Goodwill and Export Promotion Dynamics*

by

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Abstract. Federal subsidies for non-price export promotion of farm products have been criticized on the grounds that they merely substitute taxpayer dollars for private promotional expenditures. This "displacement hypothesis" is tested by estimating export demand and advertising-goodwill relationships for farm products using time series data for 1975-2008. Results suggest an added dollar of subsidy cuts industry expenditures for export promotion by between 79 cents and 87 cents. Thus, the displacement hypothesis appears to have some validity. A $1 increase in subsidy is estimated to increase total spending (government plus industry) on export promotion by $1.05 in the short run; an upper-bound estimate of the long-run effect is $1.25. Thus, the Foreign Market Development and Market Access programs operated by the U.S. Department of Agriculture appear to have had only a modest effect on goodwill creation. Given recent estimates showing that the programs have substantial consumer costs, it might be time to let them die.

Key words: export promotion, goodwill, international trade, subsidy

JEL classification: D61, F14, Q11

Goodwill and Export Promotion Dynamics

Most developed-country governments encourage industry to develop or expand foreign markets through non-price promotion activities. For example, a 2002 survey of 29 countries found that total expenditures for export promotion of agricultural, forestry, and fishery products exceeded $1.5 billion, of which $465 million, or 30 percent, represented taxpayer dollars (table 1). In the United States, the federal government has subsidized non-price export promotion of agricultural products since 1955, with outlays in recent years exceeding $225 million (Dwyer 2010). As noted by Love, Porras, and Shumway (2001), non-price promotion activities are permissible under the Uruguay Round of the General Agreement on Tariffs and Trade, which may explain their popularity as a policy instrument. Yet the programs remain controversial. Proponents argue that export promotion is a type of public good in that “the ‘new’ markets found would become open for other producers and exporters” (McCalla and Valdes 1997, p. 8). The free-rider problem implies that private-sector investment would be below the social optimum *sans* government aid. Opponents counter that such aid displaces private investment, and is inefficient (e.g., Stansel 1995).

The purpose of this research is to test the displacement hypothesis. Specifically, we address whether subsidies enlarge total spending for export promotion of farm products, or merely encourage industry to substitute public dollars for private promotional expenditures, a major contention of the U.S. Government Accountability Office (GAO). The U. S. Department of Agriculture (USDA) provides subsidies through two programs: the Foreign Market Development Program (FMD) and the Market Access Program (MAP). MAP funds are aimed at processed foods and “high-value” farm products (e.g., almonds, raisins, salmon, wine) and require dollar-for-dollar matching; FMD funds are aimed at bulk products (e.g., corn, cotton, soybeans, wheat) and require as little as one dollar of industry money per nine dollars of government money. In a series of reports, the GAO alleged that MAP funds went disproportionately to large and experienced firms that had little need of export assistance, and who would simply substitute federal dollars for private dollars (USGAO 1999). The GAO contended that redirecting the monies to small and "new-to-export" firms that had difficulty accessing international markets would improve the efficiency of the programs.

The scholarly literature is replete with studies that evaluate the effects of export promotion on market demand and producer returns (e.g., Rusmevichientong and Kaiser 2009 and references therein). However, no study to our knowledge has examined the effects of subsidies on industry spending. The closest related research is a study by Jakus *et al*. (2003) that examined whether firm size and export experience matter in the conversion of MAP funds into export sales. The study rejects GAO's contention that redirecting MAP monies to smaller, less export-experienced firms would enhance program efficiency. In a related study, Adams *et al*. (1997) found that the USDA did indeed tend to ignore small business assistance in its allocation of program funds, placing greater emphasis on the ability of the funds to expand export sales. Beyond these studies, little is known about how the programs work. The displacement hypothesis, in particular, has not been tested.

In their classic study, Nerlove and Arrow (1962, p. 130) argued that advertising is best modeled as a capital asset, which they called "goodwill," that depreciates over time, but that can be replenished by increasing current advertising outlays. The distinction between advertising outlays (a flow variable) and goodwill (a stock variable) is germane because, in a dynamic setting, the relationship between the flow and stock variables is bi-directional (Sethi 1977; Feichtinger, Hartl and Sethi 1994). Specifically, Luhta and Virtanen (1996, p. 2085) show that when advertising budgeting decisions are dependent on sales, the amount of advertising in the current period is a function of the current level of goodwill, which, in turn, is a function of current and past advertising outlays. This implies that in testing the relationship between industry expenditures on export promotion and the subsidy offered by the USDA, it is necessary to have a measure of goodwill. Thus, a secondary objective of this study is to estimate the level of goodwill associated with U.S. investments in export promotion of farm products. The objective is achieved by estimating the parameters of a Cobb-Douglas type production function for goodwill using time series data for the period 1975-2004. The estimated parameters are then used to reconstruct the levels of goodwill over the period 1980 – 2008, which are used in an advertising-goodwill relationship to test the displacement hypothesis.

The paper has five sections. The next section presents the model. We then discuss estimation procedures and results. The hypotheses are then tested, and estimates of the effect of the subsidy on total spending are presented. The final section provides concluding comments.

Model

The displacement hypothesis is tested using the two-equation system:

(1) $$A_t = a_0 + a_1 AG_t + a_2 G_t + a_3 P_t + a_4 SDR_t + \varepsilon_{1t}$$

(2) $$AI_t = b_0 + b_1 AG_t + b_2 G_t + b_3 P_t + b_4 SDR_t + b_5 \{\min(0, \Delta AG_t)\} + \varepsilon_{2t}$$

where A_t, AG_t, and AI_t are total expenditures, government expenditures, and industry expenditures on export promotion of farm products in year t; G_t is the level of goodwill in year t; P_t is real price of U.S. farm products in year t; SDR_t (for Special Drawing Rights) is the value of the U.S. dollar against a market basket of world currencies in year t; $\Delta AG_t = AG_t - AG_{t-1}$ is the change in government expenditures between adjacent years, and ε_{1t} and ε_{2t} are random error terms. The price and exchange-rate variables are specified in equations (1) and (2) to test whether promotion is pro-cyclical or countercyclical with respect to these variables. For example, Armbruster and Nichols (2001) have proposed that subsidies for export promotion be increased during periods of a strong dollar to counter the negative effects of a strong currency on export demand. Because an increase in XR implies a stronger dollar, this would imply $a_4 > 0$. Hadar (1971, p. 128) shows that for a competitive firm to engage in advertising, the difference between price and average variable cost must exceed the per-unit cost of advertising, i.e., $p - AVC \geq \bar{a}/\bar{q}$ where $\bar{q}$ is the quantity sold at advertising expenditure level $\bar{a}$. From this expression one would expect a positive relationship between advertising expenditure and price, i.e., $b_3 > 0$.

The $\min(0, \Delta AG_t)$ term is included in equation (2) to test for a “flypaper effect”. A flypaper effect occurs if, during periods of falling government expenditures, industry increases its own spending in order to maintain a constant level of investment. In other words, export promotion expenditures are “sticky” downward, a phenomenon suggested by the public finance literature (e.g., Hines and Thaler 1995; Knight 2002). The effect holds if $b_5 > 0$.

The displacement effect implies that $a_1 = 1$ and $b_1 = -1$. That is, a one dollar increase in government expenditures decreases industry expenditures by one dollar and thus increases total expenditures by one dollar. A formal test of the displacement hypothesis is:

(3a) H_N: $a_1 + b_1 = 0$

(3b) H_N: H_N not true.

Hypothesis (3) is a test of linear restrictions and hence a *t*-statistic is appropriate.

To implement the test, the level of goodwill over the sample period must be measured. For this purpose, we follow Doganoglu and Klapper (2006) and posit a Cobb-Douglas type function:

$$G_t = (1 + A_t)^{1-\lambda}(G_{t-1})^{\lambda} \tag{4}$$

where G_t is the stock of goodwill in year t, A_t is total export promotion expenditures in year t, and $0 \leq \lambda < 1$ is the goodwill retention parameter. In this formulation, if the industry ceases to advertise goodwill depreciates at a rate proportional to the level of goodwill in the preceding period.

Following Doganoglu and Klapper, the retention parameter was estimated by first specifying a demand equation in constant elasticity form:

$$\begin{aligned} \ln(MS_t) = {} & \beta_0 + \beta_P \ln(P_t) + \beta_{PS} \ln(PS_t) + \beta_{XR} \ln(XR_t) \\ & + \beta_Y \ln(Y_t) + \beta_G \ln(G_t) + \mu_t \end{aligned} \tag{5}$$

where $MS = PQ_x / Y$ is the share of foreign income spent on U.S. exports of farm products; P and Q_x are the price and quantity of U.S. farm products exported; Y is foreign income; PS is the price of substitutes; XR is an agricultural trade-weighted exchange rate; G is goodwill; and μ is a random error term. The XR variable is included to test whether foreign buyers' response to exchange-rate movements differs from their response to price movements, as suggested by Chambers and Just (1981). Equation (5) is identical in form to the model specified in Kinnucan and Cai (2010).

Performing a Koyck transformation on equation (5), i.e., lagging the equation one period, multiplying both sides by λ, and then subtracting the lagged equation from the original, yields:

$$\begin{aligned} \ln(MS_t) - \lambda\ln(MS_{t-1}) = \beta_0(1-\lambda) + \beta_P(\ln(P_t) - \lambda\ln(P_{t-1})) \\ + \beta_{PS}(\ln(PS_t) - \lambda\ln(PS_{t-1})) + \beta_{XR}(\ln(XR_t) - \lambda\ln(XR_{t-1})) \\ + \beta_Y(\ln(Y_t) - \lambda\ln(Y_{t-1})) + \beta_G(\ln(G_t) - \lambda\ln(G_{t-1})) + \mu_t - \lambda\mu_{t-1} \end{aligned} \tag{6}$$

The unobserved goodwill terms in equation (6) are eliminated by substituting equation (4) to yield:

$$\begin{aligned} \ln(MS_t) = \alpha + \beta_P(\ln(P_t) - \lambda\ln(P_{t-1})) + \beta_{PS}(\ln(PS_t) - \lambda\ln(PS_{t-1})) \\ + \beta_{XR}(\ln(XR_t) - \lambda\ln(XR_{t-1})) + \beta_Y(\ln(Y_t) - \lambda\ln(Y_{t-1})) \\ + \beta_G(1-\lambda)\ln(1+A_t) + \lambda\ln(MS_{t-1}) + \xi_t \end{aligned} \tag{7}$$

where $\alpha = \beta_0(1-\lambda)$ and $\xi_t = \mu_t - \lambda\mu_{t-1}$ is an autoregressive error term. Thus, market share is a function of traditional demand shifters, current advertising expenditures, and lagged market share. If the estimated coefficient of lagged market share is zero, the goodwill function reduces to $G_t = (1+A_t)$ and the stock of goodwill is depleted in one year. In this instance, the distinction between advertising stocks and flows vanishes and equations (1) and (2) are modified to exclude the *G* variable. If the estimated value of λ is positive, the goodwill levels over the sample period can be reconstructed using equation (4) and an assumed value for the initial level of goodwill.

Adding a trend variable and interaction term to equation (7), the empirical model estimated is:

$$
\begin{aligned}
(8) \qquad & \ln(X_t^{US} \big/ X_t^{W}) = \alpha + \beta_P(\ln(P_t^{US} / DEFL_t) - \lambda \ln(P_{t-1}^{US} / DEFL_{t-1})) + \\
& \beta_{PS}(\ln(P_t^{C}) - \lambda \ln(P_{t-1}^{C})) + \beta_{XR}(\ln(XR_t) - \lambda \ln(XR_{t-1})) + \\
& \beta_Y(\ln(X_t^{W} \big/ DEFL_t) - \lambda \ln(X_{t-1}^{W} \big/ DEFL_{t-1})) + \beta_G(1-\lambda)\ln(1+AD_t) + \\
& \tilde{\beta}_G(1-\lambda) \cdot D_t \cdot \ln(1+AD_t) + \beta_T TREND_t + \lambda \ln(X_{t-1}^{US} \big/ X_{t-1}^{W}) + \xi_t
\end{aligned}
$$

where X_t^{US} is the nominal value of U.S. agricultural exports in year *t* in U.S. dollars; X_t^{W} is the nominal value of world imports of agricultural products in year *t* in U.S. dollars; P_t^{US} is the unit-value of U.S. bulk agricultural exports in year *t* in U.S. dollars; $DEFL_t$ is a GNP deflator for the world less the United States; P_t^{C} is a Stone index of real trade-weighted exchange rates for US competitors' agricultural exports; XR_t is a world U.S. agricultural trade-weighted real exchange rate; $AD_t = A_t \cdot SDR_t / DEFL_t$ is real promotion expenditures adjusted for the strength of the U.S. dollar (Dwyer 1994); $TREND_t$ is a linear trend variable; and D_t is a dummy variable that assumes the value of one if $\Delta A_t > 0$ and zero otherwise. A precise empirical definition of all variables, including sources, is given in appendix table 1.

The interaction term $D_t \cdot \ln(1+AD_t)$ is included in equation (8) to test whether the market response to advertising increases during periods of rising expenditures, as is suggested by the "pulsing" literature (Little 1979; Simon 1982; Feinberg 1992; Feichtinger, Hartl and Sethi 1994). The pulsing hypothesis implies $\tilde{\beta}_G > 0$, which can be tested with a simple *t*-statistic. The trend variable is included to account for changes in tastes and preferences for U.S. farm exports and the effects of other unmeasured factors.

The unit value of U.S. bulk farm exports serves as a proxy for the U.S. price, and an exchange rate index reflecting competitors' agricultural export prices serves as a proxy for the substitute price. Replacing foreign income with world import expenditures on farm products in essence converts the model to a conditional demand specification. Specifically, world demand for farm exports is implicitly assumed to be weakly separable from all other goods. This assumption, coupled with the two-stage budgeting hypothesis, implies that the price and income elasticities estimated from equation (8) are properly interpreted as conditional elasticities (Phlips, 1990, pp. 71-77).

Labeling equation (8) "Model A," three restrictive forms were estimated to assess the sensitivity of parameter estimates to economic hypotheses:

Model B: $\tilde{\beta}_G = 0$ (no pulsation effect)

Model C: $\beta_Y = 0$ (homothetic preferences)

Model D: $\beta_Y = \beta_P = 0$ (homothetic preferences and unitary demand elasticity)

In testing the restrictive forms, Model A is treated as the maintained hypothesis. The model was estimated using annual data for the period 1975-2004. One observation is lost due to the goodwill specification, so the effective sample period is 1976-2004.

Estimation Procedures and Results

Augmented Dickey-Fuller tests showed all variables to be non-stationary in levels, but stationary in first differences at the 5% level or better. The Johansen test indicated the *I*(1) variables are cointegrated.[1] Based on these results, equation (8) was estimated using the Generalized Methods of Moments (GMM) estimator with P^{US} and *AD* treated as endogenous variables. Thus, we implicitly assume that the United States is a sufficiently large exporter of farm products that shocks in export demand affect price. *AD* is treated

[1] Test results are in an appendix available from the authors upon request. All estimation is done using the econometric software *EViews*, Version 7.

as endogenous because demand-side shocks affect equilibrium quantity, which, in turn, affect the funds available for promotion through industry checkoff programs (for details, see Forker and Ward 1993). Hence, $COV(P_t^{US}, \xi_t) \neq 0$ and $COV(AD_t, \xi_t) \neq 0$ are treated as maintained hypotheses, and P^{US} and AD are instrumented using D, $TREND$, $DEFL$, P^C, XR, X^W, P_{-1}^{US}, AD_{-1}, and the remaining lagged terms given in equation (8). This resulted in 13 instruments. Because the model contains 9 parameters, it is over-identified. The Sargan statistic is used to test whether the overidentifying restrictions are valid. The GMM estimates are obtained using EVIews 7 (2009). A constant term was included in the instrument list, the estimation weighting matrix was set to "HAC (Newey-West)," and weight updating was set to the sequential one-step iterative option. Significance is determined using the $p \leq 0.05$ level unless indicated otherwise.

Estimation results are satisfactory in that the estimated coefficients have the correct signs, and most are significant (table 2). The *J*-statistic indicates the over-identifying restrictions are valid, i.e., all models pass the Sargan test. Testing Models B, C, and D against Model A, the hypothesis that pulsing is unimportant (Model B) is firmly rejected (*p*-value < 0.0000). Similarly, the joint hypothesis that export demand is unitary price elastic and preferences are homothetic (Model D) is rejected (*p*-value = 0.050). However, the hypothesis that preferences are homothetic (Model C) can be rejected at the 7% level, but not the 5% level. Thus, homothetic preferences, a central feature of Armington's (1969) trade model, receive some support from our data. Because the restrictive models are rejected at normal probability levels, the remaining discussion will focus on Model A.

Demand Elasticities

The estimated value of β_P is not significant, which implies the own-price elasticity of demand for U.S. farm exports is -1. The estimated exchange-rate elasticity β_{XR} is -0.45 is significant. The hypothesis that β_{XR} = -1 is rejected at the p = 0.035 level. Hence, the Chambers-Just hypothesis that agents' responses to price movements differ from their responses to exchange-rate movements is affirmed.

The estimated cross-price elasticity β_{PS} is 0.40 and is significant. This result suggests domestic and foreign farm products are substitutes in international trade.

The estimated value of β_Y is 0.33 and is significant at the p = 0.074 level. The implied expenditure elasticity is 1.33, which suggests U.S. farm products are superior goods in international trade is expenditure elastic. This finding is consistent with the growing share "high-value" farm products (e.g., almonds, grapes, salmon, wine) in the U.S. export mix, which, over the sample period, increased from 28% to 68% of total export value.

The estimated coefficient of the trend variable is -0.016 and is significant. This suggests that in the absence of changes in relative prices, exchange rates, income, and promotion, the U.S. share of the world market for farm products would have declined at a rate of 1.6% per year. The actual average annual rate of decline over the sample period was 0.24%, which suggests that demand growth associated with the economic variables in the model largely offset the negative effects of omitted trend-related factors.

The estimated goodwill elasticity is 0.147 and is significant. Using the same data but a different specification for goodwill, Kinnucan and Cai (2010) placed the long-run promotion elasticity at 0.189. Estimating separate equations for bulk and high-value agricultural products, two recent consulting reports obtained estimates of the long-run advertising elasticity that ranged from 0.14 to 0.20 (Global Insight 2007; HIS Global Insight 2010). The consistency in estimates across a range of studies and model specifications increases confidence that the estimated demand effects are real. The 95% confidence interval for the goodwill elasticity estimated from Model A is [0.015, 0.278]. This suggests a 1% increase in promotion expenditures shifts the export demand curve in the quantity direction by between 0.015% and 0.28% when measured from the initial equilibrium point.

The interaction term is significant (t-ratio = 4.0), but the estimated effect is tiny. Specifically, $\hat{\tilde{\beta}}_G = -0.008$, which means the goodwill elasticity declines from 0.147 to 0.139 when promotion expenditures are rising. (Over the sample period promotion expenditures rose in 62% and fell in 38% of the years.) Although pulsing appears to be *empirically* unimportant, it is *statistically* important in that its exclusion causes the estimated goodwill elasticity to increase from 0.147 to 0.329 (compare models A and B in table 2).

Retention Parameter and Goodwill Measurement

Turning to λ, the key parameter in this study, its estimated value is 0.44 and is significant. Hence, the hypothesis that past levels of goodwill are independent of the current stock is rejected. By way of comparison, Doganoglu and Klapper's (2006) estimate of λ based on weekly store data for three brands of a liquid detergent sold in Germany ranges from 0.70 to 0.90. Thus, our estimate of 0.44 suggests that the "brand equity" for U.S. farm products in export markets depreciates at a faster rate than the brand equity for liquid detergents in Germany.

With an estimate of λ in hand, the levels of goodwill over the sample period were reconstructed by solving equation (4) to yield:

$$G_t = \left[\prod_{i=0}^{t-1} (1 + A_{t-i})^{\lambda^i} \right]^{(1-\lambda)} (G_0)^{\lambda^t} \tag{9}$$

where G_0 is the level of goodwill in the initial period. Luhta and Virtanen (1996, pp. 2085-86) show that if the marginal cost of goodwill is constant, the equilibrium level of goodwill under conditions of constant promotion expenditure is $G^* = A / r$ where r is the goodwill depreciation rate. Using this formula, we set $G_0 = \overline{A} / 0.33$ where $\overline{A} = \$90{,}990{,}478$ is the average level of promotion in the first three years of the sample (1975 – 1977) expressed in constant (2005) dollars, and 0.33 is the assumed depreciation rate. This yielded $G_0 = \$275{,}728{,}722$.

Since the estimate is crude, to minimize error we deleted the first five observations of the reconstructed goodwill series. Specifically, at $t = 6$ the numerical value for $(G_0)^{\lambda^t}$ is \$1.15. Hence, deleting the first five observations effectively eliminates the influence of the initial value on the data series. Consequently, equations (1) and (2) are estimated using data for the period 1980 through 2008. (The series used to estimate the demand equation terminates in 2004 because of missing values for the P^C variable after 2004.)

Hypothesis Tests

The regression estimates for equations (1) and (2) used to test the displacement and flypaper hypotheses are presented in table 3. The equations were estimated in first differences to account for a common trend in real expenditures over the sample period (figure 1). The equations were estimated as a system using Generalized Method of Moments with A, AI, AG, ΔAG, G, and P^{US} treated as endogenous. The instrument list includes the lagged values of these variables, plus the current and lagged values of *DEFL*, *SDR*, *XR*, and *FCPI* (see appendix table 1 for variable definitions). The equations are over-identified; thus a Sargan test was performed to determine whether the over-identifying restrictions are valid. The test failed to reject the null hypothesis that the restrictions are valid. A first order auto-regressive term was included in each equation to account for serial correlation. Model B imposes the restriction $a_1 + b_1 = 0$. For brevity, discussion will focus on the unrestricted model (Model A) unless indicated otherwise.

Regression results are satisfactory in that the *DW* statistic is close to 2 in each equation, and most of the estimated coefficients have the expected sign and are significant. The R^2 is 0.88 for the total expenditure equation and 0.31 for the industry expenditure equation. Bearing in mind that the models are estimated in first differences, their explanatory power appears adequate. The estimated constant term in each equation is positive and significant, which suggests that real expenditures for export promotion would have grown over the sample period even in the absence of changes in the variables specified in the models. (The estimated trend growth rate is 2.2% per annum for total expenditures and 5.6% per annum for industry expenditures.)

Farm price has a negative effect on industry expenditures (t-ratio = -3.33), but no effect on total expenditures. This suggests industry invests more in promotion when farm prices are low than when farm prices are high. Stated differently, industry appears to use promotion strategically to strengthen prices during periods of market glut.

Exchange rate has a positive effect on total expenditures (t-ratio = 4.57), but no effect on industry expenditure. This suggests the U.S. government increases its level of investment in export promotion when the U.S. dollar is strong and decreases its investment when the dollar is weak. Thus, Armbruster and Nichols (2001) proposal that subsidies for export promotion be increased during periods of a strong dollar to counter the negative effects of a strong currency on export demand appears be consistent with behavior.

The estimated coefficient of min(0, ΔAG) is positive, but is only marginally significant (t-ratio = 1.63). Hence, there is some evidence for a flypaper effect with respect to industry expenditures for export promotion. However, the effect is small, on the order of about a 10 cent increase in industry expenditure for a one dollar decline in government expenditure, *ceteris paribus*.

In Luhta and Virtanen's (1996, pp. 2087-88) model, advertising expenditures at first increase and then decrease with the level of goodwill, with the equilibrium level of advertising occurring on the downward sloping portion of the curve. This implies that if industry is optimizing its level of advertising, the slope of the expenditure function with respect to goodwill should be negative. In the present analysis, the slopes of the expenditure functions with respect to goodwill are uniformly positive, with estimates ranging from 0.47 to 0.55 depending on model specification. This suggests promotion expenditures are below the level that would maximum producer welfare, a result consistent with the bulk of the empirical literature as summarized in Rusmevichientong and Kaiser (2009).

Turning to *AG*, the key policy variable, its estimated coefficient is 0.69 in the total expenditure equation (*t*-ratio = 4.45) and -0.98 in the industry expenditure equation (*t*-ratio -4.45). Imposing the restriction that the two estimates sum to zero (Model B) yields $\hat{a}_1 = -\hat{b}_1 = 0.83$. A chi-square test of this restriction fails to reject the null hypothesis that models A and B are statistically equivalent (table 4). Thus, the displacement hypothesis appears to be valid. Specifically, an added dollar of public expenditure reduces private expenditure by 83 cents. The 95% confidence interval for this estimate is [0.79, 0.87], which suggests the displacement at the margin could be as high as 87 cents.

Total Effect of Subsidy

The total effect of the subsidy on spending can be computed from the derivative

$$(10) \qquad \frac{dA}{dAG} = \frac{\partial A}{\partial AG} + \frac{\partial A}{\partial G}\frac{\partial G}{\partial A}\frac{\partial A}{\partial AG}$$

where $\partial A/\partial AG$ is the partial (or direct) effect of the subsidy, i.e., its effect on spending when goodwill is held constant. The compound term in equation (10) is the indirect effect of the subsidy, i.e., its effect on spending that occurs as a result of the induced increase in goodwill. Factoring the partial effect gives

$$(11) \qquad \frac{dA}{dAG} = \frac{\partial A}{\partial AG}\left(1 + \frac{\partial A}{\partial G}\frac{\partial G}{\partial A}\right).$$

Whether the total effect is larger or smaller than the partial effect depends on the sign of $\partial A/\partial G$, as the other partial derivatives in equation (11) are positive by definition. In

Luhta and Virtanen's (1996) model, $\partial A/\partial G$ is negative in equilibrium, meaning that the firm continues to invest in goodwill until there are diminishing returns.

In the present analysis $\hat{a}_2 = \partial A/\partial G = 0.51$ (table 3, Model B), which implies increasing returns, and thus suboptimal investment. That is, despite the subsidies, total spending on export promotion of farm products appears to be below the level that would maximize producer welfare. Since $\hat{a}_2 = \partial A/\partial G > 0$, the estimated partial effect $\hat{a}_1 = \partial A/\partial AG = 0.83$ (table 3, Model B) understates the total effect.

To compute the total effect, an estimate of $\partial G/\partial A$ is needed. Referring to equation (4), a short-run estimate of this derivative can be obtained from the estimated replenishment elasticity $1-\hat{\lambda} = 0.56$. Specifically, evaluating this elasticity at the mean data points for *G* and *A* over the 1980-2008 sample period yields $(\partial G/\partial A)^{SR} = 0.55$. Substituting this value and the aforementioned values of the other derivatives into equation (11) gives:

$$\left(\frac{dA}{dAG}\right)^{SR} = 0.83(1+0.51(0.55)) = 1.06. \tag{12}$$

A one dollar increase in subsidy increases total spending by \$1.06. This is a short-run (one year) estimate, and thus sets the lower limit on the total effect. An upper-limit estimate can be obtained by setting $(\partial G/\partial A) = 1$ to yield:

$$\left(\frac{dA}{dAG}\right)^{MAX} = 0.83(1+0.51(1.00)) = 1.25. \tag{13}$$

Thus, the last dollar spent on subsidy is estimated to increase total spending on export promotion by *at most* \$1.25.

Concluding Comments

Federal subsidies for non-price export promotion of farm products in the United States have proved controversial, in part because a significant share of the funds have gone to large agribusiness firms that have little apparent need for export assistance. In addition to equity, this has raised concerns about efficiency. One aspect of the efficiency issue is whether the subsidies merely result in industry substituting taxpayer dollars for private promotional expenditures.

Testing this "displacement hypothesis" using annual time series data for the period 1980-2008, we find that an added dollar of subsidy cuts industry expenditures by between 79 cents and 87 cents. This suggests the displacement hypothesis has some validity. That is, while the alleged substitution is not one-for-one, a significant share of the added taxpayer dollar appears to be lost due to industry's propensity to replace private dollars with government dollars. Indeed, the estimated "total effect" of the subsidy has a maximal value of 1.25, which means the last dollar of subsidy generates at most $1.25 in additional spending. The upshot is that the Foreign Market Development and Market Access programs operated by the U.S. Department of Agriculture appear to have been relatively ineffectual in encouraging industry to expand their investments in export promotion.

Although the subsidies appear not to have enlarged total spending to any extent, this does not necessarily mean they should be eliminated. That depends on their marginal benefit, and opportunity cost. Our econometric estimates suggest industry is on the rising portion of the advertising-goodwill relationship, which suggest current spending levels are below the level that would maximize producer welfare. Marginal gains to producers from export promotion have been shown to be high, on the order of between $19 and $39 per additional dollar invested (Kinnucan and Cai 2010). The same study estimates that the social gain (producer plus consumer surplus) from incremental promotion expenditures is substantially smaller, on the order of between -$34 and $10.[2] Thus, if the goal is to increase producer welfare, even small induced increases in total spending might be sufficient to warrant continuation of the subsidies. However, if consumer and opportunity costs are deemed important, the best policy might be to let the subsidies die.

A caveat in interpreting our findings is that they are predicated on the model and estimation procedures used. It is possible that a different model or estimation procedures could alter inferences. Given the importance of the displacement hypothesis for policy efficacy, further testing is warranted.

[2] Negative gains accrue because increased export demand raises price in the domestic market, which represents a loss to domestic consumers. The loss is exacerbated when subsidies for export promotion cause industry to divert funds from domestic market promotion. In this instance, the reduced demand in the domestic market causes an added loss in consumer welfare if the promotion provides useful information or signals product quality. For details, see Kinnucan and Cai (2010).

Table 1. Expenditures for Non-Price Export Promotion of Agricultural, Forestry, and Fishery Products in 29 Surveyed Countries, 2002[a]

Country	Government Expenditures	Industry Expenditures	Total Expenditures
	---------------------------- Million US Dollars -------------------------		
New Zealand	1	352	353
United States	123	208	331
Brazil	50	70	120
Mexico	25	75	100
France	42	47	89
South Africa	2	57	59
Spain	32	21	53
Netherlands	7	45	52
Ireland	16	35	51
Australia	31	18	49
Canada	13	13	26
Chile	13	13	26
Norway	0	24	24
Korea	21	3	24
Austria	4	19	23
United Kingdom	11	9	20
Others	74	64	138
Total	465	1,073	1,538

[a] Source: Thompson (2004, p. 7). Totals may not sum due to rounding error.

Table 2. Parameter Estimates of Export Demand Equation

Parameter/ Statistic	Model A	Model B	Model C	Model D
λ	0.438	0.690	0.394	0.455
	(0.173)[a]	(0.132)	(0.166)	(0.093)
β_G	0.147	0.329	0.089	0.068
	(0.067)	(0.149)	(0.043)	(0.044)
$\tilde{\beta}_G$	-0.008	--	-0.007	-0.007
	(0.002)	--	(0.002)	(0.002)
β_P	-0.006	0.084	0.102	--
	(0.117)	(0.099)	(0.111)	--
β_{PS}	0.397	0.377	0.361	0.330
	(0.099)	(0.106)	(0.097)	(0.064)
β_Y	0.333	0.355	--	--
	(0.187)	(0.238)	--	--
β_{XR}	-0.451	-0.169	-0.676	-0.607
	(0.242)	(0.284)	(0.191)	(0.208)
β_T	-0.016	-0.012	-0.009	-0.010
	(0.004)	(0.004)	(0.002)	(0.001)
α	-3.021	-2.938	-0.631	-0.434
	(1.223)	(1.389)	(0.542)	(0.385)
Instrument Rank	13	12	13	13
J-Statistic	5.024	5.476	5.421	5.005
Sargan test (*p*-value)	0.285	0.242	0.367	0.543
Wald tests:				
Model A vs B:		19.152[b]		
		[0.0000]		
Model A vs C:			3.181[b]	
			[0.074]	
Model A vs D:				6.005[b]
				[0.0497]

[a] Asymptotic standard error. [b] Chi-square value; number below is the associated *p*-value.

Table 3. GMM Estimates of Export Promotion Expenditure Equations

Variable/ Statistic	Total Expenditure (*A*/*DEFL*)		Industry Expenditure (*AI*/*DEFL*)	
	Model A	Model B	Model A	Model B
Government Expend (*AG*/*DEFL*)	0.694	0.826	-0.976	-0.826
	(4.45)[a]	(41.07)	(-4.45)	--
Goodwill (*G/DEFL*)	0.553	0.505	0.527	0.472
	(8.14)	(18.39)	(5.67)	(21.07)
Flypaper (min(0,Δ*AG/DEFL*))	--	--	0.101	0.088
	--	--	(1.63)	(1.63)
Farm Price (P^{US}/*DEFL*)	0.017†	0.015†	-0.278†	-0.276†
	(0.55)	(0.74)	(-3.33)	(-3.73)
Exchange Rate (*SDR*)	0.291†	0.249†	0.137†	-0.0002†
	(4.57)	(5.42)	(0.60)	(-0.003)
Constant	0.022††	0.022††	0.056††	0.055††
	(7.55)	(9.71)	(11.49)	(12.77)
AR(1)	-0.248	-0.257	-0.297	-0.385
	(-5.92)	(-6.11)	(-5.51)	(-5.35)
R^2	0.88	0.89	0.31	0.38
DW-Statistic	1.92	1.93	1.96	1.93

[a] Asymptotic *t*-ratio. Note: model is estimated in first differences, so constant term is interpreted as trend effect.
†Coefficients converted to elasticities evaluated at mean data points.
†† Intercept converted from absolute to relative change.

Table 4. Test of Hypotheses

Hypothesis	Restriction	Chi-square	Probability	Result
Government dollars displace industry dollars	$a_1 + b_1 = 0$	0.574	0.448	Fail to reject
Industry expenditures increase when government expenditures decrease	$b_2 > 0$	2.65	0.103	Reject at $p < 0.10$

Appendix Table 1. Data Definitions and Sources

Variable Name	Definition	Source[a]
AG	U.S. government expenditures for export promotion	FAS
AI	U.S. industry expenditures for export promotion	FAS
A	Total U.S. expenditures for export promotion (= *AG* +*AI*)	
HV	Value of US high-value agricultural exports	ERS
BULK	Value of US bulk agricultural exports	ERS
X^{US}	Total value of US farm exports = *HV* + *BULK*	
X^{W}	Value of world imports of farm products	FAS
QBULK	Quantity of US bulk agricultural exports	FAS
P^{US}	Price of US bulk exports = *BULK*/*QBULK*	
DEFL	CPI for developed world	FAS
FCPI	World GNP deflator less the United States	ERS*
XR	World US agricultural trade weighted real exchange rate	ERS*
SDR	Special Drawing Rights (IMF exchange rate series)	FAS
PC_1	Real exchange rate for US competitors' HV farm products	FAS
PC_2	Exchange rate for US competitors' bulk farm products	FAS
MS_1	US high-value market share = HV/X^W	
MS_2	US bulk market share = $BULK/X^W$	
S_1	Normalized US high-value share = $MS_1/(MS_1+MS_2)$	
S_2	Normalized US bulk share = $MS_2/(MS_1+MS_2)$	

P^C	Stone index of substitute price = $PC_1^{(1-S_1)} \cdot PC_2^{(1-S_2)}$

[a] Abbreviations are defined as follows:
FAS = Foreign Agricultural Services, U.S. Department of Agriculture. The FAS data were obtained from personal correspondence.
ERS = Economic Research Service, U.S. Department of Agriculture. Specific sources are:
ERS = http://www.ers.usda.gov/Data/Fatus/#calendar
ERS* = http://www.ers.usda.gov/Data/Macroeconomics/

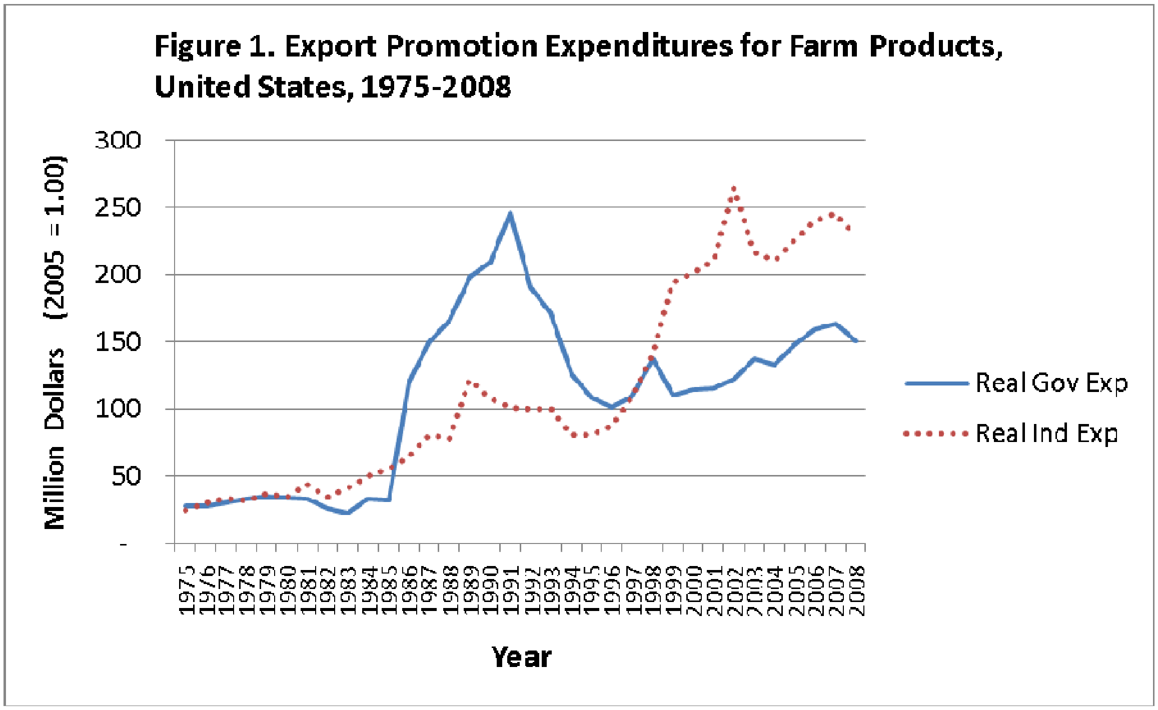

Figure 1. Export Promotion Expenditures for Farm Products,
United States, 1975-2008
Million Dollars (2005 = 1.00)
300
250
200
150
100
50
-
1975
1976
1977
1978
1979
1980
1981
1982
1983
1984
1985
1986
1987
1988
1989
1990
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
2001
2002
2003
2004
2005
2006
2007
2008
Year
Real Gov Exp
Real Ind Exp

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