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Export Demand Elasticity Estimation for Major U.S. Crops

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Elevated prices for major U.S. commodities have renewed interest in the price sensitivity of foreign demand facing the United States. Although the elasticity of foreign demand plays an important role in discussions of U.S. farm policy, it is also a parameter that is much debated with the majority of studies in this area published over 20 years ago. We provide new estimates of the elasticity of export demand for U.S. corn, soybeans, and wheat using updated data and empirical techniques. Our estimates are useful for practical policy analysis as well as for researchers seeking to parameterize large-scale simulation models.

Key Words: agricultural trade, demand, exports

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The elasticity of export demand held for U.S. agricultural products is critical for understanding the impacts of changes in farm policy (Gardiner and Dixit, 1987; Carter and Gardiner, 1988). In the past, this elasticity has been called an "elusive but highly important parameter (that) has been neglected too long" (Tweeten, 1977). This description remains relevant today. In this study we define the price elasticity of export demand as the percentage change in exports associated with a 1% increase in the U.S.' internal price. Export demand refers to the summation of all importer excess demand functions less the summation of other exporters' excess supply functions. If export demand is elastic, then U.S. exports may fall dramatically if the U.S. introduces policies such as the ethanol mandate and land retirement programs that are likely to support its commodity prices. If export demand is inelastic, on the other hand, such policies could actually serve to strengthen U.S. export revenues.

The magnitude of the elasticity of export demand has long been debated and continues to be in need of a firmer empirical foundation (Magee, 1975; Gardiner and Dixit, 1987; Carter and Gardiner, 1988; Miller and Paarlberg, 2001). Different groups of agricultural policy researchers have very different understandings of the magnitude of export elasticities, which in turn conditions agricultural policy analysis. Within the U.S. Department of Agriculture (USDA), the view on the size of the elasticity of demand for U.S. crop exports has changed over the last several decades. In the 1970s the prevailing view may have been that export demands are inelastic, whereas in the 1980s, they may have been viewed as relatively elastic. During the 1985 farm bill debate, for example, it was assumed that export demand elasticities for U.S. farm products were greater than unity in absolute values (i.e., elastic). Based on this assumption, the loan rates for wheat, feed grains, soybeans, cotton, and rice were lowered on the premise

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that lower loan rates would lead to a decline in export prices and an increase in the volume and value of U.S. exports of these commodities (Devadoss and Meyers, 1990). At present, the elasticity of demand for U.S. crop exports may be viewed as being inelastic, that is, closer to the perception of the 1970s (Paarlberg, 2009).

In this study we attempt to resolve this issue by bringing new econometric evidence to bear on the elasticity of demand for U.S. exports of three major crops: corn, soybeans, and wheat. Our primary objective is to provide both short-run and long-run estimates of the export demand elasticity facing U.S. producers of corn, soybeans, and wheat. Secondary objectives are to understand how these elasticities may have changed over the last three decades and to reconcile some of the disparate results that are currently in the literature. The long-run estimates will tell us how much export demand will contract in the long run (e.g., after more than 1 year) if there is a 1% increase in price today. The elasticities are partially influenced by changes in exchange rates, global market structure, policy, technological change, and commodity differentiation. Although any of these may influence the degree of international price competition and buyer behavior, detailed analysis of these underlying factors is beyond the scope of this study. We provide general evidence on the aggregate nature of these effects.

A point long recognized in the literature is that measurement of price elasticities in international trade is quite difficult (Orcutt, 1950; Thursby and Thursby, 1988). Our approach corresponds to the so-called "calculation method" described in Gardiner and Dixit (1987). This approach starts from a basic identity concerning total commodity exports calculated as the sum of importer consumption less importer domestic production. We then derive a comparative static result that shows how exports change with respect to a change in the export price of a commodity. In this approach, the elasticity of demand for U.S. crop exports is determined by the domestic demand and supply elasticities of foreign importers and exporters, price transmission elasticities, and the importance of the foreign country as a producer and consumer. The elasticities that we report concern only the direct effect of a price change and do not account for cross-price effects or the indirect effects that the price change has on other economic variables such as income, foreign exchange holdings, and the consumer price index. In contrast to most previous studies in this literature, we account for the importance of other exporting countries in the analysis.

The calculation method itself is not new, having origins at least as far back as Tweeten (1967), Johnson (1977), and Bredahl, Meyers, and Collins (1979). What is new is our determination to bring statistical techniques and recent data to bear. Former studies calculate the export demand elasticity based on educated guesses about underlying parameters. Instead of assuming what these values are likely to be, and assuming they are invariant across countries, we econometrically estimate these values. This makes our approach more difficult but allows the data to parameterize the model instead of the researcher.

There has been debate as to whether the export demand elasticity is really a single number or a parameter that does not change over time. Some observers recommend thinking of it as a variable (Carter and Gardiner, 1988). In this light, we provide evidence regarding the time path of the export demand elasticities, taking into account changes in international market conditions. We elaborate on previous efforts in this regard, in particular a study by Meyers, Devadoss, and Helmar (1987).

To preview our results, we find the short-run elasticity for corn and soybeans to be approximately -1 and approximately half that for wheat when looking at averages for the 2001–2011 period. Over the longer run, estimates are –1.64 for corn, –1.45 for soybeans, and –1.25 for wheat, again when looking at averages over 2001–2011. In this respect the value of exports is unlikely to increase even if less corn, for example, is available for foreign consumption resulting from its use in biofuels production.

Conceptual Model

Our mathematical representation of the export demand elasticity draws from Tweeten (1967), Johnson (1977), and Bredahl, Meyers, and Collins (1979). We divide the world into major importing and exporting regions. Let i be an index of importers, $i = 1, \ldots, m$, and j be an index of exporters other than the United States, $j = 1, \ldots, n$. For a given commodity the price internal to a country is denoted p_i, p_j, or in the case of the United States, p_{US}. Demand in a country is denoted Q_{di} or Q_{dj} whereas supply is denoted Q_{si} or Q_{sj}. Let the level of U.S. exports into the world marketplace be denoted by Q_{ef}, which is defined as

$$Q_{ef} = \sum_i (Q_{di} - Q_{si}) - \sum_j (Q_{sj} - Q_{dj}). \tag{1}$$

Quantity demanded and supplied in each foreign country is a function of that country's price, which, in turn, is a function of the U.S. price to some degree. Our interest is to know what happens to foreign import demand under a hypothetical change in U.S. domestic market prices. We can derive this response by taking the derivative with respect to p_{US}:

$$\begin{aligned} \frac{dQ_{ef}}{dp_{US}} &= \sum_i \left[\frac{dQ_{di}}{dp_i}\frac{dp_i}{dp_{US}} - \frac{dQ_{si}}{dp_i}\frac{dp_i}{dp_{US}}\right] \\ &\quad - \sum_j \left[\frac{dQ_{sj}}{dp_j}\frac{dp_j}{dp_{US}} - \frac{dQ_{dj}}{dp_j}\frac{dp_j}{dp_{US}}\right]. \end{aligned} \tag{2}$$

We then multiply through and divide by a number of terms:

$$\begin{aligned} \frac{dQ_{ef}}{dp_{US}}\frac{p_{US}}{Q_{ef}}Q_{ef} &= \sum_i \left[\frac{dp_i}{dp_{US}}\frac{p_{US}}{p_i}\left(\frac{dQ_{di}}{dp_i}\frac{p_i}{Q_{di}}Q_{di} - \frac{dQ_{si}}{dp_i}\frac{p_i}{Q_{si}}Q_{si}\right)\right] \\ &\quad - \sum_j \left[\frac{dp_j}{dp_{US}}\frac{p_{US}}{p_j}\left(\frac{dQ_{sj}}{dp_j}\frac{p_j}{Q_{sj}}Q_{sj} - \frac{dQ_{dj}}{dp_j}\frac{p_j}{Q_{dj}}Q_{dj}\right)\right] \end{aligned} \tag{3}$$

which allows us to present the derivation in percentage changes:

$$\begin{aligned} \frac{d\ln Q_{ef}}{d\ln p_{US}} &= \sum_i \left[\frac{d\ln p_i}{d\ln p_{US}}\left(\frac{d\ln Q_{di}}{d\ln p_i}\frac{Q_{di}}{Q_{ef}} - \frac{d\ln Q_{si}}{d\ln p_i}\frac{Q_{si}}{Q_{ef}}\right)\right] \\ &\quad - \sum_j \left[\frac{d\ln p_j}{d\ln p_{US}}\left(\frac{d\ln Q_{sj}}{d\ln p_i}\frac{Q_{sj}}{Q_{ef}} - \frac{d\ln Q_{dj}}{d\ln p_j}\frac{Q_{dj}}{Q_{ef}}\right)\right] \end{aligned} \tag{4}$$

or more simply as:

$$\begin{aligned} E_{ef} &= \sum_i E_{pi}\left(E_{di}\frac{Q_{di}}{Q_{ef}} - E_{si}\frac{Q_{si}}{Q_{ef}}\right) \\ &\quad - \sum_j E_{pj}\left(E_{sj}\frac{Q_{sj}}{Q_{ef}} - E_{dj}\frac{Q_{dj}}{Q_{ef}}\right) \end{aligned}$$

Note that there is no special need to explicitly distinguish importers (i) from exporters (j). If we let $i = j$, (5) can be restated more simply:

$$E_{ef} = \sum_i E_{pi}\left(E_{di}\frac{Q_{di}}{Q_{ef}} - E_{si}\frac{Q_{si}}{Q_{ef}}\right). \tag{5}$$

When country i is an importer, the term in brackets in (5) will be positive because Q_{di} exceeds Q_{si} at a given price. When country i is an exporter, the term in brackets will be negative because Q_{si} exceeds Q_{di} at a given price. For the remainder of the analysis we use only i when denoting a country other than the United States. Country i can be either an importer or exporter.

To summarize the terms in equation (5), E_{ef} is the export demand elasticity facing the United States. The price transmission elasticities are E_{pi} and concern the percentage change in a country's price associated with a one percentage change in the U.S. price. E_{di} is domestic own-price elasticities of demand and E_{si} is own-price elasticities of supply. The ratios Q_{di}/Q_{ef} and Q_{di}/Q_{ef} are domestic consumption and production in i divided by total U.S. exports to these countries.

The price transmission elasticity is of particular importance and has received little empirical attention in previous studies of export demand responsiveness. If E_{pi}, for example, is less than one, there is some kind of policy distortion or a distortionary effect played by transportation costs that prevents price changes from being fully transmitted. The lower bound of E_{pi} is zero and implies that government policies completely insulate i's internal production and consumption prices from world market prices.

In most of the rest of the study we consider how to parameterize the components of equation (5). In past studies, authors have drawn on averages of historical estimates of supply and demand elasticities. For example, Johnson (1977) assumes domestic demand elasticities are –0.2 for wheat and cotton and –0.4 for feed grains and soybeans. He further assumes that all supply elasticities are 0.2. He implicitly assumes that all price transmission elasticities are one. Bredahl, Meyers, and Collins (1979) criticize Johnson's approach and argue that in some cases, price transmission elasticities may effectively be zero. Their parameterizations are essentially educated guesses, however, and in some cases are likely too extreme, that is, there is perfect price transmission or none at all. Our alternative is to econometrically estimate the parameters.

Estimation of Price Transmission Elasticities

To estimate E_{pi} from equation (5) we follow earlier literature including Abbott (1979) and Mittal and Reimer (2008). This involves regressing how the local price varies with a 1% change in the U.S. export price. Local prices in many importing countries are constrained from directly following U.S. export prices. This may be the result of transport and transaction costs, market power, exchange rates, domestic policies, and border policies (Reimer and Kang, 2010; Reimer and Li, 2010). Among border measures, nontariff barriers such as variable tariffs, tariff rate quota, prohibitive tariffs, and technical barriers may have strong effects on price transmission (Conforti, 2004). To be clear, we are not interested in identifying these potential determinants of price transmission; rather, we are simply measuring the degree of price transmission that exists. Because full adjustments are not likely to occur within a given period, we propose the following partial-adjustment model:

$$\begin{aligned}\ln p_i^t = {} & \beta_0 + \beta_1 \ln p_i^{t-1} \\ & + \beta_2 \ln p_{US}^t + \beta_3 TREND + \varepsilon_i^t,\end{aligned} \tag{6}$$

where p_i^t is country i's internal price at time t; p_i^{t-1} is country i's lagged price; p_{US}^t is the U.S. price; $TREND$ is an annual time trend 1,2,3, . . . to allow for the possibility of general changes over time; and the βs are parameters to estimate. β_2 and $\beta_2/(1-\beta_1)$ are the short- and long-run price transmission elasticities, respectively. The error term is denoted ε_i^t and initially assumed to be independently and identically distributed with mean zero and homoscedastic variance. We test these and other assumptions subsequently.

Data on the U.S. price in year t (p_{US}^t) are taken from the *Grain and Feed Market News*, USDA Agricultural Marketing Service and—depending on the commodity being considered—correspond to either No. 2 yellow corn at U.S. Gulf ports, No. 1 yellow soybeans at Chicago, or No. 1 hard red winter wheat at Kansas City (USDA AMS, 2012). These prices, which are annual dollars per metric ton, were chosen because they are fairly representative of U.S. domestic markets for these products.

The internal price of foreign country i in year t (p_i^t) is more of a challenge to obtain. Potential sources include the U.S. Department of Agriculture Global Agricultural Trade System (USDA GATS, 2012), which provides price paid by a country for a commodity loaded free alongside ship on a unit value basis. Another source is price data from foreign Customs offices, which we were able to obtain in the case of Japan. A third source is the Food and Agricultural Organization producer price from the PriceSTAT database (FAO, 2012). In general we use the FAO price series, because this is more of an internal price than the others. Using the FAO data is especially necessary for major producers and exporters (such as Argentina and Brazil) that may import little from the United States, because they would not have GATS or Customs prices available. Some importers, by contrast, produce essentially none of a commodity at home and therefore have no FAO producer price. In these few cases we use the GATS price data as the internal price converted to local currency units using a representative exchange rate between that country and the U.S.[1] The FAO data are in local currency units and do not need to be converted using an exchange rate.

[1] The use of unit value as a proxy for price can give rise to measurement error, but the evidence available to us suggests that such error is likely to be small in this study. The extent of the error can be evaluated in the case of Japan, for which Customs prices are also available. The correlations between Japan's unit values and Customs prices are 0.98, 0.95, and 0.95 in the case of wheat, corn, and soybeans, respectively. Results associated with Customs prices vs. unit values are nearly identical.

A downside of using the FAO price as an internal price is that it has annual observations only from 1991–2011. We are therefore left with few degrees of freedom by which to estimate standard errors and carry out statistical tests. On the other hand, it must be remembered that our objectives are different than those of many studies; instead of trying to determine whether a specific policy had an effect (i.e., nullify a hypothesis), we are trying to measure magnitudes of variables that are almost certain to have a close relationship with each other. Instead of just assuming an elasticity of 1.0, for example, we let the data reveal the scope of the relationship.

International markets for the three commodities tend to be dominated by a few exporting countries on one side and by numerous importers from many regions of the world on the other. The countries used to represent foreign demand and supply for each commodity are listed in Table 1. This sample is designed to be representative of each market, with two major exporters and four major importers of each commodity.

One econometric issue in estimating (6) is the potential endogeneity of the lagged value of the dependent variable on the right-hand side (p_i^{t-1}). If this is not independent of the error terms, then ordinary least squares estimation would, in most contexts, give rise to biased estimates in small samples. This consideration is offset, however, by the fact that (6) is a partial-adjustment model. In this specific context, the parameters can be consistently and efficiently estimated by ordinary least squares and related techniques (Greene, 2004, p. 568).

Table 1. Augmented Dickey-Fuller Test of Stationarity of Price Series

	MacKinnon Approximate p Value		
	Corn	Soybeans	Wheat
Algeria	—	—	0.621
Argentina	0.720	0.878	—
Australia	—	—	0.250
Brazil	0.455	0.864	—
Canada	—	—	0.056
China	0.018	0.207	—
Egypt	0.830	0.829	0.893
Japan	0.685	0.579	0.935
Mexico	0.812	0.818	0.663
United States	0.816	0.696	0.660

Note: Null hypothesis is nonstationary.

Dickey-Fuller tests reveal that all but two of the price series are nonstationary (Table 1). The danger of using nonstationary time series is that a relationship between two variables may be found where one does not exist, because *t*-statistics are not reliable (Greene, 2004). We do not view this as a problem, however, because the point of this exercise is to show that the relationship between U.S. price changes and foreign price changes is not as close as it is normally seen to be. The implicit assumption of previous studies is that there is a perfect relationship between the two variables. This explains why some earlier studies greatly overestimate the elasticity of export demand to U.S. price changes (discussed subsequently). These studies implicitly assumed that transmission of U.S. price changes to foreign countries is perfect, whereas in actuality the relationship is less than perfect. A second point to make here is that we are careful not to make any claims about causality or statistical significance in this context; we are only trying to gauge the magnitude of price adjustments between two variables that are almost certainly going to have some relationship as a result of the United States' international importance as an exporter of the three commodities.

We report results for a restricted maximum likelihood estimator, which has good properties under first-order autoregressive (AR[1]) errors (Greene, 2004, p. 273).[2] Table 2 reports the estimation of equation (6) for each of the six countries for corn, soybeans, and wheat. There are four columns in the table associated with the estimated coefficients. The coefficient on

[2] For each commodity and country, we used a Breusch-Godfrey LaGrange multiplier test to check for up to three potential lags under a 5% level of significance. In 15 of 18 cases, we do not reject a null hypothesis that one, two, or three autoregressive coefficients are simultaneously equal to zero. In other words, in most cases there is not first-, second-, or third-order autocorrelation. Note that even in those cases with autocorrelation, it does not affect the unbiasedness and consistency of an estimator; we merely measure the interrelatedness of a commodity's price at different stages of its marketing channel.

Table 2. Price Transmission Elasticities

	ln Lagged Own Price	ln U.S. Price	Trend	Intercept
Corn				
Argentina	0.147 (0.245)	0.848** (0.338)	0.062** (0.023)	–0.228 (1.212)
Brazil	0.092 (0.181)	0.831*** (0.186)	–0.003 (0.006)	0.466 (0.710)
China	0.460** (0.230)	0.077 (0.413)	0.160 (0.025)	3.425 (2.708)
Egypt	0.303 (0.300)	0.204 (0.161)	0.045* (0.021)	3.211** (1.365)
Japan	0.058 (0.113)	0.900*** (0.112)	0.015** (0.006)	4.509*** (0.848)
Mexico	0.513*** (0.193)	0.306** (0.130)	0.028* (0.015)	1.886 (1.389)
Soybeans				
Argentina	0.005 (0.024)	1.089*** (0.224)	0.069*** (0.011)	–0.784 (1.154)
Brazil	0.480*** (0.122)	0.656*** (0.116)	0.003 (0.005)	–0.799 (0.683)
China	0.614*** (0.150)	0.432** (0.181)	0.002 (0.012)	0.719 (1.546)
Egypt	0.375* (0.224)	0.135 (0.147)	0.035*** (0.012)	3.407*** (1.250)
Japan	0.464** (0.191)	0.460*** (0.138)	0.004 (0.008)	2.888 (1.959)
Mexico	0.597*** (0.178)	0.479*** (0.183)	0.025 (0.016)	0.294 (1.723)
Wheat				
Algeria	0.512*** (0.167)	0.667*** (0.172)	–0.013 (0.009)	–0.305 (1.060)
Australia	0.203 (0.151)	0.409*** (0.102)	0.004 (0.006)	2.278** (0.891)
Canada	0.571*** (0.159)	0.389*** (0.123)	–0.007 (0.007)	0.316 (1.019)
Egypt	0.252 (0.172)	0.442* (0.235)	0.044*** (0.013)	2.452*** (0.758)
Japan	0.363** (0.141)	0.482*** (0.074)	0.009** (0.004)	0.998* (0.604)
Mexico	0.635*** (0.175)	0.490*** (0.160)	0.015 (0.017)	0.146 (1.514)

Notes: Standard error is in parentheses. R^2 is not available because the equations estimated with the restricted maximum likelihood estimator. The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. All prices are in real terms.

the natural log of U.S. price is the short-run price transmission elasticity.

Nearly all of the short-run estimates lie within the (0,1) interval, which is what we might expect in the case of imperfect price transmission (Argentina soybeans is the sole exception).[3] In the case of corn, the short-run elasticity ranges from 0.077 in the case of China to 0.900 for the case of Japan (Table 2).[4] For soybeans the short-run elasticity ranges from 0.135 for Egypt to 1.089 for Argentina. For wheat the short-run elasticity ranges from 0.389 for Canada to 0.667 for Algeria.

To the limited extent that these estimates can be compared with previous studies (e.g., Abbott, 1979; Mittal and Reimer, 2008), they appear to be largely consistent with the types of results found before. The fact that price transmission elasticities are less than one suggests that tariff and nontariff barriers are an important feature of international commodity markets.

The estimated coefficient on the lagged natural log of own price lies within the unit interval for every country and commodity (Table 2). These estimates are used to calculate the long-run price transmission elasticities by way of the $\beta_2/(1-\beta_1)$ formula. In all cases the long-run price transmission elasticity ends up being more elastic than short-run price transmission elasticity, as we might have expected.

Demand and Supply

We now turn to the estimation of supply and demand elasticities within the importing country.

[3] When we estimate (6) with ordinary least squares, all of the results are very similar to the AR(1) results in Table 2. The results also not very sensitive to inclusion of the trend variable on the right-hand side. White (robust) standard errors are used in the case of Egypt corn and Egypt wheat, in which evidence of heteroscedasticity was detected through a Breusch-Pagan test.

[4] The price transmission elasticities for Japan are slightly lower when we use Japanese Customs prices in place of unit values. The differences are small enough that they do not affect our final elasticity calculations in a meaningful way.

These models are developed using the general approaches of Sadoulet and de Janvry (1995). In this section we generally suppress the foreign country index i for notational simplicity. Let Q_d^t denote consumption of a grain at time t and Q_s^t denote production of a grain at time t. In general Q_d^t and Q_s^t will differ from each other as a result of international trade ($Q_d^t \neq Q_s^t$). For an importer, $Q_d^t > Q_s^t$ at a given price, but for an exporter $Q_s^t > Q_d^t$ at a given price.

We treat the domestic market for each commodity as operating within a small open country where price is determined in the international market. Producer price expectations are modeled as a naïve, one-period lag process. We denote p_s^{t-1} as the lagged price of a representative producer in year $t-1$. In turn, p_d^t denotes the price of a representative buyer in year t. For each importer, we propose the following system of demand and supply equations:

$$\ln Q_d^t = \delta_0 + \delta_1 \ln p_d^t + \sum_{j=2}^{m} \delta_j \ln Y_j^t + \varepsilon^t \quad \text{(demand)} \tag{7}$$

$$\ln Q_s^t = \alpha_0 + \alpha_1 \ln Q_s^{t-1} + \alpha_2 \ln p_s^{t-1} + \sum_{j=3}^{n} \alpha_j \ln Z_j^{t-1} + \nu^t \quad \text{(supply)} \tag{8}$$

where the δs and αs are parameters to be estimated; Y_j^t are demand shifters such as the prices of substitutes, income, and population; and Z_j^{t-1} are (lagged) supply shifters such as prices of alternative commodities to produce. The error terms for (7) and (8) are ε^t and ν^t. Subsequently we check whether they are independent and identically distributed with zero means and homoscedastic variances.

In the demand equation (7), δ_1 corresponds to the elasticity of domestic demand. It is expected to be negative. In the supply equation (8), the previous year's supply (Q_s^{t-1}) is included on the right-hand side to allow for the possibility of a lag in the adjustment process. The parameter α_2 is the short-run elasticity of supply response to last period's price change. The long-run elasticity of supply response is calculated as $\alpha_2/(1-\alpha_1)$. Both of them are expected to be positive.

We do not model (7) and (8) as a simultaneous system. There might be a link between the two equations if consumption Q_d^t in (7) and production Q_s^t in (8) were the same from year to year. These would equal each other in autarky but with international trade, Q_d^t and Q_s^t are different. Because we do not model this wedge explicitly, in estimating supply and demand, we effectively take exports and imports as exogenously determined. Data on these two variables, Q_d^t and Q_s^t, are from the U.S. Department of Agriculture Production, Supply and Distribution database (USDA PSD, 2012) and are in 1000 metric tons.

Equations (7) and (8) could also conceivably be linked through the internal price. However, this is lagged in equation (8), and more importantly the internal price is not the same across equations as a result of margins between the farmgate level and the level of commodity users. In demand equation (7), the internal price is given by the GATS import value converted to local currency prices, except for the case of Japan, where Japan Customs data are used. In the case of supply equation (8), we rely on FAO prices in local currency units, as described previously.

Equations (7) and (8) are estimated on an individual basis using restricted maximum likelihood. Results for the demand and supply equations are reported in Tables 3, 4, and 5 for corn, soybeans, and wheat, respectively. The standard error of the estimator is reported in the tables below the coefficient in parenthesis. Although the coefficients typically have the expected sign, the coefficient of interest is not always statistically different from zero at conventional levels of significance. The topmost row in each table corresponds to the estimated price elasticity of demand. In Argentina, for example, this value for corn is –0.036, which means that a 1% increase in domestic price is associated with a 0.036 percent fall in consumption. Looking across the tables, the price elasticity of demand for corn ranges from –0.003 to –0.251 for corn (Table 3), –0.079 to –0.397 for soybeans (Table 4), and –0.001 to –0.207 for wheat (Table 5). Most of the countries have price-inelastic demand for the three commodities. For example, Japan's price elasticity of demand is –0.109 for corn (Table 3), –0.079 for soybeans (Table 4), and –0.010 for wheat (Table 4). Egypt's price elasticity of demand is also on the inelastic side at –0.241, –0.397, and –0.010 for

Table 3. Corn Demand and Supply Results

	Argentina	Brazil	China	Egypt	Japan	Mexico
Demand Equation						
ln corn price	–0.036 (0.089)	–0.251 (0.172)	–0.003 (0.008)	–0.241 (0.293)	–0.109 (0.079)	–0.197 (0.186)
ln wheat price	—	–0.298 (0.191)	0.278* (0.017)	0.160 (0.387)	0.085 (0.053)	0.269* (0.148)
ln income	0.418 (0.098)	2.259*** (0.805)	0.059 (0.041)	–0.324*** (0.112)	0.101 (0.175)	0.017 (0.331)
ln population	2.612 (0.888)	–1.295** (0.589)	0.621*** (0.147)	1.031*** (0.472)	–0.029 (0.102)	0.801* (0.430)
Intercept	–40.422 (15.107)	6.159*** (0.816)	9.376*** (0.189)	6.193*** (0.769)	9.208*** (0.636)	7.166*** (0.352)
Supply Equation						
Lagged ln corn price	0.450 (0.275)	0.104 (0.243)	0.060 (0.055)	0.006 (0.086)	0.018 (0.190)	0.598*** (0.200)
Lagged ln corn production	0.022 (0.231)	0.243 (0.250)	0.677*** (0.208)	0.533*** (0.180)	0.787*** (0.110)	0.025 (0.217)
Lagged ln wheat/soy price	–0.465* (0.280)	–0.243 (0.267)	—	0.002 (0.062)	–0.008 (0.009)	–0.395** (0.158)
Trend	0.042*** (0.012)	0.030*** (0.010)	—	0.004 (0.007)	0.001 (0.112)	—
Intercept	9.487*** (2.135)	8.424*** (2.371)	3.378 (2.239)	3.967** (1.748)	0.217 (0.593)	8.151*** (1.813)

Notes: Standard error is in parentheses. R^2 is not available for equations estimated with the restricted maximum likelihood estimator. The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

corn, soybeans, and wheat, respectively. It is not statistically different from zero in any of these cases.

Estimates of the short-run domestic supply elasticities are reported in the lower half of Tables 3–5. In Argentina, for example, this value for corn is 0.450, which means that a 1% increase in (lagged) producer price is associated with a 0.450 percent increase in production. Looking across the tables, the short-run elasticity of supply ranges from 0.006 to 0.598 for corn (Table 3), 0.029 to 0.795 for soybeans (Table 4), and 0.008 to 0.492 for wheat (Table 5).

In Tables 3–5 we have not reported the long-run elasticities of supply for lack of space. Recall that they are calculated using the $\alpha_2/(1-\alpha_1)$ formula from the estimated coefficients in (8). These estimates are reported in Table 6. In all cases we see that long-run producer response is more elastic than short-run producer response.

Export Demand Elasticities over Time

At this point we have estimated the key parameters of equation (5) and so are able to calculate the export demand elasticities in the case of corn, wheat, and soybeans. The left-hand columns of Table 6 summarize key elements needed to make this calculation, including short- and long-run price transmission elasticities, domestic demand elasticities, and short- and long-run supply elasticities. Domestic demand elasticities are assumed not to vary across short- and long-run specifications because lags are unlikely to be an important aspect of consumer behavior (unlike their role in the case of producer behavior). Production and consumption in foreign countries divided by U.S. exports (Q_{si}/Q_{ef} and Q_{di}/Q_{ef}), which are needed for calculations in equation (5), are calculated based on data from the USDA PSD (2012).

Mean estimates of the short-run export demand elasticity are on the right side of Table 6, but before discussing these, it is useful to examine Figures 1, 2, and 3. These illustrate the time path of our calculated elasticities when the ratios (Q_{si}/Q_{ef} and Q_{di}/Q_{ef}) are allowed to vary across years, according to their actual values. For example, Figure 1 shows 36 short-run and 36 long-run corn export demand elasticities that

Table 4. Soybean Demand and Supply Results

	Argentina	Brazil	China	Egypt	Japan	Mexico
Demand Equation						
Soybean price	−0.079 (0.159)	−0.093*** (0.032)	−0.107 (0.364)	−0.397 (0.507)	−0.079 (0.151)	−0.226 (0.249)
Wheat price	—	−0.128*** (0.039)	—	—	−0.157 (0.141)	−0.351* (0.203)
Income	−0.511** (0.208)	0.324*** (0.060)	0.841 (0.852)	−1.052*** (0.397)	—	1.864** (0.824)
Population	0.419*** (0.066)	−0.012 (0.088)	1.303 (2.147)	3.078*** (0.337)	−0.012 (0.020)	−1.666 (1.044)
Intercept	11.484*** (1.675)	6.342*** (0.185)	1.360 (4.907)	−2.591 (2.924)	6.655*** (0.083)	5.412*** (1.012)
Supply Equation						
Lagged soybean price	0.029 (0.020)	0.210 (0.190)	0.140 (0.188)	0.795 (0.683)	0.070 (0.244)	0.235** (0.091)
Lagged soybean production	−0.126 (0.250)	−0.786 (0.744)	0.477** (0.198)	0.531** (0.277)	0.771** (0.283)	0.557*** (0.125)
Lagged wheat price	—	0.039 (0.051)	−0.360 (0.222)	—	—	—
Trend	0.344*** (0.092)	0.039 (0.005)	0.080** (0.035)	0.223 (0.154)	−0.215 (0.139)	0.036 (0.045)
Intercept	9.994*** (2.202)	8.728*** (0.289)	4.946*** (1.786)	−4.099 (3.942)	1.347 (4.869)	1.075 (0.796)

Notes: Standard error is in parentheses. R^2 is not available for equations estimated with the restricted maximum likelihood estimator (–). The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

would have occurred between 1976 and 2011 if our fixed behavioral parameters are true and holding other things constant as well. The behavioral parameters, which were estimated for 1991–2011, are allowed to stay constant. Because the figures go back before this period, the reader must keep in mind that the demand, supply, and price transmission parameters corresponding to the 1991–2001 period are implicitly assumed to be representative to the larger historical period. We do this because pre-1991 had interesting periods of volatility in international markets, and we seek to make the comparisons over time. More importantly, we do this to make some direct comparisons to a previous study, Devadoss and Meyers (1990), who did a similar exercise for the late 1980s.

Perhaps the most noticeable aspect of the three figures is the occasional large increase in the elasticity of export demand during some years. For example, in the case of corn, the long-run export demand elasticity fell to –2.09 in 1994, much more elastic than its mean elasticity of –1.53 over this time period (Figure 1). In this instance, the denominator in the ratios Q_{si}/Q_{ef} and Q_{di}/Q_{ef} was smaller than average, that is, the U.S. exported less in this particular year. The absolute value of the long-run corn export demand elasticity became more inelastic than -1 in only a few cases, including 1980 (–0.99) and 1983 (–0.94). The overall trend for corn in the most recent years is that the export demand elasticity seems to be becoming more elastic.

Figure 2 reports the time path of the soybean short- and long-run export demand elasticities. In this case we are unable to go back as far as for corn or wheat since some data on the USDA PSD (2012) web site are incomplete or missing before 1990. The long-run export demand elasticity was as low as –2.64 in 1990, well beyond its mean elasticity of –1.77 over the 1990–2011 time period. The long-run elasticity was at its most inelastic in 2011 (–1.11). Most apparent is a general trend of inelasticity over this period.

Figure 3 reports the elasticity time path for the case of wheat. Most apparent is that both the short- and long-run elasticities appear to get slightly more elastic over much of the time period. This reflects an increase in the ratios Q_{si}/Q_{ef} and Q_{di}/Q_{ef} and thus declining prominence of

Table 5. Wheat Demand and Supply Results

	Algeria	Australia	Canada	Egypt	Japan	Mexico
Demand Equation						
ln wheat price	–0.017 (0.200)	–0.083 (0.366)	–0.207* (0.120)	–0.013 (0.498)	–0.010 (0.043)	–0.001 (0.033)
ln corn price	—	–0.501 (0.243)	0.075 (0.129)	—	—	—
ln income	—	0.193 (0.342)	0.477 (0.302)	0.448 (0.077)	(0.098)	0.719*** (0.079)
ln population	1.760*** (0.200)	0.258 (0.048)	–0.017** (0.024)	—	—	—
Intercept	–21.374*** (3.150)	10.243*** (1.932)	7.475*** (1.482)	6.316*** (0.548)	7.650*** (1.015)	6.884*** (0.410)
Supply Equation						
Lagged ln wheat price	0.492* (0.280)	0.071 (0.244)	0.008 (0.172)	0.289 (0.620)	0.356 (0.147)	0.172 (0.263)
Lagged ln wheat production	–0.066 (0.256)	0.665 (0.620)	0.257 (0.255)	0.151 (0.287)	0.767 (0.157)	0.642** (0.262)
Lagged ln corn price	—	–0.053 (0.578)	—	0.287 (0.454)	—	–0.123 (0.369)
Trend	—	—	–0.010 (0.008)	—	0.017 (0.007)	–0.001 (0.014)
Intercept	3.161 (3.005)	7.001** (2.994)	7.537** (3.035)	3.680*** (1.173)	–1.280 (3.306)	2.581 (2.526)

Notes: Standard error is in parentheses. R^2 is not available for equations estimated with the restricted maximum likelihood estimator (–). The asterisks ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

the United States in world wheat markets. There remains considerable variation from year to year, however, as was the case for corn and soybeans.

The wheat elasticity time path in Figure 3 can be compared directly to results obtained by Devadoss and Meyers (1990); in common with our study, they report the U.S. wheat export demand elasticity for the years 1982–1987. Their estimates are generated by a different type of model and methodological approach yet mimic the pattern of our results surprisingly well. Their result lies within the boundary provided by our short- and long-run results (Figure 3). Devadoss and Meyers (1990) do not explicitly distinguish between the short and long run, but the nature of their approach makes their estimates most comparable to our long-run estimates. The similarity of their results provides a measure of reassurance regarding the accuracy of both studies.

Export Demand Elasticities on Average

Now we turn back to Table 6, where we report mean and median elasticities for the different crops based on different time periods. For corn, the mean short-run elasticity over the entire 1976–2011 period is –1.04 (Table 6). This means that a 1% increase in the price of No. 2 yellow corn at U.S. Gulf ports is associated with a 1.04 percent fall in foreign import demand, holding all else equal. The mean long-run elasticity over the entire 1976–2011 period is –1.53. This is interpreted in the same way as the short-run elasticity, with the exception that this is the percentage change occurring somewhat longer after the initial price change (e.g., after 1 year), holding all else equal. When the sample is restricted to a relatively recent time period (2001–2011), these estimates fall to –1.11 and –1.64, respectively, signaling that wheat export demand has become slightly more price sensitive than it was historically.

In the case of soybeans, the mean short-run elasticity over the 1990–2011 period is –1.03 (recall that in contrast to corn and wheat, we are unable to calculate this before 1990 as a result of incomplete data for some of the countries involved). This means that a 1% increase in the price of No. 1 yellow soybeans at Chicago is associated with a 0.35 percent fall in foreign

Table 6. Export Demand Calculations

	Short-Run Price Transmission Elasticity	Long Run Price Transmission Elasticity	Demand Elasticity	Short-Run Supply Elasticity	Long-Run Supply Elasticity	Short Run	Long Run
Corn							
Argentina	0.848	0.994	–0.036	0.450	0.460		
Brazil	0.831	0.915	–0.251	0.104	0.137		
China	0.077	0.143	–0.003	0.060	0.186		
Egypt	0.204	0.293	–0.241	0.006	0.013		
Japan	0.900	0.955	–0.109	0.018	0.085		
Mexico	0.306	0.628	–0.197	0.598	0.613		
U.S. export demand elasticity							
Mean 1976–2011						–1.04	–1.53
Mean 1991–2011						–1.09	–1.62
Mean 2001–2011						–1.11	–1.64
Soybeans							
Argentina	1.089	1.094	–0.079	0.029	0.026		
Brazil	0.656	1.262	–0.093	0.210	0.118		
China	0.432	1.119	–0.107	0.140	0.268		
Egypt	0.135	0.216	–0.397	0.795	1.695		
Japan	0.460	0.858	–0.079	0.070	0.306		
Mexico	0.479	1.189	–0.226	0.235	0.530		
U.S. export demand elasticity							
Mean 1976–2011*						—	—
Mean 1991–2011						–1.01	–1.73
Mean 2001–2011						–0.90	–1.45
Wheat							
Algeria	0.667	1.367	–0.017	0.492	0.461		
Australia	0.409	0.513	–0.083	0.071	0.212		
Canada	0.389	0.907	–0.207	0.008	0.011		
Egypt	0.442	0.591	–0.013	0.289	0.340		
Japan	0.482	0.757	–0.010	0.356	1.528		
Mexico	0.490	1.342	–0.001	0.172	0.480		
U.S. export demand elasticity							
Mean 1976–2011						–0.37	–1.10
Mean 1991–2011						–0.40	–1.14
Mean 2001–2011						–0.45	–1.25

* Incomplete data on historical values of some other exporters prevents calculation of these results.

import demand with all else equal. The long-run estimate over this timeframe is –1.77. Looking at the latter part of this period, however, (2001–2011), these estimates are slightly more inelastic, at –0.90 for the short run and –1.45 for the long run (Table 6). This confirms what we saw graphically in Figure 2.

Finally, in the case of wheat, the mean short-run elasticity over the 1976–2011 period is –0.37. This means that a 1% increase in the price of No. 1 hard red winter wheat at Kansas City is associated with a 0.37 percent fall in foreign import demand with all else equal. The mean long-run elasticity over the 1976–2011 period is –1.10. Over the 2001–2011 time period these numbers are slightly more elastic, at –0.45 and –1.25, respectively. This confirms our earlier observation that there seems to be more elasticity in the wheat market over time.

Comparison with Previous Studies

Our estimates can be placed in context by comparing them to previous estimates in the

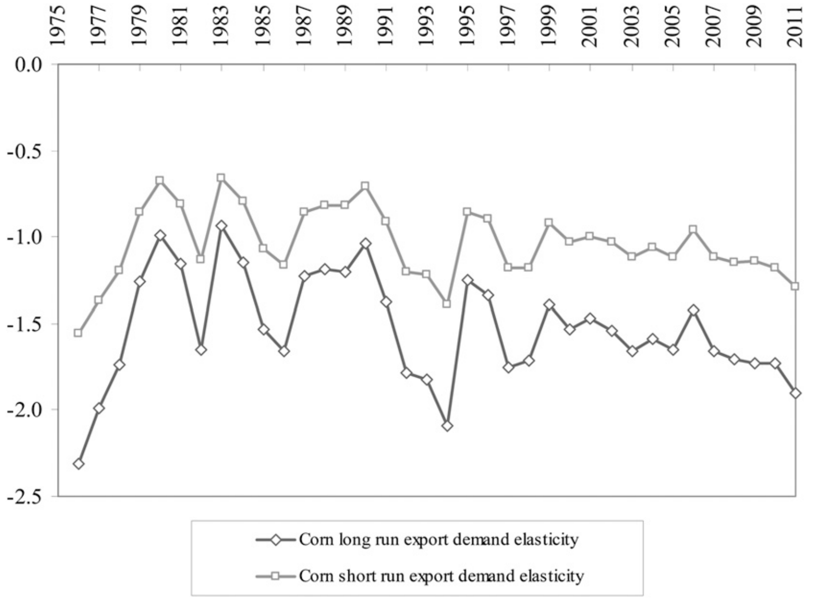

Figure 1. Time Path of U.S. Corn Export Demand Elasticity

literature. Three earlier results concerning the corn export demand elasticity are reported in Table 7. Bredahl, Meyer, and Collins (1979) report (long-run) corn elasticities varying from –0.09 for their most restrictive case (no price transmission) to –3.13 for a hypothetical free trade case (perfect price transmission). Their preferred estimate is –1.31, which lies between our short- and long-run estimates of –1.11 and –1.64 for the 2001–2011 time period. Our estimates also straddle Gardiner's (1986) long-run estimate of –1.18. Both of our estimates are somewhat more elastic than Chambers and Just's (1981) short- and long-run estimates of –0.47 and –0.63. This study used a somewhat different definition of the export demand elasticity,

Figure 2. Time Path of U.S. Soybean Export Demand Elasticity

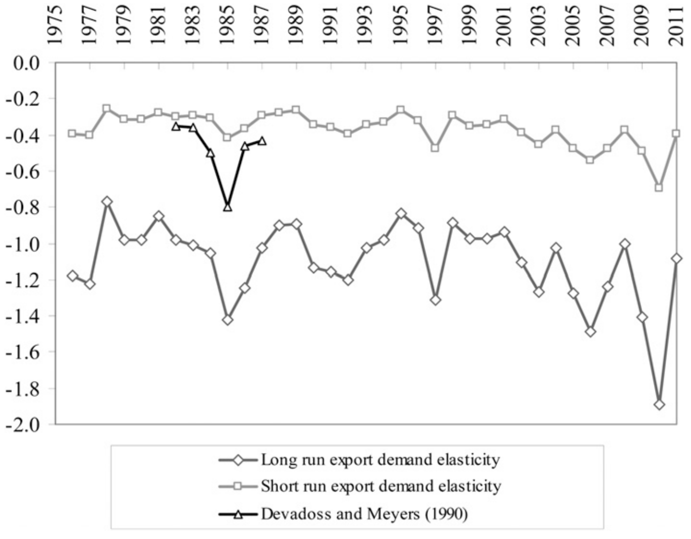

Figure 3. Time Path of U.S. Wheat Export Demand Elasticity

different estimation techniques and a much earlier sample period (before the international importance of suppliers such as Argentina).

A comparison of soybean export demand elasticities is made in the middle of Table 7. Our short- and long-run estimates of –0.90 and –1.45 straddle that of the –1.27 reported in Miller and Paarlberg (2001), for example. Our estimates are somewhat more elastic than that found by Bredahl, Meyers, and Collins (1979)

Table 7. Comparison of Export Demand Estimates

Study	Period	Short-Run Elasticity	Long-Run Elasticity
Corn			
This study	2001–2011	–1.11	–1.64
Bredahl, Meyers, and Collins (1979)	1972/1973–1975/1976	–	–1.31
Chambers and Just (1981)	1969–1977	–0.47	–0.63
Gardiner (1986)	1967–1980	—	–1.18
Soybeans			
This study	2001–2011	–0.90	–1.45
Johnson (1977)	1970 base	—	–2.80
Bredahl, Meyers, and Collins (1979)	1972/1973–1975/1976	—	–0.47
Chambers and Just (1981)	1969–1977	–0.20	–0.29
Miller and Paarlberg (2001)	1964–1999	—	–1.27
Wheat			
This study	2001–2011	–0.45	–1.25
Johnson (1977)	1970 base	—	–6.72
Bredahl, Meyers, and Collins (1979)	1972/1973–1975/1976	—	–1.67
Paarlberg (1983)	1960–1975	—	–1.82
Johnson et al. (1985)	1985	–0.16	“near –1.0”
Meyers and Helmar (1986)	1986	–0.11	—
Tyers and Anderson (1988)	1988	–1.00	–2.90
Miller and Paarlberg (2001)	1960–1999	—	–3.83
Miller and Paarlberg (2001)	1960–1984	–2.43	–2.33
Miller and Paarlberg (2001)	1985–1999	–1.65	–1.45

and by Chambers and Just (1981). They are somewhat more inelastic than the –2.80 finding reported in Johnson (1977). It is important to emphasize that Johnson (1977) assumes that there is perfect price transmission between the U.S. market and all foreign markets. As discussed previously, there is a number of reasons why there is not perfect price transmission. For example, U.S.-foreign price linkages may be sticky as a result of the insulating effects of border policies, menu or catalog pricing (i.e., prices are not renegotiated continuously), oligopoly market power, imperfections in future price markets, and imperfect exchange rate pass-through. Our estimates may be more credible because they account implicitly for such factors.

Finally, the bottom of Table 7 provides a comparison of our wheat export demand elasticities and those of several previous studies. Johnson (1977) finds a long-run elasticity of export demand of –6.72. This is much more elastic than the –1.25 estimate that we found in the long run. As discussed previously, this is probably largely driven by his implicit assumption of perfect price transmission. Our long-run wheat export demand elasticity (–1.25) is somewhat more in line with the –1.67 estimate of Bredahl, Meyers, and Collins (1979) for the early 1970s, and the –1.82 estimate of Paarlberg (1983) for the 1960–1975 time period. It also fits well with the –1.45 estimate of Miller and Paarlberg (2001) for the 1985–1999 timeframe, which is more recent than that of many of the other studies, if still not as recent as ours.

Conclusions and Caveats

We provide new estimates of the U.S. foreign export demand elasticity for corn, soybeans, and wheat. Unlike most previous studies, we econometrically estimate the price transmission, demand elasticities, and supply elasticities that underlie our export demand calculations.

During the 2001–2011 period, the export demand elasticity for U.S. corn, soybeans, and wheat averaged –1.11, –0.90, and –0.45 in the short run, and –1.64, –1.45, and –1.25 in the long run. Export demand elasticities for corn and wheat were slightly more elastic during 2001–2011 than in previous years. By contrast, the export demand elasticity for soybeans was slightly more inelastic during 2001–2011 than in previous years.

Another objective was to reconcile our results with those of earlier studies. This is difficult to do because many previous studies have a different definition of the export demand elasticity. Unlike most, we account for the excess supply of competing producers (which increases elasticity) and for imperfect price transmission (which decreases it). Perhaps for these reasons, our results typically straddle the various estimates of earlier studies. In any case, our results should reduce some of the uncertainty that has long surrounded this elusive parameter.

It is important to recognize the tradeoffs we have had to make. First, we are limited to a short data series to estimate price transmission elasticities and the supply and demand for commodities in foreign countries. Second, we do not have good time series data on all the foreign country supply and demand shifters that we might have ideally included in the supply and demand regressions. Third, we do not observe actual internal prices, and there of course could be measurement errors in the underlying price data.

Despite these shortcomings, we have been able to shed light on a much-debated parameter that is of particular importance as the United States sets about developing a new farm bill. The estimates may facilitate back-of-the-envelope policy analysis, enabling the analyst to make predictions about exports based on changing economic policies or weather shocks, for example. Our estimates should also provide a useful reference point for researchers tasked with the parameterization of large economic simulation models that have an international component.

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