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Multiple-Output Production Modeled with Three Functional Forms

Pedro A. Villezca-Becerra and C. Richard Shumway

Aggregate dual models are specified to examine multiple-output production relationships in each of four major, geographically dispersed, agricultural states (California, Iowa, Texas, and Florida). Three locally-flexible functional forms (translog, generalized Leontief, and normalized quadratic) are employed to conduct analytic simplification tests, estimate systems of output supply and input demand equations consistent with nonrejected hypotheses, derive elasticities, and determine to what extent analytic simplification tests and policy-relevant results are sensitive to functional form and state. Important differences in empirical implications were found due both to functional form and geographic unit, but differences were greater for the latter.

Key words: demand, functional form, homothetic separability, nonjointness, production, supply.

Duality theory has been used extensively in recent literature to analyze multiple-output production relationships in agriculture. Most previous studies, however, have limited empirical work to the use of a single functional form. Empirical evidence (Baffes and Vasavada; Berndt, Darrough, and Diewert; Chalfant) shows that parameter estimates and the generalization of policy-relevant results, such as elasticities, are often sensitive to choice of functional form.

Crucial simplifying assumptions also have been maintained in the specification of all models. Seldom have these simplifying assumptions been based on explicit hypothesis tests of the data. Further, while considerable empirical attention has been given to national and regional response (e.g., Huffman and Evenson; Akridge; Ball; Moschini; Kuroda; Vasavada and Chambers; Lopez; Antle), relatively few studies (e.g., Shumway 1983; Weaver 1983) have concentrated on developing policy-relevant estimates of output supply and input demand functions for individual states. Because production characteristics, market conditions, and other economically-relevant factors differ among and within states, government policies may have a different impact in each state. Regional or national estimates of impacts may have little relevance to individual states and to areas within a state (Houck and Ryan; Shumway and Alexander).

In this study, multiple-output model specification issues and production relationships will be analyzed for each of four major, geographically dispersed, agricultural states (California, Iowa, Texas, and Florida) using three locally-flexible functional forms (translog, generalized Leontief, and normalized quadratic). Specific objectives are to (*a*) conduct parametric tests of necessary and sufficient conditions for independent output modeling (nonjointness) and consistent aggregation (homothetic separability), (*b*) specify and estimate dual models consistent with analytic simplification hypotheses not rejected using any of the three functional forms, (*c*) derive short-run output supply and input demand elasticities, and (*d*) determine the extent to which test conclusions and elasticities are dependent on choice of functional form and state.

Model Specification

Assuming that each state's collection of producers behaves like a price-taking, profit-maximizing firm with a state-level aggregate production function, each state is modeled as though it were a perfectly

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Pedro A. Villezca-Becerra is a professor of agricultural economics, Universidad Autonoma Chapingo, Mexico, and formerly a research assistant at Texas A&M University. C. Richard Shumway is a professor of agricultural economics, Texas A&M University.

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competitive firm. Homothetic separability of the technology is necessary for consistent aggregation of both quantity and price indices (Pope and Hallam). With perfect competition in output and variable input markets, homothetic separability of the technology in a partition of quantity variables implies and is implied by homothetic separability in the corresponding partition of price variables in the dual restricted profit function. Therefore, subject to the maintained behavioral objective, homothetic separability can be tested just as conveniently with either a primal model or a dual model.

It is a simpler task, however, to test for nonjointness of the technology with a dual model. When the technology is nonjoint in inputs, decisions made on one output can be examined independently of decisions made on other outputs. Nonjointness implies that off-diagonal elements in the output submatrix of the profit function's Hessian matrix are zero. For a primal test, determinants of the primal Hessian submatrices must be singular when the technology is nonjoint. Therefore, short-run nonjointness and homothetic separability tests are conducted and multiple-output production relationships are examined using the restricted profit function,

$$\pi = x_0(P, Z) + P'X(P, Z) = \pi(P, Z), \tag{1}$$

where π is profit (receipts less variable costs) divided by the price of netput 0, $P = (p_1, \ldots, p_m)$ is the vector of output and variable input prices divided by the price of netput 0, $Z = (z_{m+1}, \ldots, z_n)$ is the vector of fixed input quantities and other nonprice exogenous variables, x_0 is the profit-maximizing quantity of netput 0, and $X = (x_1, \ldots, x_m)$ is the vector of profit-maximizing netput quantities (positively measured for outputs and negatively measured for inputs), which are functions of the exogenous variables P and Z.

Economic theory provides the basis for building production models such as (1), but finding the most appropriate functional form remains a pragmatic task. Since the researcher never knows the true functional form, a number of plausible functional forms have been proposed in recent years. They include the translog (Christensen, Jorgenson, and Lau), the generalized Leontief (Diewert), and the normalized quadratic (Lau 1978a). Each of these is a second-order Taylor series expansion, is linear in parameters, and is appropriately labeled a "locally-flexible" functional form (Fuss, McFadden, and Mundlak; Appelbaum). The use of locally-flexible functional forms has become increasingly popular since they impose fewer maintained hypotheses; that is, they are less restrictive than many other popular functional forms (such as the Cobb–Douglas and CES) and permit examination of comparative statics without imposing arbitrary cross-equation restrictions on choice at a point.

The translog specification is given by[1]

$$\begin{aligned}\ln(\pi) = b_0 &+ \sum_{i=1}^{m} b_i\ln(p_i) + \sum_{i=m+1}^{n} b_i\ln(z_i) \\ &+ .5\left[\sum_{i=1}^{m}\sum_{j=1}^{m} b_{ij}\ln(p_i)\ln(p_j) + \sum_{i=m+1}^{n}\sum_{j=m+1}^{n} b_{ij}\ln(z_i)\ln(z_j)\right] + \sum_{i=1}^{m}\sum_{j=m+1}^{n} b_{ij}\ln(p_i)\ln(z_j).\end{aligned} \tag{2}$$

The generalized Leontief is represented by [2]

$$\pi = c_0 + 2\sum_{i=1}^{m} c_i p_i^{.5} + 2\sum_{i=m+1}^{n} c_i z_i^{.5} + \sum_{i=1}^{m}\sum_{j=1}^{m} c_{ij}p_i^{.5}p_j^{.5} + \sum_{i=m+1}^{n}\sum_{j=m+1}^{n} c_{ij}z_i^{.5}z_j^{.5} + \sum_{i=1}^{m}\sum_{j=m+1}^{n} c_{ij}p_i z_j. \tag{3}$$

The normalized quadratic specification is

$$\pi = d_0 + \sum_{i=1}^{m} d_i p_i + \sum_{i=m+1}^{n} d_i z_i + .5\left(\sum_{i=1}^{m}\sum_{j=1}^{m} d_{ij}p_i p_j + \sum_{i=m+1}^{n}\sum_{j=m+1}^{n} d_{ij}z_i z_j\right) + \sum_{i=1}^{m}\sum_{j=m+1}^{n} d_{ij}p_i z_j. \tag{4}$$

For consistency with the competitive theory and a twice-continuously-differentiable technology, linear homogeneity of each profit function in prices is maintained through normalization (i.e., dividing profit and prices by the price of netput 0), convexity is maintained by the Cholesky factorization (Lau 1978b), and symmetry (reciprocity conditions) among the following first-derivative equations is imposed by linear parameter restrictions such that the jth price parameter in the ith equation is the same as the ith parameter in the jth equation. Monotonicity is not imposed but will be checked at each observation using the final model estimates.[3]

The estimation equations used to conduct the analysis are the linear first-derivative equations for each functional form. For the translog functional form, they are the netput share equations:

$$\partial\ln(\pi)/\partial\ln(p_i) = p_i x_i/\pi = s_i = b_i + \sum_{j=1}^{m} b_{ij}\ln(p_j) + \sum_{j=m+1}^{n} b_{ij}\ln(z_j), \qquad \text{for } i = 1, \ldots, m, \tag{5}$$

where s_i is the ith netput's share of profits. For the generalized Leontief and normalized quadratic, they are, respectively, the netput supply equations:

$$(6)\qquad \partial\pi/\partial p_i = x_i = c_i/p_i^5 + c_{ii} + \sum_{j=1}^{m} c_{ij}p_j^5/p_i^5 + \sum_{j=m+1}^{n} c_{ij}z_j, \qquad \text{for } i = 1, \ldots, m, j \neq i,$$

and

$$(7)\qquad \partial\pi/\partial p_i = x_i = d_i + \sum_{j=1}^{m} d_{ij}p_j + \sum_{j=m+1}^{n} d_{ij}z_j, \qquad \text{for } i = 1, \ldots, m.$$

Analytic Simplification Tests

With the output subvector represented by the index $i = 1, \ldots, \ell < m$, the technology exhibits short-run nonjointness in inputs for a subset of outputs $X^s = (x_e, \ldots, x_f)$, $e \geq 1, f \leq \ell$, if

$$(8)\qquad \partial^2\pi/\partial p_i p_j = 0, \qquad \forall\ i \in P^s, \qquad \forall\ j = 1, \ldots, \ell, \qquad i \neq j,$$

where P^s is the subset of normalized output prices corresponding to X^s. Short-run nonjoint production of the subset X^s can be violated due to technical interdependence in production among outputs and/or due to the presence of constraining allocatable input(s) impacting at least one output in X^s (Shumway, Pope, and Nash). While (8) does not assure that technical independence exists, it is a valid test to determine whether the researcher can simplify a short-run model of production by treating each output supply equation as being independent of other output prices (Chambers and Just). Global short-run nonjointness requires that the following linear restrictions be satisfied:

$$(9)\qquad c_{ij} = 0, \qquad \forall\ i \in P^s; \qquad j = 1, \ldots, \ell; \qquad i \neq j,$$

for the generalized Leontief functional form, and

$$(10)\qquad d_{ij} = 0, \qquad \forall\ i \in P^s; \qquad j = 1, \ldots, \ell; \qquad i \neq j,$$

for the normalized quadratic. Testing nonjointness using the translog functional form can only be done locally. At the point of approximation, i.e., $\ln(p_i) = 0$, $\forall\ i$, short-run nonjointness implies that the following nonlinear restrictions are satisfied:

$$(11)\qquad b_{ij} = -b_i b_j, \qquad \forall\ i \in P^s; \qquad j = 1, \ldots, \ell; \qquad i \neq j.$$

If the technology is homothetically separable in a partition X^s, the restricted profit function is homothetically separable in the corresponding normalized price partition P^s. The restricted profit function is homothetically separable if it is weakly separable (so all ratios of partial derivatives in P^s are independent of the magnitudes of all variables not in P^s), that is,

$$(12)\qquad \partial[(\partial\pi/\partial p_i)/(\partial\pi/\partial p_j)]/\partial p_k = 0, \qquad \forall\ i, j \in P^s, \forall\ k = \notin P^s,$$

and, if it is also homothetic in P^s (so the following modified Euler's theorem is satisfied), then

$$(13)\qquad \Sigma_{k\in s}\ \{\partial[(\partial\pi/\partial p_i)/(\partial\pi/\partial p_j)]/\partial p_k\}p_k = 0, \qquad \forall\ i, j \in P^s.$$

The restricted profit function is homothetically separable in P^s if any one of three sets of parametric restrictions is satisfied for a given functional form (Shumway 1989). Necessary and sufficient conditions for homothetic separability include both linear and nonlinear test restrictions. They are reported in table 1 for each functional form.

Data and Variable Specification

Exogenous variables included in the models were expected output prices, current variable input prices, quantities of fixed inputs (family labor and land), time, temperature, precipitation, and effective diversion payments. Annual state-level data for the period 1951–86 were used.[4] Except for pesticides, the output and input quantity and price data were compiled by Robert Evenson and his associates at Yale University for the period 1951–82, and updated to 1986 by Chris McIntosh at the University of Georgia. The pesticide price and quantity data were obtained from a data set provided by Chris McGath at the Economic Research Service. Data on government policy variables for maximum and minimum effective diversion payments for each farm program commodity were compiled by McIntosh (1989). The source of temperature and precipitation data weighted by cropland was Teigen and Singer.

Most of the price and quantity data were obtained from the U.S. Department of Agriculture's (USDA's)

Table 1. Necessary and Sufficient Restrictions for Homothetic Separability

Functional Form		Restrictions[a]	
Translog	Either:		
	(14a)	$b_i/b_j = b_{ik}/b_{jk}, \forall\ i, j \in P^s, \forall\ k,$	or
	(14b)	$\Sigma_{k\in s}\ b_{ik} = 0, \forall\ i \in P^s,$	and
		$b_{ik} = 0, \forall\ i \in P^s, \forall\ k \notin P^s,$	or
	(14c)	$b_i/b_j = b_{ik}/b_{jk}, \forall\ i, j, k \in P^s,$	and
		$b_{ik} = 0, \forall\ i \in P^s, \forall\ k \notin P^s.$	
Generalized Leontief	Either:		
	(15a)	$c_i/c_j = c_{ik}/c_{jk}, \forall\ i, j \in P^s, \forall\ k,$	or
	(15b)	$c_i = 2 \sum_{k=1}^{n} c_{ik}, \forall\ i \in P^s,$	and
		$c_{ik} = 0, \forall\ i \in P^s, \forall\ k \notin P^s,$	or
	(15c)	$c_{ik} = 0, \forall\ i \in P^s, \forall\ k.$	
Normalized Quadratic	Either:		
	(16a)	$d_i/d_j = d_{ik}/d_{jk}, \forall\ i, j \in P^s, \forall\ k,$	or
	(16b)	$d_i = \sum_{k=1}^{n} d_{ik}, \forall\ i \in P^s,$	and
		$d_{ik} = 0, \forall\ i \in P^s, \forall\ k \notin P^s,$	or
	(16c)	$d_{ik} = 0, \forall\ i \in P^s, \forall\ k.$	

[a] In our tests the index k did not include temperature, precipitation, time, or government diversion payments. Therefore, these variables must be included in both the first-stage and second-stage choice models specified for consistency with any of our nonrejected homothetic separability hypotheses.

Agricultural Statistics, Agricultural Prices, and *Field Crops Production, Disposition and Value.* Output prices were season average prices received by producers, and quantities were the harvest of the production year. Much of the fertilizer, feed, seed, hired labor, and miscellaneous inputs data were from the USDA's *State Farm Income and Balance Sheet Statistics.* Other sources of data included the USDA's *Farm Labor* and *Farm Real Estate Market Developments.*

Because of the large number of individual commercial outputs (as many as 25 in some states) and input categories (8), it was necessary initially to aggregate the data. Based on common nonrejected deterministic and stochastic nonparametric tests of separability using 1956–82 data for each of these states (Lim and Shumway), the data were aggregated into four output categories and three variable input categories.[5]

The output aggregates were crops, meat animals, milk–poultry, and other livestock. The meat animals category included cows and calves, hogs and pigs, and sheep and lambs. The milk–poultry category included milk, eggs, broilers, and turkeys. The other livestock category included all remaining commercial food animal commodities not included in the meat animal or milk–poultry aggregates.

The variable input categories were labor–capital, materials, and pesticides. The labor–capital category included hired labor, operation and repair of machinery and buildings, and service flows from machinery and service structures. The materials category included fertilizer, feed, seed, and miscellaneous inputs. The Tornqvist index was utilized in aggregating all price categories.

Guided by Lim's findings, one-year lagged output prices were used as the anticipated market prices. In a recent study, he chose this price expectation proxy over three alternative specifications (futures, univariate ARIMA, and composite) based on a comparison of measurement error sufficient for nonparametric consistency with the joint hypothesis of profit maximization, convex technology, and nonregressive technical change in two of our states, Iowa and Texas. Using a procedure adapted from Romain, expected prices of farm program commodities (corn, milk, cotton, sorghum, barley, wheat, oats, soybeans, rice, sugarbeets, peanuts, and tobacco) were specified as weighted averages of the anticipated market price and effective support price. The weights were dependent on the relative magnitudes of the anticipated market prices and effective support prices.[6] In a recent study including three of our states, McIntosh (1990) examined differences between three expectation mechanisms based on forecasting accuracy. He found that out-of-sample forecast errors were generally lowest for the weighted average specifications used here. The specification of effective diversion payments (included as a separate regressor) and effective support prices followed Houck and Ryan. The simple average of the maximum and minimum values of these variables compiled by McIntosh (1989) were used in the specification.[7]

Weather variables included temperature and precipitation either for the calendar year or for critical growing months. Exploratory analysis was conducted to determine which of ten weather variable specifications provided the greatest explanatory power. The weather variables chosen were annual average temperature and annual total precipitation in California, April–May average temperature and July–August total precipitation in Iowa, March–April average temperature and June–July total precipitation in Texas, and March–April average temperature and June–August total precipitation in Florida.

The variable input category, pesticides, and the two fixed input categories, family labor and land, were not aggregated further. A weighted average of effective diversion payments for farm program crops was constructed using profit shares in the respective states as weights.

Estimation and Tests

Systems of four output supply equations for crops, meat animals, milk–poultry, and other livestock, and two input demand equations for materials and pesticides were estimated for each state and functional form as specified in equations (5), (6), and (7) for the translog, generalized Leontief, and normalized quadratic, respectively.[8] The labor–capital input price was used to normalize profit and all output and variable input prices. Because of high collinearity arising from the large number of parameters to be estimated in their equations, the profit functions [equations (2)–(4)] and associated numeraire equations were excluded from the estimation procedure.

Error terms associated with each model were assumed to be additive, independently, and identically distributed with mean zero, and a constant contemporaneous covariance matrix. The covariance matrix used to transform the observation matrix was obtained by using the iterative version of Zellner's seemingly unrelated regression (ITSUR). Using the procedure SYSNLIN ITSUR in the SAS package, the variance-covariance matrix was iterated until it stabilized for each model. Imposition of the nonlinear inequality restrictions for maintaining convexity was accomplished by the Cholesky factorization. With the convexity restrictions imposed and using the observation matrix transformed by the iterated covariance matrix, a reduced gradient nonlinear program (Talpaz, Alexander, and Shumway) was employed using the algorithm code MINOS 5.1 (Murtagh and Saunders) to obtain least squares estimates that satisfied curvature properties for each system of output supply and input demand equations.

An exhaustive array of short-run nonjointness tests was conducted. Short-run nonjointness in inputs was tested for all four outputs, for each pair of outputs, and for individual outputs by sequentially imposing the restrictions for the respective subset outlined in equations (9)–(11) for the various functional forms.[9]

Guided by Hall's impossibility theorem of nonjointness and weak separability for a linear homogeneous production function, tests of the hypothesis of homothetic separability were conducted in partitions of outputs for which nonjointness was rejected by one or more tests. These separability tests were performed exhaustively by utilizing each of the three sufficient tests. For a given functional form, nonrejection of any of the three sufficient tests implied that the technology was homothetically separable in that partition for that state. To determine whether the conclusion was dependent on choice of functional form, the tests were conducted for each functional form.

An asymptotically valid chi-square test at the .01 level of significance was used for all individual tests.[10] Following Gallant and Jorgenson, this test was computed by estimating the unrestricted model with nonlinear ITSUR, retrieving the error covariance matrix, and estimating the restricted model with nonlinear SUR utilizing the retrieved error covariance matrix. Subtraction of the statistic labeled *OBJECTIVE*N* for the unrestricted model from the same statistic for the restricted model yielded the calculated chi-square statistic.

Empirical Results

Short-Run Nonjointness

The results of all short-run nonjointness tests conducted for each of the four states using each of the three functional forms are reported in table 2. Short-run nonjointness of all four output categories was not rejected in either Texas or California using any functional form. For Iowa, this hypothesis was not rejected using two functional forms, but it was rejected using the third (translog). For Florida, findings were the opposite of Iowa—the hypothesis was rejected using the generalized Leontief and the normalized quadratic, but was not rejected using the translog.

A similar pattern was found when conducting exhaustive tests of short-run nonjointness of pairs of outputs. Except for the meat animals and milk–poultry pair using the translog[11] and normalized quadratic in Texas, nonjointness was not rejected for any pair of output categories using any functional form in

Table 2. Short-Run Nonjointness Test Results for Texas, California, Iowa, and Florida, Using Three Locally-Flexible Functional Forms

	χ^2 Statistic[b]					
	Texas			California		
Output	TL	GL	NQ[a]	TL	GL	NQ
All Outputs	10.07	13.08	15.36	7.07	5.21	14.91
Crops, Meat Animals	7.85	12.74	14.66	7.17	4.64	14.03
Crops, Milk–Poultry	10.05	9.59	8.31	7.05	4.51	11.09
Crops, Other Livestock	10.11	6.14	11.39	2.19	5.20	12.18
Meat Animals, Milk–Poultry	—[c]	12.99	15.31*	7.61	5.19	14.62
Meat Animals, Other Livestock	9.97	11.70	11.97	2.30	4.09	14.91
Milk–Poultry, Other Livestock	6.18	11.56	14.17	1.21	5.17	10.50
Crops	4.12	1.29	4.00	1.54	2.03	4.52
Meat Animals	0.96	11.25	11.40*	1.88	3.13	13.65*
Milk–Poultry	0.47	8.69	7.97	0.63	4.49	6.23
Other Livestock	1.13	4.54	7.26	0.58	3.44	8.24

* Means hypothesis rejected at .01 level of significance.
[a] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[b] $\chi^2_{.01,6}$ for nonjointness test of all output categories; $\chi^2_{.01,5}$ for test of pair; $\chi^2_{.01,3}$ for test of individual output category.
[c] Convergence not obtained by nonlinear ITSUR procedure.

either Texas or California. For Iowa, short-run nonjointness was not rejected for any pair using two functional forms, but it was rejected with the translog for four of the six pairs. Only for crops and other livestock and for milk–poultry and other livestock was it not rejected in Iowa using any functional form. In Florida, short-run nonjointness in pairs of output categories was not rejected with any functional form only for crops and meat animals. For all other pairs, short-run nonjointness was rejected in this state by the generalized Leontief and/or the normalized quadratic. Only with the translog was the hypothesis not rejected for any pair of output categories in Florida.

Results from short-run nonjointness tests performed for individual outputs again showed generally the same tendency. With the exception of meat animals, short-run nonjointness could not be rejected for any individual output in Texas or California. For this output category, the hypothesis was rejected in both states using the normalized quadratic but not using either of the other two functional forms. In Iowa, short-run nonjointness was not rejected for any individual output using either the generalized Leontief or normalized quadratic functional form. When using the translog, nonjointness could not be rejected for only two individual output categories, crops and other livestock. In Florida, short-run nonjointness was not rejected for any individual output category using the translog. Only for crops and meat animals was the hypothesis not rejected using any functional form.

Since the objective of these tests was to identify opportunities for analytic simplification that are clearly justified for a given data set, nonjointness of all outputs was maintained in subsequent model design only where it was not rejected by any of the three functional forms at any level. With one exception, the same logic applies when considering whether to maintain nonjointness for a given pair of outputs. Because it is not possible to have joint production of only one output, rejection of the nonjointness hypothesis for a single output (when nonjointness was not rejected for any other output, any pair of outputs, or for all outputs) is not a meaningful rejection. For this reason, the reported rejection in table 1 of nonjoint production of meat animals in California using the normalized quadratic is insufficient to reject short-run nonjointness of all outputs in this state. Short-run nonjoint production was not rejected using any functional form for all outputs or for any pair of outputs in this state.

Short-run nonjoint production was not rejected in Texas for all outputs using any functional form. However, since it was rejected using the normalized quadratic both for the meat animals and milk–poultry subset and also for meat animals as an individual output, clear justification for maintaining short-run nonjointness in Texas applies only to crops and other livestock. Although specific test results differ among states, the same conclusion applies to Iowa. In Florida, only the crops and meat animals subset can be treated as nonjoint in the short run.

These results provide justification for modeling short-run supplies for each of the four output categories in California without considering changes in any other output category price. Texas and Iowa output supplies for crops and other livestock were modeled in the final specification without considering any output prices other than the own-category price. The same model simplification applied to crops and meat animal supplies in Florida.

Table 2. Continued

χ^2 Statistic						
Iowa			Florida			Critical Value
TL	GL	NQ	TL	GL	NQ	$\chi^2_{.01}$[b]
17.33*	8.61	10.62	11.37	18.88*	47.37*	16.81
16.33*	8.61	10.24	8.91	9.86	10.14	15.09
17.33*	8.50	10.06	9.47	18.09*	45.58*	15.09
8.15	4.10	5.01	11.37	10.71	24.16*	15.09
24.70*	8.48	10.24	10.31	17.95*	47.08*	15.09
24.37*	8.54	10.49	9.85	18.88*	46.73*	15.09
14.17	6.32	7.65	9.85	16.36*	44.05*	15.09
7.11	3.11	4.57	6.74	6.15	4.37	11.35
23.92*	8.41	9.73	2.57	7.19	7.97	11.35
14.11*	6.12	6.73	3.47	13.63*	42.07*	11.35
0.78	0.94	0.64	7.78	8.46	18.58*	11.35

These findings of mixed results with respect to short-run nonjoint production were similar to previously reported tests of this hypothesis. Short-run nonjointness of all outputs was previously rejected by Shumway (1983) for Texas; Moschini for Ontario, Canada; Ball for the U.S.; Chambers and Just for Israel; and Polson and Shumway for Arkansas, Mississippi, Oklahoma, and Texas. Short-run nonjointness of all outputs was not rejected by Shumway and Alexander for four of ten U.S. production regions (including regions that contained California and Iowa) or by Polson and Shumway for Louisiana. It was not rejected for selected subsets of outputs by Shumway (1983) for Texas or by Polson and Shumway for Mississippi, Oklahoma, and Texas. Thus, our nonjointness test results for California and Florida are the same as prior findings for similar areas but dissimilar to prior findings for areas containing Iowa and Texas. The dissimilarities may be due largely to different specifications of output categories. For example, when included, livestock generally was treated as a single category in the cited studies. However, crops frequently were disaggregated to some degree and sometimes to the individual crop level.

Homothetic Separability

Following the logic of Hall's impossibility theorem, homothetic separability was tested for all output partitions for which nonjointness was not clearly justified. These partitions included the meat animals and milk–poultry subset in Texas and Iowa and the milk–poultry and other livestock subset in Florida. Homothetic separability test results using all three functional forms are reported in table 3.

Homothetic separability test results provided no further justification for model simplification. With the exception of Florida, each sufficient test was rejected for each state with each functional form. The exception in Florida occurred because convergence was not obtained for the normalized quadratic and generalized Leontief for one of the three sufficient tests.

These findings rejecting homothetic separability of all output categories were similar to prior test results obtained by Weaver (1977) for North and South Dakota, Shumway (1983) for Texas, Ball for the U.S., and Polson and Shumway for five South-Central states. The failure to find support for homothetic separability in any output partition, however, is counter to the findings of Shumway (1983) and of Polson and Shumway in each state tested. As with the nonjointness tests, these differences may be due largely to differences in output category specifications.

Final Model

Based on the above nonrejected hypotheses, short-run output supply equations in California were specified in the final model specification as functions only of their own prices, prices of variable inputs, and quantities of the nonprice exogenous variables. The same applies to crop and other livestock categories in Texas and Iowa, and to crop and meat animal categories in Florida. Because no justification was found for a

Table 3. Homothetic Separability Test Results for Texas, Iowa, and Florida, Using Three Locally-Flexible Functional Forms[a]

			Test A[c]		Test B[d]		Test C[e]	
Separable Group	State	Functional Form[b]	χ^2 Statistic	Critical Value $\chi^2_{.01,8}$	χ^2 Statistic	Critical Value $\chi^2_{.01}$[f]	χ^2 Statistic	Critical Value $\chi^2_{.01}$[f]
Meat Animals and Milk–Poultry	Texas	TL	88.84	20.09	153.29	29.14	173.31	29.14
		GL	91.03	20.09	584.79	30.58	577.46	30.58
		NQ	25.05	20.09	661.58	30.58	73.60	30.58
	Iowa	TL	24.76	20.09	100.43	29.14	97.87	29.14
		GL	47.19	20.09	357.87	30.58	621.88	30.58
		NQ	59.24	20.09	294.05	30.58	95.61	30.58
Milk–Poultry and Other Livestock	Florida	TL	25.28	20.09	46.44	29.14	46.39	29.14
		GL	No Convergence		280.58	30.58	278.52	30.58
		NQ	No Convergence		254.35	30.58	64.50	30.58

[a] All hypotheses were rejected at the .01 level of significance.
[b] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[c] See equations (14a), (15a), and (16a).
[d] See equations (14b), (15b), and (16b).
[e] See equations (14c), (15c), and (16c).
[f] TL: $\chi^2_{.01,14}$; GL and NQ: $\chi^2_{.01,15}$.

higher level of data aggregation than maintained in the initial model design, the final models for each state consisted of six equations.

Summary statistics for the final model estimates are reported for each state and functional form in table 4. These estimates were obtained subject to nonrejected analytic simplifying assumptions and to linear homogeneity, symmetry, and convexity of the profit function in prices. Curvature properties were tested using the approximation test of Talpaz, Alexander, and Shumway and were not significantly violated (.05 level) in any of the states by the estimates using any functional form. Monotonicity was checked at each observation by examining the sign on every predicted output and input quantity. There were no violations of monotonicity by the translog estimates in any state or by the other functional forms in one state each. Monotonicity was significantly violated (.05 level) in Iowa by both the generalized Leontief and the normalized quadratic functional forms (six early observations were jointly significant in each case). Two nonsignificant violations were observed in California and three in Florida for the normalized quadratic, and one in Texas and two in Florida for the generalized Leontief. Therefore, except for estimates of two functional forms in Iowa, all empirical estimates were either consistent with, or not significant violations of, the joint hypothesis that the state behaves as though it were a price-taking, profit-maximizing firm.

The percent of significant parameters (.05 level) varied from 49% in Iowa to 13% in Florida; both extremes were for the translog functional form. Of the nonprice exogenous variables, time was significant in two-thirds of the estimated equations, land and family labor were significant in more than one-third, effective diversion payments and precipitation in about one-sixth, and temperature in only three of the 72 equations. No functional form consistently gave more significant parameter estimates than did any other form. All functional forms gave more significant parameter estimates in Texas than in Florida, but no other pairwise dominations were evident. Aggregate elasticities were computed from the parameter estimates of each model for the most recent observation (1986). The data used in estimation and the parameter estimates for each model are available on request from the authors.

Own-Price Elasticities

The own-price elasticities for each state and functional form are reported in table 5. Most were inelastic. The only exceptions were the generalized Leontief (GL) estimates for pesticides in California and Iowa, the GL and normalized quadratic (NQ) estimates for other livestock in Iowa, the translog (TL) estimates for milk–poultry and other livestock in Iowa, and the TL estimates for crops, meat animals, milk–poultry, and labor–capital in Florida. The only estimates that were consistently elastic across functional forms were for other livestock in Iowa, but none of these were significant at the .05 level.

A number of similarities were observed among these results across functional forms. For example,

Table 4. Summary Statistics for the Final Model Estimates for California, Iowa, Texas, and Florida, Using Three Locally-Flexible Functional Forms

	F-Statistic, Convexity[a]			Monotonicity: Number of Violations[b]			Monotonicity: χ^2 Statistic, Monotonicity			Percent of Significant Parameters (.05 level)		
State	TL	GL	NQ	TL	GL	NQ	TL	GL	NQ	TL	GL	NQ
California	0.53	0.48	0.70	0	0	2	NA	NA	1.53	36.5	28.6	26.9
Iowa	0.14	1.08	1.20	0	6	6	NA	39.22*	49.77*	49.2	22.2	31.7
Texas	1.03	0.21	0.29	0	1	0	NA	1.0E−7	NA	20.6	38.1	39.7
Florida	0.32	0.44	0.49	0	2	3	NA	1.90	7.24	12.7	30.2	30.2

Note: Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic. NA is not applicable.
* Violations significant at .05 level of significance.
[a] Critical value at .05 level of significance: $F_{63,153} = 1.32$.
[b] Potential number of violations is 216.

magnitude differences of .2 or less were observed for the elasticities of crops in California, crops in Iowa, and other livestock in Florida. No comparable similarities were observed across all functional forms in Texas.

Considering pairs of functional forms, the TL and GL gave similar own-price elasticities for meat animals, other livestock, materials, and labor–capital in California; for labor–capital in Iowa; for crops, other livestock, and pesticides in Texas; and for pesticides in Florida. The TL and NQ gave similar elasticities for materials and labor–capital in Iowa and for other livestock in Texas. The GL and NQ gave similar elasticities for meat animals, milk–poultry, and materials in California and Iowa; for meat animals, milk–poultry, materials, and labor–capital in Texas; and for crops, meat animals, milk–poultry, and labor–capital in Florida. Thus, the largest number of similar elasticities were for the GL and NQ, and the least were for the TL and NQ.

Seven of the 21 own-price elasticities estimated by the three functional forms were significant at the .05 level in California, four in Iowa, two in Texas, and five in Florida. Although relatively few of the estimated elasticities were significantly different from zero, more lay outside the 95% confidence interval of the same elasticities estimated by a different functional form.[12] For example, in California, neither the TL nor the GL estimates for meat animals, milk–poultry, materials, or pesticides were within the 95% confidence interval (CI) of the corresponding NQ estimate. The TL estimate for milk–poultry was also outside the GLCI, and the GL estimate for pesticides was outside the TLCI. The only NQ estimate that was outside the CI of one of the other functional forms was for milk–poultry.

In Iowa, the TL estimate was outside the NQCI for milk–poultry, materials, and pesticides, and outside the GLCI for meat animals and milk–poultry. The GL estimate was outside the TLCI for milk–poultry and outside the NQCI for pesticides. The NQ estimate was outside the TLCI for milk–poultry and pesticides and outside the GLCI for pesticides.

In Texas, the TL estimate was outside the GLCI for meat animals and outside the NQCI for meat animals, milk–poultry, materials, and pesticides. The GL estimate was outside the NQCI and the NQ was outside the TLCI for pesticides.

In Florida, the TL estimate was outside both the GLCI and NQCI for crops, meat animals, milk–poultry, and labor–capital, and outside the NQCI for materials and pesticides. The GL was outside the TLCI for meat animals and outside the NQCI for materials and pesticides. The NQ was outside the TLCI for meat animals.

More than half the elasticity estimates that lay outside the 95% CI of an alternative functional form were TL. More than half the CIs that did not include another functional form's estimate were NQ. The NQ estimates tended to have the greatest precision, and the TL estimates tended to be the furthest removed statistically from other functional form estimates.[13]

Cross-Price Elasticities

Cross-price elasticities for output–output and input–input pairs are reported in table 6. All other cross-price elasticities are in the appendix tables. Cross-price elasticities also were mostly inelastic. Since the estimation of the aggregated models for each state and functional form was performed maintaining nonjointness in some output categories, no output–output cross-price elasticities were derived in California

Table 5. Output Supply and Input Demand Own-Price Elasticities, 1986

	State and Functional Form					
	California			Iowa		
Output or Input	TL[a]	GL	NQ	TL	GL	NQ
Crops	.254	.315	.143	.107	.034	.018
	(.284)[b]	(.183)	(.099)	(.383)	(.129)	(.054)
Meat Animals	.417	.254	.126	.708	.246	.186
	(.182)	(.346)	(.041)	(.528)	(.155)	(.555)
Milk–Poultry	.401	.124	.061	1.127	.160	.175
	(.143)	(.109)	(.016)	(.269)	(.385)	(.119)
Other Livestock	.589	.756	.022	1.030	3.368	2.433
	(1.622)	(2.291)	(.878)	(8.447)	(15.520)	(4.756)
Materials	−.621	−.512	−.316	−.089	−.296	−.267
	(.447)	(.221)	(.083)	(.860)	(.183)	(.085)
Pesticides	−.458	−1.017	−.091	−.950	−1.177	−.040
	(.276)	(.527)	(.151)	(.275)	(.452)	(.078)
Labor–Capital	−.808	−.731	−.322	−.268	−.302	−.068
	(1.277)	(.606)	(.307)	(2.766)	(.564)	(.254)

[a] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[b] Approximate standard errors are in parentheses.

Table 6. Output-Output Supply and Input-Input Demand Cross-Price Elasticities, 1986

		State and Functional Form			
		California			Iowa
Quantity	Price	TL[a]	GL	NQ	TL
Outputs:					
Meat Animals	Milk–Poultry	—[b]	—	—	−.274
					(.093)[c]
Milk–Poultry	Meat Animals	—	—	—	−1.612
					(.544)
Milk–Poultry	Other Livestock	—	—	—	—
Other Livestock	Milk–Poultry	—	—	—	—
Inputs:					
Materials	Pesticides	.041	.073	.051	−.054
		(.044)	(.029)	(.026)	(.044)
Pesticides	Materials	.363	.647	.458	−.547
		(.396)	(.261)	(.228)	(.449)
Materials	Labor–Capital	.135	.020	.066	−.045
		(.690)	(.277)	(.129)	(1.272)
Labor–Capital	Materials	.154	.023	.075	−.065
		(.790)	(.317)	(.147)	(1.851)
Pesticides	Labor–Capital	.561	1.193	.017	.285
		(.764)	(.686)	(.419)	(.789)
Labor–Capital	Pesticides	.072	.153	.002	.041
		(.098)	(.088)	(.054)	(.114)

[a] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[b] Not computed because estimation was performed maintaining nonjointness for selected output categories in each state.
[c] Approximate standard errors are in parentheses.

Table 5. Continued

State and Functional Form					
Texas			Florida		
TL	GL	NQ	TL	GL	NQ
.222	.213	.001	1.355	.100	.102
(1.542)	(.352)	(.136)	(1.657)	(.104)	(.066)
.646	.031	.000	3.796	.081	.058
(1.721)	(.162)	(.035)	(.989)	(.248)	(.038)
.419	.095	.052	2.960	.015	.002
(.479)	(.180)	(.030)	(1.512)	(.196)	(.046)
.095	.249	.000	.015	.142	.110
(.917)	(1.506)	(.589)	(1.016)	(.762)	(.201)
−.506	−.260	−.122	−.874	−.643	−.266
(2.622)	(.290)	(.136)	(2.310)	(.189)	(.062)
−.634	−.564	−.210	−.556	−.608	−.165
(.198)	(.445)	(.069)	(.911)	(.377)	(.082)
−.622	−.212	−.155	−2.601	−.379	−.241
(7.581)	(.723)	(.324)	(6.166)	(.350)	(.188)

Table 6. Continued

State and Functional Form							
Iowa		Texas			Florida		
GL	NQ	TL	GL	NQ	TL	GL	NQ
−.041	−.035	.294	.014	−.001	—	—	—
(.023)	(.015)	(.502)	(.012)	(.006)			
−.239	−.208	.920	.045	−.002	—	—	—
(.137)	(.089)	(1.568)	(.039)	(.019)			
—	—	—	—	—	−.072	−.006	.006
					(.213)	(.079)	(.034)
—	—	—	—	—	−.367	−.033	.029
					(1.092)	(.406)	(.176)
−.007	.004	−.123	−.051	−.032	−.054	−.052	−.037
(.012)	(.008)	(.123)	(.016)	(.009)	(.247)	(.016)	(.009)
−.068	.038	−1.879	−.784	−.483	−.312	−.305	−.217
(.118)	(.078)	(1.877)	(.239)	(.137)	(1.433)	(.092)	(.053)
.063	.064	.101	.066	.113	−1.137	.449	.147
(.229)	(.115)	(3.886)	(.348)	(.177)	(4.392)	(.229)	(.110)
.092	.094	−.133	.087	.148	−.829	.327	.107
(.333)	(.167)	(5.110)	(.457)	(.233)	(3.202)	(.167)	(.080)
1.310	.037	−.411	1.102	.578	−.107	.351	.143
(.538)	(.206)	(2.804)	(.576)	(.221)	(3.145)	(.506)	(.189)
.189	.005	−.036	.095	.050	−.013	.044	.018
(.078)	(.030)	(.242)	(.050)	(.019)	(.394)	(.063)	(.024)

and only two were derived in each of the other states for each functional form. Regarding the signs of the output–output cross-price elasticities, the functional forms were all consistent only in identifying meat animals and milk–poultry as short-run gross substitutes in Iowa. Both the TL and NQ estimates of these elasticities were significantly (.05 level) different from zero in this state. However, the TL estimate lay outside the 95% CI of either the GL or the NQ estimates, and the GL and NQ estimates lay outside the TLCI. While a different sign was estimated by the NQ than by the TL and GL on output–output cross-price elasticities in other states, only the TL estimate in Texas lay outside the CI of any other estimate; it lay outside the CI of both the GL and NQ.

Results for the input–input elasticities showed more consistency across both functional forms and states. In California, the signs of these elasticities consistently implied that variable inputs were short-run gross substitutes. Nevertheless, only two estimates were significantly different from zero, GL and NQ estimates of the materials : pesticides relationship. The GL estimate of the pesticides : labor–capital relationships lay outside the NQCI. All other input–input cross-price elasticity estimates in California lay within the 95% CI for the corresponding estimate by alternative functional forms.

All functional forms implied that the pesticides input was a gross substitute to labor–capital in Iowa and a gross complement to materials in Texas and Florida. Except for the TL estimates, labor–capital was found to be a gross substitute in all states to both materials and pesticides. Except for California and the NQ estimate in Iowa, the materials input was found to be a gross complement to pesticides in all states. For the materials : pesticide relationships, the TL estimate lay outside both the GLCI and the NQCI in Iowa and Texas, and the GL estimate lay outside the NQCI in Texas. For materials : labor–capital, the TL estimate lay outside the GLCI and NQCI in Florida. For pesticides : labor–capital, the NQ estimate was outside the GLCI and the GL estimate was outside the NQCI in Iowa; the TL estimate was outside the GLCI and NQCI and the GL estimate was outside the NQCI in Texas. Although some elasticity signs differed, no other statistically substantial differences were attributable to differences in functional form.

The finding of largely inelastic output supply and input demand estimates obtained from all three functional forms was consistent with much prior literature at the state, regional, and national levels (e.g., Shumway 1983; Antle; Vasavada and Chambers; Shumway and Alexander; Huffman and Evenson). Few (e.g., Weaver 1983; Ball) have estimated elastic responses for a large portion of output supplies and input demands.

Conclusions

Dual models of agricultural production for Texas, California, Iowa, and Florida using the translog, generalized Leontief, and normalized quadratic locally-flexible functional forms were specified. Each state was modeled as a competitive industry with a twice-continuously-differentiable multiproduct transformation function and facing exogenous output and variable input prices. The initial model specification included four output and three variable input categories based on separability hypotheses not rejected by prior nonparametric tests. Exhaustive dual tests of short-run nonjointness (production independence) and homothetic separability (consistent aggregation and two-stage choice) of outputs were conducted to determine potential for analytic simplification and to determine whether conclusions were dependent on choice of functional form. When valid for a production system, either property reduces the number of parameters that must be estimated, thus conserving precious degrees of freedom and often reducing collinearity.

Short-run nonjointness was not rejected by any functional form for some or all of the four output categories in each state. However, homothetic separability was rejected by all functional forms in all tested partitions of outputs in each state. Given the empirical models designed for this study and the use of three equally-plausible functional forms, justification for legitimate analytic simplification was provided only in the form of imposing nullity restrictions on the matrix of independent parameters requiring estimation. Although additional possibilities for consistent aggregation were consistently rejected, degrees of freedom were conserved by maintaining short-run nonjointness in the final models. Monotonicity in prices of the final estimated profit functions was rejected only at early observations in Iowa for the generalized Leontief and normalized quadratic specifications. Convexity was not rejected for any state or functional form.

Most output supply and input demand relationships were inelastic. Many of the elasticity estimates were quite sensitive to choice of functional form. However, less than one-fourth of the own-price elasticity estimates lay outside the 95% confidence interval of any other functional form's estimate. Thus, while a considerable number of large differences due to functional form were noted, far fewer differences appeared to be important in a statistical sense.

Across states, the normalized quadratic own-price elasticity estimates were the most similar to each other for most output and input categories, and the translog estimates were generally the least similar.

Among cross-price relationships, all three functional forms revealed that meat animals and milk–poultry were short-run gross substitutes in Iowa, and all variable inputs were gross substitutes in California. In all states except California, labor–capital was generally found to be a gross substitute to materials and pesticides while materials and pesticides were generally found to be gross complements.

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Notes

[1] Because the effective diversion payment had a value of zero for some observations, a complete data series in logarithms could not be obtained. In addition, changes in the relative magnitudes of time have no intuitive meaning. Therefore, these two nonprice exogenous variables were included in original form (i.e., without transformation into logarithms or square roots) in all models.

[2] This form of the generalized Leontief is a slight modification from a second-order Taylor-series expansion in square roots. The variables in the last term are not transformed to square roots. This modification was incorporated in the final model to simplify the restrictions required to maintain convexity. In the model used to test for analytic simplification, all variables in all terms were expressed in square root form.

[3] Convexity and monotonicity were not maintained when conducting analytic simplification tests. Jorgenson and Lau (pp. 71–72) and Rothenberg (pp. 49–58) have shown that the asymptotic properties of these structural tests are the same with and without convexity being maintained. The inequality nature of the curvature restrictions does not affect the asymptotic distributions of the unconstrained test statistics. The allowable region for the estimates is reduced when imposing curvature, but the dimensions of the region are not. Therefore, although our sample size is modest, curvature was not maintained in the econometric estimation in order to reduce computational burden. Monotonicity was not imposed since prior empirical work (e.g., Shumway and Alexander; Lopez; Weaver 1983; Moschini) has shown this property rarely to be violated.

[4] Only data for 1951–82 were used in the analytic simplification tests.

[5] No violations of the generalized axiom of revealed preference, a necessary condition for a separable partition, were observed for materials in all four states, labor–capital in Texas, crops in three states, or milk–poultry in two states. In all but two of the cases where violations occurred, the true data could have satisfied both necessary and sufficient conditions for separability in these partitions if the observed quantity data had been measured with a 6% error. The remaining two cases required a 9% measurement error for consistency. Because measurement errors of these magnitudes are common in commodity production data, separability in these partitions was not rejected.

[6] When there was no announced support price for the commodity, its expected price was lagged market price. When the effective support price (Houck and Ryan) exceeded lagged market price, expected price was the effective support price. When lagged market price was at least as great as effective support price, both were included in the specification of expected price. When they were equal, the weight on lagged market price was .5; when it was double the effective support price, its weight was .67; when triple, its weight was .75. Thus, some weight was given to the effective support price whenever price supports were available for the commodity.

[7] Farm programs changed dramatically over the data period. For example, in some years the only price support mechanism was a nonrecourse loan. In other years, prices were supported also by target price, deficiency payment, acreage set-aside programs, payment in kind, and/or farmer-owned reserves. Our two policy variables were constructed for program commodities in affected years, attempting to use as consistent a set of procedures as possible. However, no additional policy variables or dummy variables were included as regressors in the model to account for major changes in the nature of farm programs.

[8] A multi-state model could have been estimated if the time series data could be pooled across states. Nonrejection of the hypothesis of identical technologies across the pooled states would justify either geographic aggregation or data pooling. Although not tested here, this hypothesis was rejected by Polson and Shumway for all pairs of states in two contiguous production regions.

[9] There are no independent nonjointness tests for a subset that consists of three output categories. Nonjointness of any subset of three outputs implies nonjointness of all four.

[10] Because the specification of each estimated model relied partially on the exploratory analysis of weather variables reported, significance levels are conditional on the estimated model specification being the "true" specification.

[11] No conclusion is available for this test with the translog since convergence was not obtained.

[12] Confidence intervals are asymptotic and conditional on the model specification.

[13] The fact that so few of the own-price elasticity estimates were statistically significant obviously limits their practical usefulness for policy analysis purposes. These estimates by each of the functional forms generally were not very precise.

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Appendix Table 1. Output-Input Supply Cross-Price Elasticities, 1986

		State and Functional Form											
		California			Iowa			Texas			Florida		
Quantity	Price	TL[a]	GL	NQ	TL	GL	NQ	TL	GL	NQ	TL	GL	NQ
Crops	Materials	−.062	−.108	−.031	−.020	−.045	−.020	−.257	−.148	.000	−.429	−.099	−.048
		(.229)[b]	(.075)	(.044)	(.447)	(.062)	(.034)	(2.096)	(.192)	(.111)	(1.093)	(.039)	(.028)
	Pesticides	.053	.030	.009	−.045	−.002	−.002	−.086	−.000	.003	−.028	−.015	−.004
		(.026)	(.015)	(.015)	(.025)	(.005)	(.004)	(.100)	(.013)	(.007)	(.122)	(.010)	(.006)
	Labor–Capital	−.246	−.237	−.121	−.042	−.012	.004	.121	−.064	−.005	−.899	.014	−.050
		(.382)	(.201)	(.109)	(.673)	(.149)	(.064)	(3.118)	(.411)	(.176)	(2.106)	(.115)	(.071)
Meat Animals	Materials	−.614	−.384	−.326	−.167	−.165	−.163	−.364	−.052	.002	−2.070	.082	−.025
		(.490)	(.147)	(.086)	(.651)	(.098)	(.056)	(2.153)	(.080)	(.052)	(2.643)	(.078)	(.048)
	Pesticides	−.095	.071	.048	−.072	.035	.017	−.080	−.010	.000	−.622	−.094	−.053
		(.087)	(.057)	(.044)	(.036)	(.013)	(.009)	(.100)	(.007)	(.004)	(.325)	(.050)	(.026)
	Labor–Capital	.292	.060	.151	−.196	−.075	−.005	−.497	.016	−.001	−1.104	−.069	.021
		(.805)	(.441)	(.105)	(.980)	(.195)	(.081)	(3.190)	(.200)	(.063)	(5.080)	(.322)	(.067)
Milk–Poultry	Materials	−.335	−.206	−.096	.020	−.100	−.085	−.517	−.215	−.139	−.137	.093	.010
		(.270)	(.051)	(.024)	(.657)	(.126)	(.078)	(1.974)	(.082)	(.038)	(2.198)	(.052)	(.027)
	Pesticides	−.038	.023	.021	.043	−.152	−.050	−.114	−.061	−.034	.199	.017	−.004
		(.045)	(.027)	(.014)	(.136)	(.106)	(.061)	(.102)	(.028)	(.014)	(.381)	(.063)	(.027)
	Labor–Capital	−.028	.059	.015	.422	.331	.168	−.707	.136	.122	−2.951	−.118	−.013
		(.438)	(.140)	(.032)	(1.025)	(.457)	(.179)	(2.930)	(.219)	(.054)	(4.367)	(.257)	(.069)
Other Livestock	Materials	2.847	1.994	.156	.962	2.395	2.178	.392	.158	−.001	−.056	−.110	−.026
		(1.482)	(1.010)	(.559)	(5.680)	(4.143)	(2.661)	(2.433)	(.646)	(.397)	(.652)	(.179)	(.108)
	Pesticides	−.903	−.137	.019	4.832	.352	−1.007	−.087	.485	−.001	.030	−.293	−.104
		(1.110)	(.713)	(.424)	(6.715)	(5.240)	(2.977)	(.444)	(.340)	(.223)	(.817)	(.247)	(.116)
	Labor–Capital	−2.532	−2.613	−.197	−6.824	−6.115	−3.603	−.400	−.892	.002	.378	.294	−.009
		(3.516)	(3.006)	(1.124)	(16.183)	(19.024)	(6.210)	(3.817)	(1.918)	(.745)	(2.215)	(1.001)	(.311)

[a] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[b] Approximate standard errors are in parentheses.

Appendix Table 2. Input-Output Demand Cross-Price Elasticities, 1986

		State and Functional Form											
		California			Iowa			Texas			Florida		
Quantity	Price	TL[a]	GL	NQ	TL	GL	NQ	TL	GL	NQ	TL	GL	NQ
Materials	Crops	.131	.227	.066	.031	.068	.031	.230	.133	−.000	1.348	.310	.150
		(.480)[b]	(.157)	(.091)	(.679)	(.095)	(.052)	(1.874)	(.172)	(.099)	(3.438)	(.121)	(.087)
	Meat Animals	.134	.084	.071	.163	.161	.159	.351	.050	−.001	.637	−.025	.008
		(.107)	(.032)	(.019)	(.635)	(.095)	(.055)	(2.079)	(.077)	(.051)	(.813)	(.024)	(.015)
	Milk–Poultry	.223	.137	.064	−.003	.017	.014	.160	.066	.043	.074	−.050	−.005
		(.179)	(.034)	(.016)	(.109)	(.021)	(.013)	(.611)	(.025)	(.012)	(1.189)	(.028)	(.015)
	Other Livestock	−.042	−.030	−.002	−.002	−.006	−.005	−.010	−.004	.000	.006	.012	.003
		(.022)	(.015)	(.008)	(.013)	(.010)	(.006)	(.064)	(.017)	(.010)	(.069)	(.019)	(.011)
Pesticides	Crops	−.999	−.569	−.163	.694	.031	.026	1.172	.005	−.040	.506	.267	.067
		(.480)	(.279)	(.289)	(.375)	(.083)	(.058)	(1.363)	(.172)	(.095)	(2.233)	(.188)	(.108)
	Meat Animals	.184	−.137	−.094	.704	−.341	−.169	1.180	.145	−.006	1.113	.168	.096
		(.169)	(.111)	(.086)	(.357)	(.132)	(.090)	(1.475)	(.101)	(.065)	(.582)	(.089)	(.047)
	Milk–Poultry	.228	−.136	−.124	−.072	.254	.083	.538	.289	.160	−.626	−.053	.012
		(.266)	(.161)	(.083)	(.227)	(.178)	(.101)	(.481)	(.133)	(.066)	(1.197)	(.200)	(.085)
	Other Livestock	.120	.018	−.003	−.114	−.008	.024	.035	−.193	.000	−.018	.180	.064
		(.147)	(.095)	(.056)	(.159)	(.124)	(.071)	(.177)	(.136)	(.089)	(.500)	(.151)	(.071)
Labor–Capital	Crops	.590	.570	.290	.092	−.027	−.010	−.143	.075	.005	2.061	−.031	.115
		(.918)	(.483)	(.261)	(1.487)	(.329)	(.142)	(3.665)	(.484)	(.207)	(4.829)	(.263)	(.164)
	Meat Animals	−.073	−.150	−.038	.278	107	.007	.631	−.021	.002	.248	.016	−.005
		(.201)	(.110)	(.026)	(1.391)	(.277)	(.115)	(4.052)	(.253)	(.081)	(1.140)	(.072)	(.015)
	Milk–Poultry	.021	−.045	−.011	−.102	−.080	−.041	.288	−.055	−.050	1.164	.047	.005
		(.334)	(.107)	(.024)	(.247)	(.110)	(.043)	(1.192)	(.089)	(.022)	(1.722)	(.101)	(.027)
	Other Livestock	.043	.045	.003	.023	.021	.012	.014	.031	−.000	−.029	−.023	.001
		(.060)	(.051)	(.019)	(.055)	(.065)	(.021)	(.131)	(.066)	(.026)	(.170)	(.077)	(.024)

[a] Functional form codes: TL is translog, GL is generalized Leontief, NQ is normalized quadratic.
[b] Approximate standard errors are in parentheses.