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Total Factor Productivity and Processed Food Trade: A Cross-Country Analysis

Munisamy Gopinath and Jason Carver

Processed food products account for a growing share of global agricultural trade. Growth in total factor productivity and intersectoral linkages between the agricultural and processed food sectors are hypothesized as factors explaining this phenomenon. Estimating the neoclassical trade model using an internationally comparable database, we find evidence of (*a*) Hecksher-Ohlin (factor endowments) and Ricardian-type (technology) effects in agricultural and processed food trade, and (*b*) transfer of comparative advantage from the primary agricultural sector to the processed food sector. Thus, public policies protecting primary agriculture can adversely affect processed food sectors, while those supporting R&D efforts can bring about dynamic comparative advantage.

Key words: dynamic comparative advantage, processed food trade, productivity growth

Introduction

Processed food products account for a growing share of agricultural trade. Currently, two-thirds of globally traded agricultural products, valued at over $400 billion, undergo some form of value addition prior to shipment (Coyle et al.). Most countries perceive this trend as beneficial, because processed products are generally of high value, and the addition of value to primary agricultural products generates employment. Not surprisingly, countries tending to export a higher share of processed foods [e.g., those within the European Union (EU)] continue to emphasize the need to maintain and increase that trend. Countries tending to export more bulk commodities (e.g., the United States) have recognized the need to adjust to the changing structure of agricultural trade.[1]

The purpose of this study is to examine the effects of technology and factor supplies (labor, capital) on specialization within agriculture. In particular, we seek to identify the sources of comparative advantage in the processed food and primary agricultural sectors of the United States and major developed countries.[2]

Munisamy Gopinath is an associate professor in the Department of Agricultural and Resource Economics, Oregon State University, Corvallis; Jason Carver is an agricultural economist with the USDA, Foreign Agricultural Service, Forestry and Fishery Products Division, Washington, DC.

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[1] Although it is a global phenomenon, the increasing share of processed food in agricultural trade is more pronounced in developed countries. Coupled with the liberalization policies underway (e.g., the EU 2000 Agenda), intra-EU trade is important, and so it is included in this study.

[2] Based on the International Standard Industrial Classification (ISIC) coding system, agricultural products are those classified under ISIC 1, while the products under ISIC 31 (excluding tobacco) are processed food products.

Several explanations have been offered for the increasing share of processed foods in global agricultural trade (Coyle et al.). The exhaustive list includes demand, supply, and policy factors. The level and composition of food demand vary in a systematic way with per capita income levels and associated changes in income elasticity levels, and thus affect the pattern of trade.

On the supply side, the Rybczynski theorem of trade theory predicts that, when the endowment of one factor increases faster than others, the sectors using this factor most intensively increase their output faster compared to other sectors. For instance, if inputs used in food processing accumulate faster than those used in agriculture, one may observe an increase in the share of processed food products in agricultural trade. Finally, reduction in transport and communication costs and differential levels of protection can affect output and the structure of trade. For example, if protection levels for livestock products are lowered at a faster rate than those of grain products, then beef imports may rise relative to wheat imports.

Previous investigations of the pattern of agricultural trade appear to have overlooked a key explanatory variable, namely technological change in agricultural and processed food sectors and the intersectoral linkages.[3] The linkages between primary agriculture (which produces bulk products) and the processed food sector (which adds value to the bulk products) through prices are obvious. However, the effect of productivity growth in agriculture on the processed food sector and vice versa has received little or no attention. Specifically, productivity growth in primary agriculture in a global context places downward pressures on real agricultural price. These price declines in primary agriculture translate into declines in the procurement (input) costs of the food processing sector, because the cost share of bulk commodities in food processing is substantially large (U.S. Department of Commerce, Bureau of Economic Analysis). Such a cost advantage is likely to allow the expansion of the processed food sector. In turn, the growth of the processing sector may translate into an increased demand for bulk commodities, and the price decline that would otherwise occur through a lack of growth in demand will likely be mitigated.

While most researchers agree that demand, supply, and policy factors explain trade patterns, their relative contributions have been a subject of debate. Adding productivity growth to the list of factors affecting trade patterns necessitates confirmation as well as assessment of its relative contribution. The objective of this study is to identify and analyze the supply-side effects—i.e., productivity growth, intersectoral linkages, and factor endowments—on export patterns and performance of agricultural and processed food sectors. In doing so, some of the policy effects are taken into account [such as the exchange rate and common agricultural policy (CAP) of the EU countries], while the richness of a panel database and procedure is used to capture demand-side and other policy effects.

Results indicate that productivity, intersectoral linkages, and capital accumulation have important effects on agricultural and processed food trade patterns. Specifically, accumulation of knowledge and the resulting efficiency gains in the primary agricultural sector are shared with the downstream processed food sectors. These factors lend support to the notion that comparative advantage in the context of open economies is not

[3] This analysis focuses on the supply side because demand (per capita income growth) and policy factors (liberalization) of other countries are not necessarily influenced by exporting countries.

necessarily static, especially when knowledge, and physical and human capital can be accumulated (Krugman; Romer; Grossman and Helpman). For instance, countries exporting primary agricultural products need not do so over a prolonged period. Comparative advantage can be changed by conscious efforts directed toward the accumulation of various factors, including knowledge.

Theoretical and Empirical Framework

The approach used to test for and identify effects of technological change on the pattern of agricultural trade draws on the conceptual contributions of earlier studies (Dixit and Norman; Woodland), as well as previous empirical applications (Harrigan; Trefler 1993, 1995; Kohli). Consider a small, open economy with fixed factor supplies, constant returns-to-scale technology, and competitive markets. The maximization of final output value characterizes the general equilibrium of this economy. Formally, the maximized value of gross domestic product (GDP) is given by:

$$(1) \qquad G(\mathbf{p}, \mathbf{v}) = \text{Max } \mathbf{p} \cdot \mathbf{y}, \quad \text{s.t.: } \mathbf{y} \in Y(\mathbf{v});\ \mathbf{p}, \mathbf{y} \in \mathbb{R}^n;\ \mathbf{v} \in \mathbb{R}^m,$$

where $G(\mathbf{p}, \mathbf{v})$ is the revenue or GDP function, $\mathbf{y}$ is the final goods vector, $\mathbf{p}$ is the row vector of final goods prices, and $Y(\mathbf{v})$ is the convex production set for factor endowments $\mathbf{v}$. The notation $\mathbf{p} \cdot \mathbf{y}$ denotes the scalar product of the two vectors. If $m \geq n$, and $G(\mathbf{p}, \mathbf{v})$ is twice differentiable, the gradient of $G(\mathbf{p}, \mathbf{v})$ with respect to $\mathbf{p}$ yields the vector of net output supplies, $\mathbf{y}(\mathbf{p}, \mathbf{v})$.

In order to explain the pattern of agricultural trade, this basic model must be extended in two significant ways. First, the GDP function framework is extended to model exports and imports (Burgess; Kohli). In the definition of production possibilities, export goods are considered as outputs, while imports are designated as inputs.[4]

Consistent with Kohli, consider the partition of the final goods vector $\mathbf{y}$ as domestic output, exports, and imports $(\mathbf{y}_D, \mathbf{y}_X, \mathbf{y}_M)$, with the corresponding price vector $\mathbf{p} \equiv (\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M)$. By modeling exports and imports, the above GDP function can be extended to a gross national product (GNP) function as follows (see also Diewert and Morrison):

$$(2) \qquad G(\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M, \mathbf{v}) = \text{Max } \mathbf{p}_D \cdot \mathbf{y}_D + \mathbf{p}_X \cdot \mathbf{y}_X + \mathbf{p}_M \cdot \mathbf{y}_M, \quad \text{s.t.: } \mathbf{y} \in Y(\mathbf{v});\ \mathbf{p}, \mathbf{y} \in \mathbb{R}^n;\ \mathbf{v} \in \mathbb{R}^m,$$

where $\mathbf{y}_M$ (the quantity of imports) is negative. Note that $\mathbf{p}_D \cdot \mathbf{y}_D$, $\mathbf{p}_X \cdot \mathbf{y}_X$, and $\mathbf{p}_M \cdot \mathbf{y}_M$ are scalar products. Similarly, if $m \geq n$, and $G(\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M, \mathbf{v})$ is twice differentiable, the gradients of $G(\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M, \mathbf{v})$ with respect to $\mathbf{p}_X$ and $\mathbf{p}_M$ yield the vectors of export supplies $\mathbf{y}_X(\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M, \mathbf{v})$ and imports $\mathbf{y}_M(\mathbf{p}_D, \mathbf{p}_X, \mathbf{p}_M, \mathbf{v})$, respectively. Following Diewert, one can show that the aggregate GNP (revenue) function can be a sum of sectoral GNP (revenue) functions—i.e., individual revenue functions, such as agricultural, food processing, manufacturing, and others, exist such that their sum equals economywide GNP.

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[4] Both Kohli and Burgess recognize the fact that most internationally traded goods are intermediates, and even finished goods go through commercial channels (services sectors) before serving final demand.

The (sectoral) GNP function can be extended to incorporate technological change, as illustrated by Dixit and Norman, and applied by Kohli and by Harrigan. The level of technology is denoted by the parameter θ and is considered to represent Hicks-neutral technical change at the sectoral level, which translates into product-augmenting technical change at the aggregate level. Empirically, θ suggests productivity levels and rates can vary among sectors (Dixit and Norman; Harrigan).

For specification of the empirical model, the translog functional form is chosen. Moreover, each of the price and goods vectors ($\mathbf{p_D}$, $\mathbf{p_X}$, $\mathbf{p_M}$) and ($\mathbf{y_D}$, $\mathbf{y_X}$, $\mathbf{y_M}$) is reduced to a scalar. Alternatively, the various domestic (export or import) goods are aggregated into a single domestic (export or import) good. For each sector—agriculture and processed food—a translog GNP function [$G(\mathbf{p}, \mathbf{v}, \theta)$], including exports and imports, is specified as:

$$(3) \quad \begin{aligned} \ln G(\mathbf{p}_t, \mathbf{v}_t, \theta_t) = {} & \alpha_0 + \sum_{j=D,X,M} \alpha_j \ln(p_{jt}) + \frac{1}{2} \sum_{j=D,X,M} \sum_{n=D,X,M} \alpha_{jn} \ln(p_{jt}) \ln(p_{nt}) \\ & + \sum_{i=R,L,K} \beta_i \ln(v_{it}) + \frac{1}{2} \sum_{i=R,L,K} \sum_{m=R,L,K} \beta_{im} \ln(v_{it}) \ln(v_{mt}) \\ & + \sum_{j=D,X,M} \sum_{i=R,L,K} \gamma_{ji} \ln(p_{jt}) \ln(v_{it}) + \delta_1 \ln(\theta_t) + \frac{1}{2} \delta_2 \ln(\theta_t)^2 \\ & + \sum_{j=D,X,M} \zeta_j \ln(p_{jt}) \ln(\theta_t) + \sum_{i=R,L,M} \eta_i \ln(v_{it}) \ln(\theta_t), \end{aligned}$$

where indexes $j, n = D, X, M$ denote, respectively, outputs, exports, and imports; indexes $i, m = R, L, K$ denote land, labor, and capital stock (factor endowments), respectively; and index t denotes time. All Greek characters (α, β, γ, δ, ζ, and η) are parameters.

For homogeneity conditions, the following restrictions should be imposed on the parameters of equation (3):

$$\sum_{j=D,X,M} \alpha_j = 1; \quad \sum_{j=D,X,M} \alpha_{jn} = 0; \quad \sum_{n=D,X,M} \alpha_{jn} = 0; \quad \sum_{i=R,L,K} \beta_i = 1; \quad \sum_{i=R,L,K} \beta_{im} = 0;$$

$$\sum_{m=R,L,K} \beta_{im} = 0; \quad \sum_{j=D,X,M} \gamma_{ji} = 0; \quad \sum_{i=R,L,K} \gamma_{ji} = 0; \quad \sum_{j=D,X,M} \zeta_j = 0; \quad \sum_{i=R,L,M} \eta_i = 0.$$

Further, to ensure symmetry conditions, the parameters of (3) should be subjected to the following restrictions:

$$\alpha_{jn} = \alpha_{nj}; \quad \beta_{im} = \beta_{mi}; \quad \gamma_{ji} = \gamma_{ij}$$
$$\text{for } j, n = D, X, M; \quad i, m = R, L, K.$$

Differentiating $\ln G(\mathbf{p}_t, \mathbf{v}_t, \theta_t)$ with respect to the export price $\ln(p_{Xt})$ gives the export share in GNP:[5]

$$(4) \quad \frac{p_{Xt} y_{Xt}}{p_{Dt} y_{Dt} + p_{Xt} y_{Xt} + p_{Mt} y_{Mt}} \equiv s_{Xt} = \alpha_X + \sum_{n=D,X,M} \alpha_{nX} \ln(p_{nt}) + \sum_{i=R,L,K} \gamma_{iX} \ln(v_{it}) + \zeta_X \ln(\theta_t).$$

[5] Import share equations can be derived in a similar fashion, but the focus here is on exports.

To focus on the effect of factor endowments and technologies, which differ among countries, the price expression in equation (4) can be reduced to a "country-dependent" parameter (Harrigan).[6] Choosing a country as reference, and using subscript c to denote countries, the estimable share equation becomes:

$$(5) \qquad s_{Xct} = \alpha_X + \lambda_{Xc} + \mu_X s_{Xct-1} + \sum_{i=L,K} \gamma_{iX} \ln\left(\frac{v_{ict}}{v_{Rct}}\right) + \zeta_X \ln(\theta_{ct}) + \psi_X \ln(E_t) + \varepsilon_{ct},$$

where

$$\lambda_{Xc} = \sum_{n=D,X} \alpha_{nX} \ln(p_{nct}/p_{Mct})$$

is a country-dependent parameter which could be treated either as fixed or random, E_t is the real exchange rate, and ε_{ct} is an error term. The introduction of a lagged dependent variable acknowledges that reallocation of factors among sectors occurs with a lag in response to changes in technology and factor endowments (Harrigan). If λ_{Xc} is not equal to zero, the short- and long-run effects of factor endowments and productivity on export shares are different. The normalization of factor endowments [the summation term in equation (5)] imposes the condition that the agricultural export share equation is homogeneous of degree zero in the factor endowments of land, labor, and capital:

$$\sum_{i=R,L,K} \gamma_{ji} = 0.$$

Symmetry restrictions are not imposed because data constraints prevent the estimation of other share equations.[7]

Denomination of domestic, export, and import revenues using a common currency (US$) to derive export shares necessitates the inclusion of the real exchange rate (E_t) in the share equation. More specifically, prices in local currencies are multiplied by the exchange rate to the common currency (US$) to obtain an internationally comparable set of revenue and cost accounts. To the extent the real exchange rate and its change differ among countries, introducing it in the export share equation can capture at least a portion of the differences in prices. Note that equation (5) can be estimated over a panel of countries identified in the following section.

Equation (5) allows for differences in technology to explain the variations in export shares of GNP among countries. While a positive ζ_X suggests technological improvements over time have favored exports, it also implies that countries with a relatively higher level of technology are likely to experience bigger gains in the export share of GNP (Kohli). Similar to the accumulation of factor endowments, growth in productivity, which depends crucially on investments in R&D, can bring about changes in comparative advantage.[8]

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[6] To the extent prices differ between countries, the price expression reflects protection or promotion policies. Hence, this model can be extended to include these policies. Neither the price data in common currency nor producer subsidy equivalents and export subsidies are available for the 1975–95 time period.

[7] The inclusion of other share equations (e.g., imports, input shares) will lead to a seemingly unrelated regression (SUR) system of equations in a panel setting. Any efficiency gain is outweighed by data constraints and the possibility that there are a large number of asymptotically efficient estimators. (For such panel SUR estimators, see Baltagi, chapter 6.)

[8] The dependent variable in the empirical model is value of exports normalized by sectoral GNP. In a world characterized by positive growth rates, a rise in the growth rate of one country's (normalized) exports relative to that of another country implies the former is gaining bilateral competitiveness. A similar comparison can be drawn for multilateral settings.

The case of the food processing industry and its export share of production differs from the above specification because food processing uses raw and bulk commodities as intermediate inputs. Woodland has demonstrated that the GNP function, $G(\mathbf{p}, \mathbf{v}, \theta)$, can be easily extended to account for intermediate inputs. Modeling intermediate inputs in this framework does not alter the structure and properties of the GNP function, except for expanding the price vector ($\mathbf{p}$) to include prices of intermediates. Specifically, the export share equation for the food processing sector (with a superscript F) is derived from a translog GNP function (3) as follows:

$$(6) \qquad \frac{p_{Xt}^F y_{Xt}^F}{p_{Dt}^F y_{Dt}^F + p_{Xt}^F y_{Xt}^F + p_{Mt}^F y_{Mt}^F + p_{It}^F y_{It}^F} = s_{Xt}^F$$
$$= \alpha_X^F + \sum_{n=D,X,M,I} \alpha_{nX}^F \ln(p_{nt}^F) + \sum_{i=L,K} \gamma_{iX}^F \ln(v_{it}^F) + \zeta_X^F \ln(\theta_t^F),$$

where p_I denotes the price of domestic agricultural intermediates employed in the food processing sector, and y_M and y_I are negative by definition.[9] The price of primary agricultural commodities (p_I) is specified as $p_I(\theta^A)$, reflecting the fact that agricultural technological change (θ^A) leading to supply shifts causes agricultural prices to fall, and thus benefits food processing.[10]

Once again, choosing a country as reference, and using subscript c to denote countries, the estimable export share equation can be rewritten as:

$$(7) \qquad s_{Xct}^F = \alpha_X^F + \lambda_{Xc}^F + \mu_X^F s_{Xct-1}^F + \gamma_{KX}^F \ln\left(\frac{v_{Kct}^F}{v_{Lct}^F}\right) + \zeta_X^F \ln(\theta_{ct}^F)$$
$$+ \psi_X^F \ln(E_t) + \alpha_{IX} \ln(\theta_{ct}^A) + \varepsilon_{ct}^F.$$

Similar to equation (5), the price expression has been reduced to a "country-dependent" parameter which could be treated either as fixed or random. The lagged dependent variable (s_{Xct-1}) is included to distinguish between short- and long-run effects of factor endowments, productivity, and exchange rates on export shares. As before, the normalization of factor endowments imposes one of the homogeneity conditions,

$$\sum_{i=L,K} \gamma_{ji} = 0,$$

and conversion of revenues to a common currency to derive export shares adds the real exchange rate (E_t) to the share equation.

The next-to-last term in equation (7) suggests that differences in levels of agricultural total factor productivity (TFP) (θ_{ct}^A) between countries can increase the export share of production in the processed food sectors. Ignoring country and time subscripts, note that $\partial G/\partial p_I < 0$, and $\partial p_I/\partial\theta^A < 0$. Therefore, the parameter α_{IX} is expected to be positive.

[9] Note: (*a*) land is not considered as an input into the food processing sector, (*b*) imported intermediate inputs are different from domestic agricultural intermediate inputs, and (*c*) nonagricultural intermediate inputs are not considered due to data constraints.

[10] In a global context with open economies, productivity growth in primary agriculture and the lower price and income elasticities (Engel's law) of food demand in the United States, as well as its competitors, have led to a global decline in the prices of agricultural output.

Effectively, it captures the transfer of comparative advantage from primary agriculture to the food processing sectors due to technological progress. Hence, cross-country differences in the levels of agricultural and food processing TFP are considered as explanatory variables in the latter's export share of production. The last term (ε_{ct}^{F}) in (7) is the random error.

Data and Estimation Procedures

Estimation of (5) and (7) requires internationally comparable data. Such a set of comparable accounts for developed countries is provided by the Organization for Economic Cooperation and Development (OECD). The OECD databases used in this study include: (*a*) the 1998 *International Sectoral Database* (*ISDB*), (*b*) the 1997–98 *Stan Database for Industrial Analysis* (*SDIA*), and (*c*) the 1996–98 *Economic Accounts for Agriculture* (*EAA*). Based on availability and compatibility, a panel data set is assembled comprised of annual data from 1975–95 for 13 OECD countries: Australia, Belgium, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Sweden, the United Kingdom, and the United States.

Variables defined below are used in the estimation of both the agricultural export share of agricultural GNP and the food export share of food GNP equations. Table 1 presents descriptive statistics on the data used in this study.

The total agricultural output data are taken from the OECD's *EAA* database, while the *SDIA* is used to obtain final food processing output, both denominated in current U.S. dollars for all countries. Exports and imports of primary agricultural and processed food products (ISIC 1 for agriculture and ISIC 31 for food manufacturing excluding tobacco) for all 13 countries (in current U.S. dollars) are obtained from the Economic Research Service, U.S. Department of Agriculture (USDA) (Gehlhar).

Sectoral gross capital stock (constant 1990 U.S. dollars) and total employment data for both sectors are taken from the OECD's *ISDB* database for 1975–95. Data on arable land, included only in the agricultural export share equation, are obtained from the 1997 *FAOSTAT* statistical database (Food and Agriculture Organization of the United Nations). The real exchange rate is taken from various issues of the International Monetary Fund's *International Financial Statistics.*

Total factor productivity (TFP) levels are computed based on indexes developed by Caves, Christensen, and Diewert, and applied by Harrigan. Let value added (VA) be a function of labor (L) and capital (K). Ignoring subscripts for industry and time, the index for any two countries a and b is given by:

$$(8) \qquad TFP_{ab} = \frac{VA_a}{VA_b}\left(\frac{\bar{L}}{L_a}\right)^{\sigma_a}\left(\frac{\bar{K}}{K_a}\right)^{1-\sigma_a}\left(\frac{L_b}{\bar{L}}\right)^{\sigma_b}\left(\frac{K_b}{\bar{K}}\right)^{1-\sigma_b},$$

where $\bar{L}$ and $\bar{K}$ are the geometric mean of labor and capital, respectively, over all observations in the sample, and $\sigma_c = [(s_c + \bar{s})/2]$ for $c = a, b$, where s_c is labor's share of value added, and $\bar{s}$ is the geometric mean of s_c over all observations in the sample. The index in equation (8) is superlative, meaning it is exact for the translog functional form. In addition, index (8) is transitive, i.e., $TFP_{ad} = TFP_{ab} * TFP_{bd}$, such that the choice of base year and country does not affect its computation. Relative TFP indexes are calculated

Table 1. Sample Means (Standard Deviations) for Selected Variables, 1975–95

Country	Export Share		TFP Level[a] (base year = 1975)		Labor[b] (000 employees)		Capital[c] (mil. 1990 US$)		Land (000 acres)
	Ag	Food	Ag	Food	Ag	Food	Ag	Food	
Australia	0.25	0.34	1.65	1.32	413	178	36,662	13,807	45,639
	(0.06)	(0.04)	(0.22)	(0.34)	(13)	(8)	(1,693)	(2,014)	(2,132)
Belgium	0.57	0.34	0.88	0.82	103	106	8,816	8,972	800
	(0.21)	(0.06)	(0.04)	(0.04)	(12)	(6)	(275)	(1,793)	(92)
Canada	0.35	0.11	1.83	0.82	573	261	56,183	16,377	45,558
	(0.05)	(0.03)	(0.18)	(0.09)	(20)	(10)	(829)	(202)	(521)
Denmark	0.11	0.48	2.68	1.71	168	91	24,132	8,862	2,572
	(0.03)	(0.06)	(0.14)	(0.44)	(31)	(4)	(1,268)	(805)	(88)
France	0.17	0.20	0.94	0.96	1,531	585	77,678	56,579	17,802
	(0.03)	(0.03)	(0.06)	(0.13)	(337)	(18)	(3,315)	(8,571)	(355)
Germany	0.08	0.12	2.84	0.81	1,184	865	141,900	65,294	11,915
	(0.03)	(0.02)	(0.16)	(0.07)	(242)	(40)	(144)	(269)	(169)
Italy	0.09	0.10	2.55	0.77	2,549	401	193,839	38,950	9,092
	(0.01)	(0.02)	(0.36)	(0.08)	(424)	(27)	(30,437)	(8,079)	(361)
Japan	0.01	0.01	2.65	0.64	6,542	1,533	341,107	77,054	4,195
	(0.01)	(0.01)	(0.60)	(0.28)	(1,186)	(145)	(80,485)	(26,831)	(125)
Netherlands	0.58	0.59	1.05	1.23	267	166	26,136	16,609	834
	(0.06)	(0.08)	(0.06)	(0.31)	(10)	(10)	(3,911)	(4,094)	(45)
Norway	0.01	0.06	2.25	1.84	146	52	17,358	5,716	860
	(0.01)	(0.02)	(0.32)	(0.38)	(17)	(3)	(1,921)	(801)	(44)
Sweden	0.05	0.07	1.85	1.27	198	82	15,720	6,109	2,896
	(0.02)	(0.01)	(0.17)	(0.26)	(35)	(7)	(1,191)	(587)	(84)
UK	0.10	0.07	1.27	1.12	610	638	40,197	48,267	6,717
	(0.03)	(0.01)	(0.09)	(0.12)	(44)	(78)	(1,023)	(4,528)	(351)
USA	0.16	0.04	1.60	0.75	3,125	1,678	445,950	188,950	186,927
	(0.03)	(0.01)	(0.22)	(0.11)	(91)	(44)	(34,048)	(28,527)	(1,253)

[a] TFP level is relative to the UK (base year 1975 = 1.00).
[b] Labor denotes number of full-time equivalent employees (in thousands).
[c] Capital denotes gross capital stock (in millions of 1990 US$).

using (8) and value-added, labor, and capital data available from the *ISDB* for both the agriculture and food processing sectors.[11]

A number of estimation issues arise with export share equations (5) and (7). The first is the use of a lagged dependent variable to capture the rate of adjustment in the shares of exports in GNP. In using a one-period lag, the dependence of regressors (lagged dependent variable) and the error yields inconsistent parameter estimates, which can be corrected by instrumenting the lag with the two-period lag (for details on consistency, see Hsiao, chapter 4; Nerlove).

Second, we address the question of whether a fixed- or random-effects specification is superior. For instance, Greene (p. 469) notes, "... it might be appropriate to view individual specific constant terms as randomly distributed across cross-sectional units. This

[11] Ball et al. compare levels of farm productivity using appropriately constructed data on purchased inputs and land quality. Based on their findings, the level of U.S. agricultural TFP (1973–93) is on average lower than several developed countries, few of which are the same as in the current study. In fact, the United States ranks No. 8 out of the 10 countries compared in the Ball et al. estimates. In this study, the United States ranks No. 9 out of the 13 countries compared. Observations such as the average level of Denmark's (UK's) TFP is greater (smaller) than that of the United States also suggest similarity.

would be appropriate if we believed that the cross-sectional units are drawn from a large population." Mundlak argues the individual effects should be treated as random, because the fixed-effect model is simply analyzed conditionally on the effects present in the observed sample. In addition, the fixed-effects model is costly in terms of degrees of freedom lost. On the other hand, there is no justification for treating the individual effects as uncorrelated with other regressors, as is assumed in the random-effects model. Hence, in our empirical analysis, we use a Hausman test to identify any misspecification resulting from the use of the random-effects models.

The third estimation issue addressed here is the possibility of measurement errors in the computation of TFP. An instrument such as

$$\theta_{ct} = \frac{1}{(C-1)} \sum_{d \neq c}^{C} \theta_{dt},$$

where C denotes the total number of countries (13), is appropriate for these models (Harrigan). A major share of the sample consists of EU countries. Consequently, the implications of CAP for the econometric specifications need to be explored.

Finally, the validity of imposing a homogeneity condition on the parameters of factor endowments on equations (5) and (7) should be subjected to an F-test. In the following section, F- and Hausman test results are reported before discussing the parameter estimates of export share equations (5) and (7).

Estimation Results

The regression results of the fixed- and random-effects models are summarized in tables 2 and 3 for the export share equations of the agricultural and processed food sectors, respectively. Parameter estimates and standard errors are reported, along with the R^2 for each model. As noted earlier, the random-effects model may be misspecified due to correlation between the panel-specific component of error and the explanatory variables, which can be verified by a Hausman test. Using this test, the null hypothesis of "no misspecification" in the random-effects model is rejected because the calculated $\chi^2_{[6]}$ values for agriculture (46.248) and food (63.465) are greater than the critical value (12.591) at the 1% level of significance. Only country-specific fixed effects are included because the addition of time-specific effects (dummies) neither improved the fit (R^2) nor affected the results. The two fixed-effects specifications yielded high R^2 statistics as well.[12]

The restriction that the intercepts of European countries under CAP are equal is rejected by an F-test ($F_{8,241} = 8.195$), suggesting different initial conditions for these countries. Testing for the differential impact of factor endowments and productivity on export shares of CAP and non-CAP countries is carried out using slope dummies. The null hypothesis that the slope dummies are equal to zero is rejected by an F-test ($F_{6,235} = 12.344$). This result does not suggest the CAP does not affect EU agriculture, because the independent variables without country dummies explained only 55% (60%) of the variation in the agricultural (food) export shares of GNP. We are unable to decompose the additional gain in R^2 (38%–39%), which comes from a variety of sources including

[12] We also tested for the restriction that the parameters on factor endowments sum to one, which is required to ensure homogeneity conditions of the translog GNP function. F-test results indicated the restriction is valid for the agricultural export share equation, but not for the food export share equation. However, we only report results from restricted models.

Table 2. Parameter Estimates of Agricultural Export Share Equations: Fixed- and Random-Effects Models

Variable	Fixed Effects Estimate	Fixed Effects Std. Error	Random Effects Estimate	Random Effects Std. Error
Intercept[a]	-0.0731	0.1727	0.0066	0.1100
Lagged Share of Exports	0.4596**	0.0527	0.8407**	0.0470
Ag Labor	-0.0152	0.0422	-0.0023	0.0033
Ag Capital	0.0182*	0.0096	0.0025	0.0042
Land	-0.0030	0.0185	-0.0002	0.0016
Exchange Rate	-0.0346**	0.0093	-0.0110	0.0071
Ag TFP	0.0159*	0.0085	0.1144**	0.0168
Fixed Effects:				
Australia	0.1448**	0.0447		
Belgium	0.3308**	0.1145		
Canada	0.1984**	0.0330		
Denmark	0.0591**	0.0097		
France	0.1055**	0.0132		
Germany	0.0512**	0.0061		
Italy	0.0507**	0.0128		
Japan	0.0033	0.0034		
Netherlands	0.3145**	0.0651		
Sweden	0.0355**	0.0174		
UK	0.0664**	0.0125		
USA	0.0871**	0.0302		
R^2 Statistic	0.9281		0.9193	
Hausman Test: $\chi^2_{[6]} = 46.248$, p-value = 0.0001				

Note: Single and double asterisks (*) denote significance at the 10% and 5% levels, respectively.
[a] The dummy variable for Norway is dropped.

policy structure. Panel-specific serial correlation was detected only in the agricultural export share equations, and was corrected for in the final results. The serial correlation coefficients ranged from 0.23 to 0.87, and are significant at the 1% level.

Looking first at the fixed-effects specification of the share of exports in agricultural GNP, note that four out of six coefficients are significant at the 5% or 10% level (table 2). Moreover, with the exception of Japan, all country dummies are significant. Current export shares are significantly influenced by lagged shares, and the magnitude of the parameter (0.4596) indicates the rate of adjustment is relatively fast (a value closer to one suggests sluggish adjustment, while a value closer to zero denotes rapid adjustment). This result is likely due to the accumulation of large stocks of grains and other bulk commodities by the sample countries (USDA 1993). Moreover, both European countries and the United States made extensive use of export subsidies during the sample period.[13]

[13] Tracking all components of the export assistance, including subsidies, promotion, export credits, and guarantees, for the time period of our study (1975–95) is a formidable challenge. As noted in the theory section, price differences among the sample countries can be captured only by fixed effects in our model.

Table 3. Parameter Estimates of Food Export Share Equations: Fixed- and Random-Effects Models

	Fixed Effects		Random Effects	
Variable	Estimate	Std. Error	Estimate	Std. Error
Intercept[a]	0.0583	0.1257	0.2002*	0.1367
Lagged Share of Exports	0.5753**	0.0472	0.8945**	0.0287
Food Labor	-0.0220**	0.0107	-0.0005	0.0119
Food Capital	0.0220**	0.0107	0.0005	0.0119
Exchange Rate	-0.0673**	0.0163	-0.0448**	0.0168
Ag TFP	0.0577**	0.0247	0.0404	0.0332
Food TFP	0.0437*	0.0301	0.0104	0.0455
Fixed Effects:				
Australia	0.1382**	0.0175		
Belgium	0.1319**	0.0165		
Canada	0.0431**	0.0111		
Denmark	0.1934**	0.0207		
France	0.0673**	0.0107		
Germany	0.0401**	0.0096		
Italy	0.0206**	0.0085		
Japan	-0.0144	0.0015		
Netherlands	0.2387**	0.0254		
Sweden	0.0197**	0.0092		
UK	0.0190**	0.0094		
USA	-0.0078	0.0084		
R^2 Statistic	0.9916		0.8330	
Hausman Test: $\chi^2_{[6]} = 63.465$, p-value = 0.0001				

Note: Single and double asterisks (*) denote significance at the 15% and 5% levels, respectively.
[a] The dummy variable for Norway is dropped.

The effect of agricultural capital (0.0182) is positive and significant, while the parameter estimates for labor and land are negative and insignificant. Therefore, as capital accumulates in the agricultural sector, a country is more likely to export agricultural products (a Rybczynski-type effect as in the Heckhser-Ohlin model). Note, factor intensities here are measured relative to the domestic (import-competing) goods sector. The insignificant effects of land and labor on agricultural export shares (not exports) indicate export production is less land- and labor-intensive relative to capital.[14] The addition of services such as transportation, insurance, and others may likely make export production less land-intensive relative to domestic goods.

Appreciation of a country's currency, as represented by the real exchange rate index (table 2), has negative effects on export shares (-0.0346), a result commonly observed by other studies (Chambers and Just). The effect of agricultural TFP is positive (0.0159) and significant. Within the group of countries considered here, technological progress

[14] The effects of factor endowments on export shares are relative because of the imposition of a homogeneity condition. While the possibility that Rybczynski elasticities can be negative has been established in the general case (n goods and n inputs), the negative effects on export shares need not translate into negative effects on export supply.

is biased in favor of exports; those countries with higher rates of TFP growth in agriculture are more likely to export bulk commodities, and thereby gain comparative advantage. This result lends credibility to the Ricardian assertion that technological differences affect specialization.

In the case of the fixed-effects model of food export shares (table 3), all coefficients are significant at the 5% level except the coefficient for food TFP, which is significant at the 15% level. Once again, the lagged share variable is important in the estimation of current export shares (0.5753). Compared to agriculture, however, the rate of adjustment is relatively slower for processed food export shares. As noted earlier, publicly supported stocks of bulk commodities and export subsidies may explain the divergence in the rates of adjustment between the agricultural and food processing sectors.

A Rybczynski-type effect emerges again from the model, suggesting food export production is more intensive in its use of capital (0.0220) than labor (-0.0220). This result is hardly surprising, as developed countries' food processing sectors are typically capital intensive. The effect of the real exchange rate (-0.0673)—i.e., appreciation lowers export shares—is consistent with prior studies.

A most striking result from the food export share equations is the effect of agricultural TFP (0.0577) which is positive and significant (a Ricardian effect), suggesting productivity growth in agricultural sectors benefits the expansion of food sectors' export shares. This positive effect is consistent with supply shifts due to productivity growth, resulting in lower agricultural prices passed on to the food processing sectors. By lowering the agricultural input costs, this transfer enables food processing sectors to compete more effectively in foreign markets. Thus, the food export share of GNP rises. In essence, comparative advantage in primary agriculture is transferred to the food processing sector, which, in turn, achieves higher export shares. Finally, food TFP has a positive effect (0.0437), implying some bias in food industries' technology in favor of export production.

Table 4 presents the long-run effects and standard errors of factor endowments, TFP, and the real exchange rate on export shares of the agricultural and food sectors. Dividing the parameter estimates of tables 2 and 3 (fixed-effects model) by one minus the respective parameter estimates on lagged shares (lagged dependent variable) provides these long-run effects. The standard errors are computed using Fieler's theorem (Arnade and Gopinath). Note the pattern of significance exhibited in tables 2 and 3 also carries over to the long-run effects. Given the relatively faster rates of adjustment of export shares in both sectors to their respective equilibrium values, the long-run effects are two to three times larger than those for the short run.

The effects of TFP and capital on the evolution of agricultural export shares are consistent with the notion that developed countries have experienced significant capital deepening and technological progress over the past few decades (Ball et al.). The effects of land and labor on agricultural export shares remain insignificant. Growth in capital, agricultural TFP, and food TFP are significant growth sources for processed food export shares in the long run.

The short- and long-run effects reported in tables 2, 3, and 4 suggest significant roles for factor endowments, productivity, and exchange rate movements in the evolution of export shares of the agricultural and processed food sectors. The parameter estimates in tables 2 and 3 can also be used to decompose the fitted export shares into their components for both sectors (Gopinath and Roe). As noted earlier, the country-specific

Table 4. Long-Run Effects of Factor Endowments and Technology on Export Shares

Variable	Agricultural Export Shares	Variable	Food Export Shares
Ag Labor	-0.0281	Food Labor	-0.0518**
	(0.0428)		(0.0099)
Ag Capital	0.0337**	Food Capital	0.0518**
	(0.0078)		(0.0099)
Land	-0.0056	Exchange Rate	-0.1585**
	(0.0181)		(0.0190)
Exchange Rate	-0.0640**	Ag TFP	0.1359**
	(0.0077)		(0.0257)
Ag TFP	0.0294**	Food TFP	0.1029**
	(0.0085)		(0.0321)

Notes: Double asterisks (**) denote significance at the 5% level. The long-run effects are computed by dividing parameter estimates (fixed effects) in tables 2 and 3 by one minus the parameter estimate on the lagged dependent variable. The standard errors (in parentheses) are computed using Fieler's theorem:

$$[\mathrm{var}(a/b) = \mathrm{var}(a) + (a/b)^2\mathrm{var}(b) - 2(a/b)\mathrm{cov}(a, b)].$$

Table 5. Contributions of Factor Endowments and Technology to Predicted Shares

Variable	Agricultural Export Share (%)	Variable	Food Export Share (%)
Lag Share	45	Lag Share	55
Ag Labor	-105	Food Labor	-142
Ag Capital	188	Food Capital	215
Land	-14	Exchange Rate	-159
Exchange Rate	-82	Ag TFP	76
Ag TFP	68	Food TFP	55
Sum:	100%	Sum:	100%

effects accounted for 38% and 39% of the variation in agricultural and food export shares, respectively.

The results presented in table 5 break up the remaining 55% (60%) of the explained variation in agricultural (food) export shares into its components. The largest single contributor to the increase in the agricultural and food export shares is growth in capital stock (188% and 215%, respectively). However, the departure of labor from the food and agricultural sectors had negative effects on export shares. The total (Rybczynski) effect of the factor endowments is 69% (-105 + 188 - 14) for agricultural export shares, while that of processed food export shares is 73% (-142 + 215).

Exchange rate effects are negative, but larger, on food export shares relative to agricultural export shares (table 5). The effect of agricultural TFP on export shares of both sectors is large and comparable to the total effect of endowments. The Ricardian effects are approximately equal to the Rybczynski effects in the agricultural sector, while the former is relatively larger for the processed food sector.

In sum, some factor endowments and technology are shown as important determinants of a country's pattern of agricultural trade. The effects of relative expansion in capital and food labor endowments, and TFP offset the negative effects of real exchange rate movements on the export shares in both sectors.

Conclusions

The hypothesis of this study is that productivity growth and intersectoral linkages explain agricultural and processed food export patterns and performance. Comparative advantage in the context of open economies is not necessarily static, especially when knowledge, and physical and human capital can be accumulated. Comparative advantage can be changed by conscious efforts directed toward the accumulation of various factors, including knowledge. Using a panel database (13 countries over the 1975–95 time period), this study identified and assessed the supply-side effects—productivity growth, intersectoral linkages, and factor endowments—on export patterns and performance of the agricultural and processed food sectors. Due to data constraints, the effects of policy and demand-side phenomena are proxied by country-specific effects in the panel econometric procedures used here.

We find evidence of both Rybczynski-type (Hecksher-Ohlin) and Ricardian (technology) effects in agricultural and processed food trade of OECD countries. Both effects contribute significantly to the growth in the export shares of the two sectors. There is evidence of transfer of comparative advantage. The efficiency gains in the primary agricultural sector are transferred to the food processing sector, allowing food processors to compete more effectively in foreign markets. Thus, policies which tend to distort markets, for example, by raising the price of primary agricultural products, can adversely affect the competitiveness of processed food sectors. However, policies supporting R&D in agriculture can bring about comparative advantage in processed food sectors.

Future studies may focus on modeling the demand and policy-side effects on trade patterns more explicitly, and explore their relative contributions to growth in exports. Innovations in empirical procedures (e.g., simultaneous or multiple-equation panel model procedures) and data construction (e.g., common-currency price data, details on intermediate input use and prices, protection measures) would greatly aid further investigations of trade patterns.

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