. Sign is important: a positive multiplier suggests that each percentage change in the shock variable directionally coincided with the shock variable changes (hereafter, "similarly directed" responses), while a negative multiplier suggests that a variable's response was in the opposite direction of the shock (hereafter, "oppositely directed" responses).

Following previous VAR econometric research, Kloek and Van Dijk's (1978) Monte Carlo simulation methods were used to generate t-statistics for the impulse responses (Babula 2000; Babula et al. 1994, 1996, 1998; and USITC 1994, pp. II.80–II.96 and appendix N). We focused our analysis on the impulse responses which were statistically nonzero at the 10-percent-significance level.

The VAR model's impulse-response function was simulated with three shocks discussed below. Following accepted procedures in the literature, imposed shocks are of the magnitude of a single standard error of the variable's innovation (Bessler 1984; Babula et. al. 1994). Nonetheless, the model is linear, so the chosen size of the shock is arbitrary.[14] Three VAR model simulations were chosen:[15]

Simulation 1: a one-standard-error (11.5 percent) rise in QDURUM to examine the dynamic aspects of the elicited quarterly responses in PDURUM, PSEMOL, and PPASTA. This shock could be the result of increased domestic production or imports from relaxed quantity quota restrictions.

Simulation 2: a one-standard-error (10.9 percent) decline in PDURUM to examine the dynamic aspects of elicited quarterly responses in QDURUM, PSEMOL, and PPASTA. This PDURUM shock could arise from a reduction in import tariffs.

Simulation 3: a one-standard-error (2.37 percent) rise in PPASTA to examine the dynamic aspects of the elicited quarterly responses in PSEMOL, PDURUM, and QDURUM. This shock could arise from raising the zero import duty on dry pasta, especially given that domestic and imported pasta products are highly substitutable.

Table 1 provides the dynamic and empirical aspects of the responses generated from these three reduced-form model simulations which could possibly arise from, among other things, a rise in U.S. durum wheat market access (simulation 1), a decline in the U.S. tariff on durum wheat (simulation 2), and a tariff-induced rise in U.S. pasta price (simulation 3). It is important to point out that there is some subjective leeway in identifying the source of the shocks imposed on this (or on another) reduced-form model. While the assumed sources of the shocks in the simulations above are valid, the shocks could have arisen from other sources, since the VAR model's estimated reduced-form relations are neither prices nor quantities supplied or de-

[14] Babula et al. (1994, p. 377) pointed out that because of a VAR model's linearity one can characterize the impulse response simulations of a 20 percent shock simply by multiplying the impulse responses from a 10 percent shock by the scalar 2.0. Likewise, one can obtain the simulation results for a negative shock by multiplying the results from the simulation of a positive shock by the scalar -1.0.

[15] As required for each simulation's Choleski decomposition, a theoretically-based Wold causal ordering is chosen for each simulation. The orderings are those reflected in the descriptions of each simulation with the shock variable placed atop each ordering. So for example, simulation 1's ordering is QDURUM, PDURUM, PSEMOL, and PPASTA.

**Table 1. Dynamic Aspects of Responses to Shocks in QDURUM, PDURUM, and PPASTA**

| Respondent variable | Reaction times (quarters) | Response directions | Patterns of quarterly responses | Response Durations | Multipliers |
|---|---|---|---|---|---|
| Simulation 1: Rise in durum wheat quantity (QDURUM) | | | | | |
| PDURUM | 0 quarters | decline | U-shaped | 4 | -0.90 |
| PSEMOL | 0 quarters | decline | sharp, then decaying | 4 | -0.88 |
| PPASTA | | NSR | | | |
| Simulation 2: Decline in durum wheat price (PDURUM) | | | | | |
| QDURUM | 0 quarters | increase | U-shaped | 7 | -0.59 |
| PSEMOL | 0 quarters | decline | sharp, then decaying | 8 | 0.96 |
| PPASTA | 1 quarter | decline | increasing magnitudes | 5 | 0.08 |
| Simulation 3: Rise in pasta price (PPASTA) | | | | | |
| PSEMOL | | NSR | | | |
| PDURUM | | NSR | | | |
| QDURUM | 0 quarters | decrease | U-shaped | 5 | -2.91 |

Notes: Impulse responses were considered statistically nonzero at the 10%-significance level. "NSR" denotes that there were no impulse responses that were statistically nonzero at the 10%-significance level.

manded but market clearing prices which emerge after a full interplay of all—and often counterbalancing—demand and supply adjustments (Babula 2000). Other sources could have given rise to the same shocks. For example, simulation 1's shock of a presumed rise in QDURUM from increased imports could be explained by a rise in production while presumed tariff-induced shocks in wheat and pasta prices for simulations 2 and 3, respectively, could also have arisen from changes in production costs. A shock in a reduced-form model's price or quantity can therefore have a number of valid hypothesized sources. We chose to look at these shocks in the context of change in trade policy (e.g. changes in import, tariff levels, or negotiated changes in a quantity- or price-influencing farm policy).

*Simulation 1: An Imposed Rise in Durum Wheat Quantity (QDURUM).*

An increase in durum wheat quantity, perhaps resulting from a surge in U.S. durum wheat production or highly substitutable imports, was imposed on the model. The QDURUM increase induces a series of oppositely directed durum price responses (declines), which have what is herein considered an "immediate" or zero-quarter reaction time—that is the responses begin during the same quarter as (within 89 days of) the imposed shock. These declines take a "U-shaped pattern"—that is, a pattern of declines that initially accelerate, level off, and then decelerate in (absolute) magnitude. The PDURUM declines last an average of four quarters. The multiplier of -0.9 suggests that each one-percent rise in QDURUM elicits an average decline of 0.9 percent in durum wheat price. The reduced-form model cannot explain the reasons for the U-shaped response pattern for PDURUM, and we appeal to theory and market knowledge for such explanations. In the context of trade it is plausible that an injection of durum wheat imports, for example, will cause a decline in the price of durum wheat since the market will be supplied with both domestic and imported durum wheat. Domestic farmers, upon seeing the lower prices for durum wheat, may reduce production in an attempt to raise prices, but the effect of this behavior will occur with a lag, given the planting decisions involved. Prices will therefore fall until domestic planting decisions (namely, to lower production) have been accounted for in the marketplace, providing the bell-shaped shock in this market. So a hypothetical surge in imports would likely have a sustained negative impact on prices as the effects of the imports are transmitted through the economy and stocks of durum wheat.

Semolina price responses to the positive QDURUM shock are similar to durum wheat price responses, with a decline of about 0.9 percent for each percent rise in QDURUM. Cheaper durum wheat will clearly induce cheaper semolina. As with PDURUM's response pattern, we rely on theory and market knowledge for insights to help explain PSEMOL's pattern of responses—a sharp-then-decaying pattern which differs from the PDURUM response pattern just examined. PSEMOL historically begins responding sharply during the imposed shock's same quarter. Subsequent responses, which last over a four-quarter period, are much less dramatic than the first quarter response. In response to the price drop in semolina, users of semolina may increase market coverage (through 30 to 120 day contracts) of semolina at the lower price. Prices thus adjust quickly in the same quarter as the initial injection of durum wheat into the system. The surge in initial demand could moderate price pressures in future quarters, however. Pasta prices, on the other hand, do not seem to be affected directly by shocks in QDURUM. Instead, results suggest that durum wheat quantity changes influence the pasta market through impacts on durum wheat and semolina prices.

The data-embedded long-run market forces suggest that QDURUM increases elicit rather immediate declines in durum wheat and semolina prices. These PDURUM and PSEMOL declines last four quarters and are nearly proportional to the percentage rise in QDURUM. Increases in QDURUM—for instance, through an increase in imports—are thus likely to swiftly and noticeably affect durum wheat and semolina prices.

*Simulation 2: An Imposed Decline in Durum Wheat Price (PDURUM).*

Simulation 2 imposed a decline in U.S. durum wheat price that could arise from a reduction in the tariffs on durum wheat or a negotiated change in U.S. wheat policy. In response to the imposed reduction in PDURUM, the reduced-form model suggests that durum quantity supplied to and consumed in the market begins rising during the shock's same quarter, possibly due to augmented durum wheat demand. Each one-percent decline in durum wheat price ultimately elicits an average 0.6-percent increase in QDURUM. The QDURUM increases take a U-shaped pattern and last for an average of seven quarters. Theory and market knowledge provide a similar rationale for the shape of QDURUM's response pattern as in the case where there is an injection of durum wheat in the system. The fall in prices may allow for greater demand for durum wheat, supplied initially from imports that are part of QDURUM that was actually modeled. These price movements would likely be abated, however, as domestic farmers adjust their planting decisions to take into account the lower price received for durum wheat.

As expected, the imposed decline in durum wheat price elicits declines in both semolina and pasta prices. Semolina price decreases occur during the PDURUM shock's same quarter and take on a pattern of quarterly responses which initially are more pronounced in magnitude and then decay gradually over a period of up to eight quarters. The reasons for the shape of the shock are consistent with the results in the previous simulation. With PSEMOL's response multiplier of 0.96, semolina price declines are about proportional with declines in PDURUM.

PDURUM-induced pasta price declines take a full quarter to respond (that is, they have a one-quarter reaction time) and assume a pattern of accelerating magnitudes for about one year. The PPASTA response is muted, however, as reflected by the 0.08 multiplier that suggests each percent decline in PDURUM has elicited only an average 0.08-percent decline in PPASTA. Compared to the more marked PSEMOL response, the muted level of PPASTA in response to a change in durum wheat price is not surprising. Pasta is more of a manufactured good than semolina and embodies substantial value added through processing and manufacturing, so durum wheat price has less of a proportional impact on pasta prices.

More significantly, the myriad pricing, discount, and promotion practices that govern retail pasta sales (see USITC (1996), pp. 4.8–4.10 and appendix I) may cause wholesale prices actually paid to differ from published wholesale prices and hence from PPASTA constructed by Labor, BLS (1999) from these published prices. Three practices exist. First, some suppliers (mainly foreign suppliers) practice "line pricing" by charging one unit-price averaged across a full line of pasta products, ranging from low levels of specialization and low unit-production costs (e.g. spaghetti) to more specialized products with higher levels of specialization and higher unit-production costs (e.g. large shells) (USITC 1996, pp. 4.8–4.10). As a result, line prices may not vary with production costs for specific pasta products. Second, slotting fees—fees which pasta wholesalers or suppliers are required to pay retail chains in return for optimal amounts and locations of shelf space—represent lump-sum reductions in the prices charged by pasta producers/wholesalers to the grocery outlets that retail the pasta products and are thus discounts that are not reported in the wholesale prices (USITC 1996, pp. v.9–v.10). Third, pasta producers/wholesalers have a set of discount and promotion procedures that subsidize certain retailer or grocery promotion and merchandising activities on behalf of the pasta products. These activities include retailer discounts based on sales volumes, cooperative advertising allowances that subsidize grocery store advertising in local newspapers and other media, in-kind goods payments to the grocery retail outlets in place of slotting fees, and retailer or manufacturer discount coupons given directly to the retail consumer (USITC 1996, pp. v.8–v.10 and appendix I). Wholesale pasta prices actually paid are net of such slotting fees and promotional/discount arrangements, but published wholesale prices and PPASTA may not be.[16] Thus it is likely that movements in PPASTA actually modeled in the VAR model would be more muted or sluggish than the prices actually paid and may partly explain the sluggish response implied by the PPASTA multiplier.

Therefore, a drop in durum wheat price—perhaps from the reduction of an import tariff on highly substitutable durum wheat imports—will elicit a rise in the quantity of durum wheat supplied to and consumed in the U.S. market. This QDURUM increase will begin reacting immediately, will be enduring and may last up to two years, and will accumulate average increases of about 0.6 percent for each one-percent drop in durum wheat price. The imposed PDURUM decline influences the pasta industry through swift, pronounced, and sustained declines in semolina prices and in more delayed and muted declines in pasta prices.

*Simulation 3: An Imposed Increase in Pasta Price (PPASTA).*

The third and final simulation involved imposing an increase in pasta prices. Such an increase in pasta price could arise from increasing import duties on pasta product imports that are close substitutes with U.S.-produced products, allowing U.S. producers to raise pasta prices by the entire or partial margin of the tariff. The USITC (1996, p. I.23) uncovered evidence that domestically produced and imported dry pasta are probably moderately to highly substitutable for one another.

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[16] One author contacted an analyst of the U.S. Department of Labor involved with collecting information for and calculating the PPI for pasta products (Series Number WPU 021402) used as PPASTA in this study. The exact degree to which the series' values account for all of the slotting fees, discounts, and promotion practices is unknown. While the analyst noted that a sampling of the responding agent's questionnaires indicated that some of the discounts were subtracted from the prices underlying the PPI values, these subtracted items varied among responding agents. Furthermore, the analyst did not find any specific "slotting fees" in the sample of questionnaires he reviewed. So, while there is an effort to account for such discounts and promotional items, the exact degree to which Series WPU021402 accounts for such items appears unknown.

The summarized dynamics suggest that the sluggishly responsive pasta price has probably not fluctuated enough to directly influence semolina and durum wheat prices. This result is not surprising given that the U.S. pasta prices as published (and PPASTA formulated from these published prices) may not be prices actually paid and may not include variations inherent in line pricing, slotting fees, and the promotional/discount arrangements previously discussed (USITC 1996, pp. v.8–v.10 and appendix I). These results are supported by the results of the first two simulations showing that shocks to durum wheat quantity and durum wheat price have little or no effect on PPASTA.

More interestingly, while pasta price movements do not appear to directly influence durum wheat and semolina prices, pasta price increases do seem to influence the quantity of durum wheat, which is likely to ultimately influence the other durum-related variables. Each one-percent rise in the pasta price results in an immediate and greater-than-proportional (2.9 percent) decline in the quantity of durum wheat supplied to and consumed in the U.S. market that endures for up to five quarters. While the magnitude of the multiplier for QDURUM is relatively large it is not surprising in the context of the durum wheat market. For most non-durum varieties of wheat, there are many end-products (breads, flours, pastries, starches) that can be derived from these relatively substitutable types of wheat, so a price change for one particluar end product would exert a relatively small impact on the wheat market. In the case of durum wheat, however, there are few alternative uses for quality durum wheat other than the production of pasta (see USITC 1994, pp. II.5). Thus the impact from a shock in the pasta market would be expected to produce noticeable repercussions in the durum wheat market that negotiators and policy makers may well find interesting. Model results suggest that durum quantity's decline is U-shaped, with the largest decline occurring in the third and fourth quarters. Since the reduced-form model cannot provide reasons for this response pattern's shape, we appeal to theory and market knowledge for insights. We posit that while suppliers of durum wheat would react favorably to an increase in pasta prices, this effect would be initially overwhelmed by a negative effect on demand. Over subsequent quarters production may respond favorably to the price increases for pasta—although possibly not by enough to overwhelm the disincentives to consumption—creating favorable production effects that offset some of the demand-driven declines in QDURUM.

As noted earlier, the PPASTA increase imposed as a model shock could arise from, among other things, an increase in a tariff when imported and domestic pasta are highly substitutable. Given the high degree of substitutability between imported and domestically-produced pasta (USITC, 1996) and the duty-free status accorded to dry pasta not subject to the mentioned CVDs and ADs[17] (USITC, HTS 2000, chapter 19), the results of scenario 3 suggest that raising the general duty rate on dry uncooked pasta—as occurred with the AD/CVD duties placed on certain Italian and Turkish manufacturers of pasta—would likely have adverse upstream effects on the U.S. durum wheat market. An increase in the duty on imported dry pasta may allow domestic producers to raise prices on their import-substitutable products. The reduced-form model results suggest that the rise in pasta price may generate declines in derived demand for QDURUM, given that durum wheat has virtually no end-uses other than semolina and pasta. These net declines in QDURUM could result from demand declines for both domestic and imported pasta, both of which are included in the modeled QDURUM variable.[18] Each percent rise in PPASTA would immediately induce a series of five quarters' worth of QDURUM declines, which would ultimately register a 2.9 percent drop for each one-percent rise in pasta price.

## Strength of Relationships: Analyses of Forecast Error Variance Decompositions

Analysis of decompositions of forecast error variance (FEV) is another tool of VAR econometrics for discerning relationships among the mod-

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[17] Dry, uncooked, and otherwise unprepared pasta and packaged without sauces or other preparations is classified in Sections 1902.11.20 (if it contains eggs) and 1902.19.20 (if it does not contain eggs) of the Harmonized Tariff Schedule and enters the United States duty-free.

[18] Such QDURUM declines would be net and presumably would occur because demand-induced declines in QDURUM would outweigh any incentives to increase durum wheat production, since pasta, virtually the sole end use of durum wheat, is now higher-priced.

eled system's time series. As noted by Bessler (1984, p. 111), analysis of FEV decompositions is closely related to Granger causality analysis, and both tools provide evidence concerning the simple existence of a causal relationship among two modeled variables. But analysis of FEV decompositions goes further than Granger causality tests. A modeled endogenous variable's FEV is attributed at alternative time horizons to shocks in each modeled endogenous variable (including itself), and not only provides evidence of the existence of a relationship among two endogenous variables but illuminates the strength and dynamic timing of such a relationship (Bessler 1984, p. 111). Error decompositions attribute within-sample variance to alternative series and thus provide measures that are useful in applied work (Bessler 1984). Table 2 provides the FEV decompositions for the estimated VAR.

A highly exogenous variable has large proportions of its FEV attributed to its own variation and lower proportions to variation in other endogenous variables. Similarly, a highly endogenous variable has small proportions of its FEV attributed to own variation, and large FEV proportions attributed to the innovations of other variables (Bessler 1984, p. 111).

Perhaps one of the most evident results in Table 2 is durum wheat price's high degree of exogeneity, and its role as the modeled system's central and driving force. Durum wheat price is largely determined by production and planted acreage in Canada

**Table 2. Decompositions of Forecast Error Variance**

| | | Percent explanation of forecast error variance from | | | |
|---|---|---|---|---|---|
| Variable explained: | STEP | PDURUM | QDURUM | PSEMOL | PPASTA |
| PDURUM | 1 | 99.04 | 0.67 | 0.23 | 0.06 |
| | 2 | 97.99 | 0.84 | 0.74 | 0.43 |
| | 4 | 95.81 | 0.85 | 2.09 | 1.25 |
| | 6 | 94.09 | 0.78 | 3.42 | 1.71 |
| | 8 | 92.88 | 0.73 | 4.52 | 1.87 |
| QDURUM | 1 | 25.03 | 61.17 | 0.95 | 12.84 |
| | 2 | 31.47 | 50.03 | 1.73 | 16.76 |
| | 4 | 40.37 | 39.19 | 3.03 | 17.42 |
| | 6 | 45.26 | 34.13 | 4.13 | 16.48 |
| | 8 | 47.79 | 31.52 | 5.03 | 15.66 |
| PSEMOL | 1 | 93.31 | 0.48 | 6.03 | 0.18 |
| | 2 | 91.92 | 0.44 | 7.22 | 0.42 |
| | 4 | 89.67 | 0.37 | 9.21 | 0.76 |
| | 6 | 88.10 | 0.32 | 10.70 | 0.87 |
| | 8 | 87.05 | 0.30 | 11.76 | 0.89 |
| PPASTA | 1 | 4.43 | 2.98 | 1.14 | 91.46 |
| | 2 | 9.18 | 2.67 | 1.40 | 86.75 |
| | 4 | 18.48 | 2.38 | 2.04 | 77.11 |
| | 6 | 25.10 | 2.18 | 2.75 | 69.97 |
| | 8 | 29.10 | 2.05 | 3.44 | 65.42 |

as well as in the United States, and, in turn, on what farmers received at the end of the previous market year. As a result, data-embedded regularities suggest that PDURUM drives QDURUM to a far-greater degree than QDURUM drives PDURUM. U.S. durum wheat price's high degree of exogeneity may arise from the dependence of price on market actions—particularly planting decisions, past price movements, and stock movements—on *both* sides of the U.S./Canadian border, a political division runs through the center of North American durum wheat production. It is therefore not surprising that at all of Table 2's horizons more than 90 percent of the uncertainty of durum wheat price is attributed to its own variation.

Durum wheat quantity is relatively endogenous compared to PDURUM, with no more than 61 percent and no less than about 32 percent of its variation attributed to its own movements. The degree of QDURUM's exogeneity rapidly declines at horizons beyond two quarters. As expected, other than its own variation, the key explicator of durum wheat quantity behavior is durum wheat price, especially at longer horizons. Durum wheat price has moderate influence on QDURUM at shorter horizons (25 to 31 percent), but takes a more proactive role in explaining QDURUM movements at longer horizons, when nearly half of QDURUM's behavior is attributed to durum wheat price variation. One result that is perhaps surprising but that coincides with the third simulation's results is the noticeable proportion (up to about 17 percent) of QDURUM's variation attributed to pasta price movements.

Semolina price is highly endogenous, with no more than about 12 percent of its uncertainty attributed to own variation. As expected, the price of durum wheat—semolina's primary productive input—accounts for the majority (no less than 87 percent) of the variation in semolina price. This result coincides closely with the result of scenario 2, in which a price shock to durum wheat elicits an immediate, lengthy, similarly directed, and nearly proportional change in PSEMOL.

Pasta price appears largely exogenous, with no less than about 65 percent of its uncertainty attributed to own variation. Supporting impulse-response results of the second simulation, where PDURUM changes elicit similarly directed changes in PSEMOL, FEV decompositions suggest that durum wheat price contributes moderately to explaining pasta price variation. Such PDURUM contributions are lower at shorter horizons, and increase (up to 29 percent) at longer horizons.

**Summary and Conclusions**

The specification, estimation, and simulations of the VAR model reveal a rich set of data-embedded, long-run, and dynamic forces that not only govern the U.S. durum wheat and pasta markets, but characterize the nature of the two markets' interaction as well. These dynamics are highly relevant to farm- and trade-policy makers, agribusiness agents, and researchers involved with issues relevant to U.S. durum wheat and pasta markets. This is because changes in prices or quantities produced or consumed—arising perhaps from negotiated changes in tariff and nontariff trade barriers or from legislated or negotiated farm policy changes—affect prices and quantities in the immediate market as well as the related markets upstream or downstream.

The paper's first result emerges from the structural change test evidence: changing pasta consumption patterns and the 1995–1996 imposition of antidumping and countervailing duties on certain U.S. imports of Italian and Turkish dry pasta failed to induce structural change. This may arise from what the USITC (1996, pp. II.11 and II.15) reported as moderate to high degrees of substitutability, and hence fungibility, between U.S. and imported dry pasta, and between U.S. imports from competing foreign suppliers. The imposed duties failed to reduce imports because the U.S. market turned from assessed or subject dry-pasta imports to competing and substitutable non-assessed supplies in order to offset any duty-induced shortages (see Rich 1999).

The impulse-response simulations and analyses of FEV decompositions generated several results. First, a rise in the U.S. quantity of durum wheat—resulting perhaps from increased domestic production or a surge in imports—affects the market immediately with about a year's worth of price declines for both durum wheat and semolina. The pasta market, however, does not appear directly affected by the imposed durum wheat quantity increase. Second, policy changes that reduce durum wheat prices, such as changes in tariffs or farm policies that lead to lower farm prices for durum wheat, affect all modeled markets, including the pasta market. The impact of such a policy immediately affects the durum wheat and semolina markets with durum quantity increases and price declines lasting up to 7 or 8 quarters. The reduction in durum wheat prices causes durum wheat quantity to increase, as the reduced-form model captures demand increases that likely outweigh the supply declines, and causes a nearly proportional decline in semolina price as lower input costs are passed on to semolina millers. Pasta prices also decline for over a year, but such declines are far less proportional than the decline in PDURUM, presumably because pasta prices are far less durum-dependent than are semolina prices. The sluggishness or reluctance of PPASTA response may also arise from the exclusion by PPASTA of price variation from slotting fees, line pricing, and the various discount and promotional procedures that influence the pasta prices actually paid. Finally, while changes in durum wheat quantities and prices have little or no effect on pasta price, the final simulation—an increase in pasta price that could result from a tariff increase—shows adverse upstream effects on the durum wheat market: oppositely directed movements (declines) in durum wheat quantity. Decreased volumes of domestically demanded and imported durum wheat, both components in the modeled QDURUM variable, may account for this effect. This is not surprising given that there are few other demands for high-quality durum wheat and semolina other than for pasta production.

Future research along these lines could be threefold. First, in terms of the pasta-durum market, policy simulations incorporating greater sources of data, such as reliable quantity data for semolina and pasta, would provide a much richer analysis and allow for a wider range of policy simulations that would be of importance to practitioners. Given that both quantity and stock data are unavailable for pasta products on a highly periodic (monthly or quarterly) basis, the methods used herein to invoke pasta market demand and supply elements through estimation of a single pasta price reduced-form equation were the only located method of obtaining highly periodic (quarterly) pasta-related econometric estimates of market parameters. Second, expanding this analysis to other countries of interest, particularly Canada and Italy, would be instructive in comparing the market effects in these countries relative to those in the United States. Such an analysis would elucidate market conditions and interactions that could cause bottlenecks in trade negotiations. More significantly, this type of analysis could be useful if extended to the analysis of other food markets, where empirical estimates of the parameters behind all segments of the value-added chain remain few and far between. This would be especially insightful in the context of the next round of trade negotiations. The Uruguay Round did little to reduce tariffs among many processed food products, as tariffs were often reduced by only the minimum amount (15 percent) or, where tariffs replaced non-tariff barriers such as quotas, the tariffs devised were prohibitively high. A greater empirical understanding of the structure of the processed food market from the finished product down to the raw inputs (which are often protected in foreign markets) would be extremely beneficial for trade practitioners to formulate modalities that could bring forth meaningful tariff reductions in this sector.

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