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International Trade in Manufactured Products: A Ricardo-Heckscher-Ohlin Explanation with Monopolistic Competition

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Abstract

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A large data set on trade in manufactured products is used to evaluate the performance of a model that combines both the Ricardian and Heckscher-Ohlin effects and incorporates monopolistic competition. The paper estimates a relation implied by the model to explain relative sectoral exports of major countries to a number of important markets, using 1970–90 data for nine manufacturing sectors. The relation fits the data well and variables suggested by both traditional and new trade models play an important role in explaining relative exports.

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I. Introduction

What factors explain international trade flows is a basic issue that has been addressed by a large body of empirical research. Early research focused on testing simple explanations suggested by the Ricardian and Heckscher-Ohlin models but did not yield clear-cut conclusions. MacDougall's (1951,1952) well-known test of the Ricardian model found a positive cross-industry association between the ratio of US to UK labor productivity and the US-UK export ratio. Although this evidence is still widely cited, it had little influence on the subsequent research because MacDougall's test was not rigorously derived from the Ricardian model and the model's one-factor framework was considered too simple to adequately explain international trade flows. The multi-factor Heckscher-Ohlin model provided a more appealing framework but Leontief's (1953) paradoxical findings that US export goods were less capital intensive than US import-competing products raised serious doubts about the empirical validity of the basic Heckscher-Ohlin model.

Current research has focused mainly on the Heckscher-Ohlin-Vanek (HOV) version that explains trade in factor content. This version initially appeared to rescue the Heckscher-Ohlin theory by suggesting a resolution of the Leontief paradox.[2] But further research on the HOV model [e.g., Bowen, Leamer and Sveikauskas (1987)] soon demonstrated major departures from the basic model that assumes the same technology and factor prices everywhere. Recently, however, a number of studies have been successful in improving the empirical performance of the HOV model by refining and modifying it in a number of ways. Trefler's (1995) influential study identifies key discrepancies between the data and the predictions of the HOV model and explores the potential of several simple amendments of the model's assumptions (regarding technology and absorption) to improve its fit. He shows that the assumption of Hicks-neutral technical differences across countries performs the best, as compared to a number of other modifications of the identical technology assumption. Davis and Weinstein (1998) examine a number of further variations of the HOV model. They find that the performance of the model (with two factors) improves considerably if international differences in production techniques attributed to unequal factor prices are incorporated in the model in addition to the Hicks-neutral technical differences.

The introduction of more general differences in production techniques between countries also helps the model explain certain features of the data. Hakura (1999, 2000) shows, for example, that the failure of a bilateral version of the HOV model can be largely explained by accounting for actual differences in production techniques within the European Community. In related research on international specialization in production, Harrigan (1997) finds that, in addition to relative factor endowments, Ricardian type non-uniform (Hicks-neutral) technology

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[2] Leamer (1980) showed that in the presence of more than two factors, Leontief's findings were not inconsistent with the HOV model. Brecher and Choudhri (1982), however, noted another paradox about Leontief's results produced by the HOV model.

differences across sectors are important in explaining cross-country variation in output shares of different sectors.

The above evidence favors a generalized Heckscher-Ohlin framework with (possibly Ricardian type) international differences in technology and unequal factor prices between countries. The ability of such a framework to explain international trade in goods, however, remains largely unexplored.[3] One basic problem in pursuing this question is that both the Heckscher-Ohlin and Ricardian models assume perfect competition and it is difficult to derive clear predictions about trade flows in individual sectors under this assumption. New trade models have emphasized the role of product differentiation and monopolistic competition and these factors can be fruitfully incorporated in traditional models to derive testable implications for goods trade.[4] Choudhri and Schembri (2000) have shown, for example, that a MacDougall type relation can be rigorously derived from the Ricardian model by adding monopolistic competition to the model. Applying such a relation to Canada-US trade, they find that relative productivity differences are a significant determinant of the two country's relative shares in each country's markets.

The present paper builds on the Choudhri-Schembri approach to explore how well a model that combines both the Ricardian and Heckscher-Ohlin effects and incorporates monopolistic competition would explain international trade in manufactured goods. The paper empirically implements a relation implied by the model, which (like the MacDougall relation) explains relative sectoral exports of two countries to a third country (or a group of countries). This relation includes variables that capture not only the Ricardian and Heckscher-Ohlin effects but also the influence of monopolistic competition. A large trade data set—that includes exports of six major countries to a number of important markets—is used to evaluate the performance of the Ricardo-Heckscher-Ohlin model with monopolistic competition. The export relation fits the data well and all variables in the relation are found to be significant determinants of relative exports.

The theory underlying the export relation is discussed briefly in Section II. Section III discusses the methodology for empirical implementation of the export relation using OECD's International Sectoral Database. Section IV describes key features of the data for our sample

[3] There is considerable empirical literature, however, that estimates cross-industry regressions to examine the relation between some measure of export performance (e.g., net exports) and industry characteristics such as capital and skill intensities [see Deardorff (1984) for a review of early literature]. These regressions, however, do not have a clear theoretical basis, as discussed by Leamer and Bowen (1981), for example.

[4] As Helpman and Krugman (1985) show, however, monopolistic competition does not affect the predictions of the basic HOV model for trade in factor content. This result perhaps explains why new trade models have not received much attention in the empirical research on the HOV model.

that includes France, West Germany, Italy, Japan, UK and US. The key results are presented in section V, which discusses the estimation of the basic regressions. Some variations of the basic regressions are considered in section VI. Conclusions are offered in section VII.

II. Theory

This section discusses a model that incorporates monopolistic competition into the Ricardo-Heckscher-Ohlin framework and uses this model to derive a relation explaining relative exports of two countries to a third country (or a group of countries). Let there be I monopolistically-competitive industries in J countries. Assume that consumer demand for each industry's varieties is based on the Dixit-Stiglitz (1977) model of the utility function with symmetrically differentiated varieties. To simplify the exposition, suppose initially that no intermediate goods are produced. The demand in country m for a variety produced in country j can then be expressed as

$$D_i^{jm} = E_i^m (P_i^j B_i^{jm})^{-\sigma} / \sum_{k\in J} n_i^k (P_i^k B_i^{km})^{1-\sigma} \ , \qquad (1)$$

where subscript $i = 1...I$ indexes industries (time subscripts are omitted for notational simplicity); E_i^m is the total expenditure in country m on all (domestic and foreign) varieties in the industry; P_i^j denotes the home price of each variety produced in country j 's industry: B_i^{jm} represents an index of industry trade barriers for country j 's exports to country m (so that $P_i^j B_i^{jm}$ represents the price of country j 's variety in country m 's market); n_i^j stands for the number of varieties in country j 's industry; and σ is the elasticity of substitution (assumed to be the same for all industries).

Letting $X_i^{jm} \equiv n_i^j P_i^j D_i^{jm}$ denote industry exports from country j to country m, the relative exports of a country pair (j, k) to country m 's market are determined by (1) as

$$X_i^{jm} / X_i^{km} = (n_i^j / n_i^k)(B_i^{jm} / B_i^{km})^{-\sigma} (P_i^j / P_i^k)^{1-\sigma} . \qquad (2)$$

Equation 2 represents the basic building block for our export relation. Factors emphasized by Ricardian and Heckscher-Ohlin models can be introduced into (2) by using the production conditions to link the price ratio to relative productivity and relative factor prices. Assume that each variety is produced by a distinct firm, requires a fixed amount of headquarter services and is manufactured at a plant under constant returns to scale. For country j, we express the plant production function as:

$$Q_i^j = \alpha_i^j F_i^{pj} , \qquad (3)$$

where Q_i^j is the plant output, and α_i^j and F_i^{pj} are the productivity and the quantity of the composite factor used in the plant. Letting $\mathbf{V}_i^{pj}$ represent a vector of primary factors employed at each plant, define $F_i^{pj} \equiv \phi_i(\mathbf{V}_i^{pj})$. The function $\phi_i(.)$ is assumed to be homogeneous of degree one and the same for all countries. Thus, the production relation (3) allows for only Hicks-neutral technology differences between countries. The cost of using one unit of the composite factor can be expressed as $C_i^j = \chi_i(\mathbf{W}^j)$, where $\mathbf{W}^j$ is the price vector for primary factors. In view of (3), the unit variable cost equals C_i^j / α_i^j. Profit maximization by the firm then implies that

$$C_i^j / \alpha_i^j = (1 - 1/\sigma) P_i^j . \tag{4}$$

Assume that headquarter operations require a fixed amount of the composite factor given by $\overline{F}_i^h \equiv \phi_i(\mathbf{V}_i^{hj})$, where $\mathbf{V}_i^{hj}$ represents the primary factors vector for headquarters.[5] Fixed headquarter costs thus equal $\overline{F}_i^h C_i^j$. The zero-profit condition can be stated as $\overline{F}_i^h C_i^j / Q_i^j + C_i^j / \alpha_i^j = P_i^j$. This condition along with (3) and (4) can be used to determine

$$F_i^{pj} = (\sigma - 1)\overline{F}_i^h . \tag{5}$$

Let A_i^j denote the industry total factor productivity (TFP) for country j. Note that $A_i^j = n_i^j Q_i^j / F_i^j$, where $F_i^j \equiv n_i^j (F_i^{pj} + \overline{F}_i^h)$ represents the total employment of the composite factor in the industry (since each firm uses the amount F_i^{pj} at the plant and $\overline{F}_i^h$ at the headquarters). Using (3) and (5), we can link the plant technology index to TFP as

$$\alpha_i^j = A_i^j \sigma / (\sigma - 1) . \tag{6}$$

We can also use (5) to relate the industry employment of the composite factor to the number of varieties (firms) in the industry as

$$F_i^j = n_i^j \sigma \overline{F}_i^h . \tag{7}$$

This relation allows us to use F_i^j as a proxy for (difficult-to-measure) n_i^j.

We now make use of (4), (6) and (7) to restate the export relation (2) as

$$X_i^{jm} / X_i^{km} = (F_i^j / F_i^k)(B_i^{jm} / B_i^{km})^{-\sigma} (C_i^j / C_i^k)^{-(\sigma-1)} (A_i^j / A_i^k)^{\sigma-1} . \tag{8}$$

[5] For simplicity, headquarter technology is assumed to be the same for all countries. Allowing $\overline{F}_i^h$ to vary across countries, however, would not substantively affect the export relation derived below.

In this relation, the Ricardian effects are represented by the TFP ratio, which is assumed to be determined exogenously. The Heckscher-Ohlin effects are captured by the composite-factor cost ratio. This ratio depends on relative factor prices and sectoral factor intensities (via the $\chi_i(.)$ function). Although factor prices are determined endogenously at the economy level, they (and thus the cost ratio) would be exogenous to an individual (small) industry.[6] The influence of the new trade theory is reflected in the composite-factor quantity ratio (or the ratio of number of varieties that it proxies), which is determined endogenously even at the industry level. We expect this ratio to be positively related to the relative size of the two countries.[7]

We can also distinguish two interesting special cases of our general model. The first case assumes Hicks-neutral technical differences to be uniform across industries (i.e., $\alpha_i^j = \alpha^j$). In this case, the TFP ratio would not vary across countries and the Ricardian effects would be absent. The second case assumes that the function defining the composite factor is the same for all industries [i.e., $\phi_i(.) = \phi(.)$] and thus there are no factor-intensity differences between industries. The cost ratio in this case would be identical in all industries and the Heckscher-Ohlin influences would be suppressed.[8]

Our model can be extended to include differentiated intermediate goods produced under monopolistic competition. Each firm in monopolistically-competitive industries can be viewed as producing its variety for final demand by consumers as well as intermediate demand by firms. Assume that each variety enters utility and production via the same Dixit-Stiglitz aggregator and (1) now represents country m 's total (final plus intermediate) demand for country j 's variety. Thus for monopolistically competitive industries, (2) still determines relative exports (now for both final and intermediate use) of a pair of countries.

On the production side, we continue to use (3) as the plant production function but redefine the composite factor as $F_i^{pj} \equiv \hat{\phi}_i(\mathbf{V}_i^{pj}, \mathbf{Z}_i^{pj})$, where $\mathbf{Z}_i^{pj}$ is the intermediate inputs vector with each element representing a quantity aggregator of (home and foreign) intermediates purchased from a particular industry, and the function $\hat{\phi}_i(.)$ is homogeneous of degree one. Assume that headquarters require $\overline{F}_i^h$ amount of the redefined composite factor, so

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[6] Relative factor endowments do not directly enter the export relation but they would exert an indirect influence via relative factor prices.

[7] Such a relation would potentially represent the home market effect discussed by Krugman (1980)

[8] In this case, $\mathbf{W}^j$ could still differ from $\mathbf{W}^k$ because of international differences in factor endowments and productivity. The cost ratio [$= \chi_i(\mathbf{W}^j)/\chi_i(\mathbf{W}^k)$], however, would not vary across industries since $\chi_i(.)$ would be the same for all i.

that $\overline{F}_i^h \equiv \hat{\phi}_i(\mathbf{V}_i^{hj}, \mathbf{Z}_i^{hj})$, where $\mathbf{Z}_i^{hj}$ is a vector representing the headquarter use of intermediate inputs. The cost of one unit of the composite factor changes to $C_i^j = \hat{\chi}_i(\mathbf{W}^j, \mathbf{P}_i^{zj})$, where $\mathbf{P}_i^{zj}$ denotes the price vector for intermediate inputs used in the industry. Each element of $\mathbf{P}_i^{zj}$ represents a price index that is dual to the quantity aggregator of the corresponding element of $\mathbf{Z}_i^{pj}$. Although the quantity and the cost of the composite factor now depend on the use and the price of intermediate inputs, conditions (4)–(7) continue to hold. Thus the relation explaining the export ratio remains the same and is still given by (8).

We can also modify the model to allow for some intermediates to be homogeneous goods produced by perfectly-competitive industries. In this case, $\mathbf{Z}_i^{pj}$ (and $\mathbf{Z}_i^{hj}$) can be redefined to include a subset that consists of homogeneous intermediate goods. The corresponding subset of $\mathbf{P}_i^{zj}$ would now simply represent the domestic price vector for these goods. While $\mathbf{Z}_i^{pj}$ and $\mathbf{P}_i^{zj}$ are now different, this modification would not alter the export relation.

III. Empirical Implementation

This section discusses the empirical implementation of the model for explaining exports of manufactured products by major countries. The export relation can be estimated in the following log-linear form:

$$\ln(X_{it}^{jm} / X_{it}^{km}) = \beta_0 \ln(B_{it}^{jm} / B_{it}^{km}) + \beta_1 \ln(F_{it}^j / F_{it}^k) + \beta_2 \ln(A_{it}^j / A_{it}^k) + \beta_3 \ln(C_{it}^j / C_{it}^k) + e_{it}^{jkm}, \quad (10)$$

where $\beta_0 = -\sigma$, $\beta_1 = 1$, $\beta_2 = \sigma - 1$, $\beta_3 = -(\sigma - 1)$, e_{it}^{jkm} is an error term and a time subscript has been added. This relation can be used to explain relative industry exports of any country pair (j, k) to a particular market (m) in time period t. Apart from trade barriers, explanatory variables in this relation require industry-level data only for countries j and k. OECD's International Sectoral Database (ISDB) provides sectoral data for member countries on a comparable basis. This data are available for a number of manufacturing sectors (generally at the 2-digit ISIC level) but cover only the value-added activity and two factors (capital and labor). We assume a Cobb-Douglas form for the function defining the composite factor. With additional assumptions, this form enables us to account for intermediate goods in our estimation based on ISDB data (supplemented by some data from other sources).

It is difficult to construct a satisfactory index of trade barriers (B_{it}^{jm}) that adequately captures all types of trade costs and border effects. Lacking such an index, our approach is to assume that relative trade barriers for an exporting pair (in a particular market at a given time) do not differ much from one sector to another and consider this variable to be invariant across sectors in our empirical analysis. However, we allow relative barriers to vary across markets, country pairs and time periods, and use fixed effects to capture this variation.

To explain the measurement of other variables, first define the industry use of the composite factor in country j in the presence of M intermediate goods as

$$\ln F_{it}^{j} = \theta_i^K \ln K_{it}^j + \theta_i^L \ln L_{it}^j + \sum\nolimits_{r\in M} \theta_i^{zr} \ln Z_{it}^{jr} , \qquad (11)$$

where K_{it}^j and L_{it}^j represent amounts of capital and labor used in the industry; Z_{it}^{jr} is an index (quantity aggregator) for amounts of industry r 's intermediate goods used in industry i, and θ_i^K, θ_i^L and θ_i^{zr} are the shares of capital, labor and industry r 's intermediate goods in the value of output (the sum of the shares equals one). The total factor productivity is given by

$$\ln A_{it}^j = \ln Q_{it}^j - \ln F_{it}^j . \qquad (12)$$

Lacking data on Z_{it}^{jr}, F_{it}^j and A_{it}^j cannot be directly estimated from (11) and (12). However, the Cobb-Douglas form can be exploited to estimate these variables via the following value-added function:

$$\ln Y_{it}^j = \ln \widetilde{A}_{it}^j + \widetilde{\theta}_i^K \ln K_{it}^j + \widetilde{\theta}_i^L \ln L_{it}^j , \qquad (13)$$

where Y_{it}^j is value-added output; $\widetilde{\theta}_i^K$ and $\widetilde{\theta}_i^L$ are shares of capital and labor in value added; and $\widetilde{A}_{it}^j$ is the TFP in the value added activity, which can be estimated from ISDB data. Letting θ_i^Y denote the share of value added in the value of output, we can use (11)–(13) to link the two measures of TFP as $\ln A_{it}^j = \theta_i^Y \ln \widetilde{A}_{it}^j$.[9] This relation can be utilized to estimate A_{it}^j and then (12) can be used to estimate F_{it}^j if data on Q_{it}^j are available. Also the TFP ratio is given by $\ln(A_{it}^j / A_{it}^k) = \theta_i^Y \ln(\widetilde{A}_{it}^j / \widetilde{A}_{it}^k)$.

The cost of one unit of the composite factor can be derived from (11) as

$$\ln C_{it}^j = \theta_i^K \ln R_t^j + \theta_i^L \ln W_t^j + \sum\nolimits_{r\in M} \theta_i^{zr} \ln P_{it}^{jr} , \qquad (14)$$

where R_t^j and W_t^j represent the country's rental and wage rates (assumed to be the same in all sectors because of free mobility of factors), and P_{it}^{jr} is the price index for Z_{it}^{jr}. In the estimation of our basic regression, we assume that all intermediate goods are produced in monopolistically-competitive industries and are traded. We suppose, moreover, that inter-country differences in

[9] Noting that $\theta_i^K = \theta_i^Y \widetilde{\theta}_i^K$ and $\theta_i^L = \theta_i^Y \widetilde{\theta}_i^L$, we can use (11) and (13) to express $\ln Q_{it}^j = \theta_i^Y \ln Y_{it}^j + \sum_r \theta_i^{zr} \ln Z_{it}^r = \theta_i^Y \ln \widetilde{A}_{it}^j + \ln F_{it}^j$. The link between the two measures can then be established by using (12).

the intermediate-goods price index are not too large.[10] We then simplify our estimation of the composite-factor cost ratio by letting P_{it}^{jr} be the same for all j. Note that the cost of the composite factor in value added equals $\ln \widetilde{C}_{it}^{j} = \widetilde{\theta}_{i}^{K} \ln R_{t}^{j} + \widetilde{\theta}_{i}^{L} \ln W_{t}^{j}$. Under our assumption that $P_{it}^{jr} = P_{it}^{kr}$, (14) and (15) imply that

$$\ln(C_{it}^{j} / C_{it}^{k}) = \theta_{i}^{Y} \ln(\widetilde{C}_{it}^{j} / \widetilde{C}_{it}^{k}). \tag{15}$$

Thus, we can use the data on value added activity to estimate the cost ratio.

We also consider a variant of our basic regression, which allows certain intermediate inputs (originating from sectors excluding manufacturing, mining and agriculture) to be non-traded goods. Let N represent the non-traded subset of M intermediate goods. We assume that this subset is produced under perfect competition, so that $P_{it}^{jr} = C_{rt}^{j} / A_{rt}^{j}$ for $r \in N$. Using this condition and (14), and recalling that $P_{it}^{jr} = P_{it}^{kr}$ for $r \notin N$, we can express the cost ratio as

$$\ln(C_{it}^{j} / C_{it}^{k}) = \theta_{i}^{Y} \ln(\widetilde{C}_{it}^{j} / \widetilde{C}_{it}^{k}) + \sum_{r \in N} \theta_{i}^{zr} \ln(C_{rt}^{j} / C_{rt}^{k}) - \sum_{r \in N} \theta_{i}^{zr} \ln(A_{rt}^{j} / A_{rt}^{k}) \tag{16}$$

This measure of the cost ratio includes the influence of relative TFP in non-traded intermediates. It is useful, however, to separate the effects of factor prices from those of TFP. Our empirical analysis thus uses an alternative measure of the cost ratio that excludes the third term on the right hand side of (16). This term is included in the corresponding measure of the TFP ratio. We further simplify our empirical analysis by aggregating all non-traded sectors into one sector.

IV. Data

This section briefly discusses some key features of our data. Further details are provided in the Data Appendix. We focus on exports of large countries that are likely to account for a significant share of most markets. We use a sample of six exporting countries, which includes France, West Germany, Italy, Japan, UK and US.[11] Exports of Italy are not as large as the other countries in the sample but this country is included to increase the variation in factor prices and productivity performance within the sample.

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[10] This index depends on the numbers as well as the prices of home and foreign varieties and is defined as $P_{it}^{jr} = \left[\sum_{k \in J} n_{rt}^{k} (P_{rt}^{k} B_{rt}^{kj})^{1-\sigma}\right]^{1/(1-\sigma)}$. B_{rt}^{kj} would vary across k and j but we assume that the sum in the square bracket is similar across j.

[11] These countries account for 44 to 48 percent of world exports in goods and services during the 1975-1990 period (based on data reported in various issues of the IMF's *World Economic Outlook*).

For these countries, W_t^j is measured by the annual wage in manufacturing expressed in US dollars while an estimate of the user cost of a comparable unit of capital (i.e., worth one US dollar in 1990 prices) in US dollars is used to measure R_t^j.[12] Using annual ISDB data for nine 2-digit ISIC manufacturing sectors, we averaged each sector's shares of capital and labor in value-added over the six countries and the 1970–1990 period to estimate $\tilde{\theta}_i^K$ and $\tilde{\theta}_i^L$, and use these estimates to calculate $\ln \tilde{A}_{it}^j$ and $\ln \tilde{C}_{it}^j$. The six-country 1970–90 averages of sectoral shares of value added in output (obtained from the OECD's STAN database) are used to estimate θ_i^Y. This estimate is then used to measure $\ln(A_{it}^j / A_{it}^k)$, and $\ln(C_{it}^j / C_{it}^k)$ under the assumption that prices of intermediate goods are the same in all countries.

ISDB does not provide data on real (gross) output by sector (Q_{it}^j). Lacking such data, we use the composite factor in value added (i.e., $\ln \tilde{F}_{it}^j \equiv \ln Y_{it}^j - \ln \tilde{A}_{it}^j$) to measure F_{it}^j. This measure would be a good proxy for the relative quantity of the composite factor if the ratio of the intermediate-goods to the value-added composite factor does not vary much across countries.[13] Another limitation of the ISDB data is that it converts real industry (value-added) outputs to internationally comparable units using GDP purchasing power parities. This procedure does not account for international differences in relative prices across sectors and would introduce errors in measures of $\ln \tilde{A}_{it}^j$ as well as $\ln \tilde{F}_{it}^j$. Our estimation method (discussed in section V) addresses this problem.

For each country in our sample, Table 1 shows the 1970–1990 averages of absolute and relative factor prices as well as aggregate measures of relative factor supplies and productivity performance based on data for all manufacturing. According to these measures, the US is the most capital abundant as well as the most productive country. It also has the highest wage rate and the lowest rental rate (and thus a rent/wage ratio that is substantially lower than other countries). There is considerable variation in aggregate capital/labor ratio and TFP across countries. These variables, however, are not strongly related. West Germany, for example, has a relatively high TFP but a relatively low capital/labor ratio. The data do suggest a negative

[12]This cost is calculated simply by multiplying the price of a comparable unit of capital in US dollars by the sum of the real interest rate and a fixed depreciation rate. These rates are measured as in Caballero and Lyons (1990).

[13] Using (11) and (13) and recalling that $\theta_i^K = \theta_i^Y \tilde{\theta}_i^K$ and $\theta_i^L = \theta_i^Y \tilde{\theta}_i^L$, we can express $\ln F_{it}^j = \ln \tilde{F}_{it}^j + (1-\theta_i^Y)\ln(Z_{it}^j / \tilde{F}_{it}^j)$, where $\ln Z_{it}^j \equiv \sum_{r \in M} [\theta_i^{zr} / (1-\theta_i^Y)] \ln Z_{it}^{jr}$ represents an aggregate quantity index of the intermediate-goods composite factor. If the ratio, $Z_{it}^j / \tilde{F}_{it}^j$, is the same for all j, $\ln(F_{it}^j / F_{it}^k)$ would indeed equal $\ln(\tilde{F}_{it}^j / \tilde{F}_{it}^k)$.

relation between the rent/wage and the capital/labor ratios as well as a positive relation between the absolute wage rate and the TFP.

Figures 1 and 2 explore the inter-industry variation in comparative productivity and costs for the six countries. Figure 1 exhibits the long-run behavior of each country's comparative TFP in value added [i.e., the 1970–90 average of $\ln(\widetilde{A}_{it}^{j} / \widetilde{A}_{it}^{k})$ with k = US] across the nine sectors, ordered according to German comparative TFP. If international differences in TFP were uniform across sectors, these comparative TFP relations would be flat and parallel to each other. As Figure 1 shows, however, the TFP relations crisscross each other and each country's comparative productivity varies considerably from one sector to another.

Figure 2 shows how each country's long-run comparative cost of the composite factor in value added [i.e., the 1970–90 average of $\ln(\widetilde{C}_{it}^{j} / \widetilde{C}_{it}^{k})$ with k = US] varies from one sector to another. The sectors are ordered according to capital intensity as measured by the share of capital in value added (the last sector has the lowest share). Since the US has the lowest rent/wage ratio, the sectoral comparative cost of each non-US country falls as the sectoral capital share decreases.[14] The figure suggests important inter-industry differences as comparative costs vary significantly moving from one end of the capital intensity scale to the other.[15] The figure does reveal substantial international differences in absolute costs of the composite factor. The cost in Italy and UK, for example, is about half of the US cost for most sectors.

We examine the relative exports of our sample countries to a number of markets. Individual countries in our sample are used as markets since they are generally large enough to import sizable amounts from other countries in the sample. In addition, we aggregated a number of small countries within a region to form large markets that are likely to be reasonably integrated. We consider three such multiple-country markets: (1) European Union (EU), comprising of Belgium/Luxembourg, Denmark, Greece, Ireland, Netherlands, Portugal and Spain; (2) Other Western Europe (OWE), consisting of Austria, Finland, Norway, Switzerland Sweden, and Turkey; and (3) East Asia (EA), including Hong Kong, Indonesia, Malaysia, Philippines, Singapore, South Korea, Taiwan and Thailand.

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[14] Using (15), we can express the comparative cost ratio as $\ln(\widetilde{C}_{it}^{j} / \widetilde{C}_{it}^{k}) = \ln(W_t^j / W_t^k) + \widetilde{\theta}_i^K [\ln(R_t^j / W_t^j) - \ln(R_t^k / W_t^k)]$. Given that the expression in the square bracket is positive (for k = US), this ratio increases in $\widetilde{\theta}_i^K$.

[15] Note, however, that if the most and the least capital-intensive sectors (i.e., Food, Beverages and Tobacco and Chemicals at the one end and Textiles at the other) are excluded, the range of variation in comparative costs is small (between 4 and 8 percent). Thus only a few sectors have significantly different factor shares.

V. BASIC RESULTS

As the export relation is based on a long-run model, we estimate it using 5-6 year averages of our annual data to smooth out the influence of short-run fluctuations. Our sample period is thus divided into four sub-periods, 1970–74, 1975–79, 1980–84 and 1985–90. Using the US as the reference country, we focus on five exporting pairs (with each pair combining a non-US country with the US).[16]

Letting δ^{jm} and δ_t^m denote country and time fixed effects for market m and simplifying our notation, we estimate the export relation (10) in the following form:

$$x_{it}^{jm} = \delta^{jm} + \delta_t^m + \beta_1 f_{it}^j + \beta_2 a_{it}^j + \beta_3 c_{it}^j + e_{it}^{jm}, \tag{17}$$

where $x_{it}^{jm} \equiv \ln(X_{it}^{jm} / X_{it}^{km})$, $f_{it}^j \equiv \ln(F_{it}^j / F_{it}^k)$, $a_{it}^j \equiv \ln(A_{it}^j / A_{it}^k)$, $c_{it}^j \equiv \ln(C_{it}^j / C_{it}^k)$; j = France, West Germany, Italy, UK, Japan; k = US; m = EU, OWE, EA, France, West Germany, Italy, UK, Japan; $i = 1,\ldots,9$; $t = 1,\ldots,4$; and j unequal to m. For each market, we pool the data for all pairs and time periods to estimate (17). Equations for all markets are estimated as a system to examine cross-market constraints and allow for the possibility that residuals are correlated across markets.[17] Country and time fixed effects are introduced into (17) to account for the influence of the omitted trade-barriers term, $\ln(B_{it}^{jm} / B_{it}^{jk})$, which is assumed to be the same for all industries. The fixed effects, however, could also pick up the effect of short-run departures that are not eliminated by averaging annual data over 5-6 year intervals.

A problem with estimating (17) is that (as discussed in section II) the composite-factor size variable (f_{it}^j) is endogenously determined. To deal with this problem, we use a set of instrumental variables that capture the effect of the relative size of the exporting countries but do not depend on sector-specific variables (and thus are likely to be independent of the error term). This set is created by interacting relative real GDP and population size (in logs) with industry dummy variables.[18] Although the Ricardian approach treats the productivity variable

[16] The use of the US as a reference country ensures that each pair contains countries that have significantly different factor prices and productivity.

[17] Note that the system of equations represents an unbalanced sample because regressions for aggregate markets include all five pairs while regressions for individual markets exclude the pair formed by the home country.

[18] The set is thus defined as $\ln(Y_t^j / Y_t^k)ID_1,\ldots,\ln(Y_t^j / Y_t^k)ID_9$, $\ln(PS_t^j / PS_t^k)ID_1,\ldots,$ $\ln(PS_t^j / PS_t^k)ID_9$, where Y_t^j and PS_t^j denote real GDP and population size for country j in period t, and $ID_1 \ldots ID_9$ are the dummy variables for the nine industries.

(a_{it}^j) as an exogenous variable, it is often viewed as an endogenous variable that is potentially influenced by sector-specific factors. We consider both possibilities. For the case of endogenous productivity, we develop an additional set of instrumental variables that represent interactions between relative R&D effort in aggregate manufacturing and industry dummy variables (these are likely to be correlated with relative sectoral productivity and would be independent of sector-specific variables).[19] Our instruments would also address the problem (mentioned above) that the ISDB value-added data does not adequately account for international differences in relative prices and could introduce sector-specific measurement errors in estimates of the productivity (as well as the size) variable.

The first two columns of Table 2 present the basic results of estimating (17) as a system using three stage least squares (to allow for instrumental variables). These results show the estimates of the export relation under the restriction (implied by the model) that the coefficient of each of the three variables is the same in all markets. Country and time dummy variables are used to estimate fixed effects but these effects are not reported in the table. Column 1 of the table shows estimates for the case of exogenous productivity and column 2 for the endogenous-productivity case. In both of these regressions, the size variable is assumed to be endogenous while the cost variable (c_{it}^j) is considered to be exogenous.

The model implies that the coefficient of the size variable equals one while coefficients of the productivity and cost variables equal $\sigma - 1$ and $-(\sigma - 1)$, respectively. The results in both columns 1 and 2 are consistent with these predictions. In conformity with the theory, the effect of the size variable is significantly positive and not significantly different from one.[20] Moreover, as the theory predicts, the productivity variable exerts a significantly positive effect, the cost variable exerts a significantly negative effect, and the coefficient of the productivity variable is not significantly different from the absolute value of the cost-variable coefficient. The absolute values of the coefficients of productivity and cost variables are, in fact, identical in the endogenous-productivity regression and these values suggest an estimate of σ close to 3.5. To compare the explanatory power of the model in different markets, the table also shows R^2 for individual market regressions. The two Asian markets, especially Japan, exhibit the lowest R^2. One reason for this result may simply be that the sample consists mainly of European countries whose relative exports to Asian markets are much smaller than those to European markets and are thus likely to be more noisy.

To further explore inter-market differences, we also estimated unrestricted regressions that allow the coefficient of each variable to differ across the eight markets. Table 3 shows the

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[19] This additional set is defined as $\ln(RD_t^j / RD_t^k)ID_1, \ldots, \ln(RD_t^j / RD_t^k)ID_9$, where RD_t^j represents country j's real R&D expenditures in total manufacturing for period t.

[20] We simply use the conventional 5% level as the criterion for significance throughout the paper.

results for the case of endogenous productivity.[21] Estimates without cross-market restrictions indicate some departures from the predictions of the model. The coefficients of the size and cost variables differ significantly across markets. The size coefficient is substantially below one for two markets (Other Western Europe and East Asia) and substantially above one for another two (Italy and Japan). The productivity coefficient is significantly different from the absolute value of the cost coefficient in one market (EU). Although the data for certain markets does not fully conform to the model, the coefficients of all three variables have the predicted signs in all markets and are significantly different from zero in all except two cases (the cost variable is not significant in the markets for Italy and UK). The average effect of each variable over all markets (as captured by the restricted regressions in Table 2), moreover, accords very well with the model.

Table 3 also reports country and time fixed effects for each market. The regressions include all country and time dummy variables except the time dummy for period 4 (note that the constant term is not included in the regression). Thus the country dummies show the country effect in the excluded period 4 while time dummies show the effect in a particular period additive to the excluded period. Country fixed effects are significant and their pattern across markets suggests a strong role for geographical location as a determinant of market access.[22] For instance, fixed effects for European countries are significantly positive in the case of all European markets but significantly negative in the case of East Asia and insignificant in the case of Japan. On the other hand, the effect for Japan is significantly positive in East Asia and significantly negative or insignificant in European markets. The magnitudes of country effects suggest large differences in relative barriers between European and Asian markets. If we assume, for example, that only relative trade barriers give rise to country effects [i.e., $\delta^{jm} = -\sigma \ln(B_i^{jm} / B_i^{km})$] and use an estimate of σ equal to 3.5 (as suggested by the restricted endogenous-productivity regression), the regression estimates of country effects (for period 4) suggest that while trade barriers for (say) France are about 50% of US barriers in EU, they are about 150% of US barriers in East Asia.

The time fixed effects do not differ significantly across markets. However, these effects do reveal significant shifts between periods. The time effects show a tendency to both increase over time and fluctuate from one period to another. A reduction in relative trade barriers due to regional trade liberalization in Europe and East Asia could account for the positive trend in the

[21] The results for the exogenous-productivity regressions are not much different but are not reported to conserve space.

[22] Gravity models highlight distance as an important determinant of trade costs. We explored the use of a variable based on relative distances as an alternative measure of relative trade barriers. This variable was significant but had lower explanatory power as compared to country dummies.

time effects.[23] Short-term influences related to cyclical factors or exchange rate fluctuations, however, would appear to be responsible for the up and down movements of these effects.

VI. Variations

This section explores a number of variations of the basic model. We estimated both restricted and unrestricted regressions for each variation. To simplify the discussion and facilitate comparisons of parameter estimates between different variations, we report the results only for the restricted regressions. Unrestricted regressions, however, yield similar conclusions.

First, we estimate two special cases of our model. The first case eliminates the Ricardian effects by assuming a Heckscher-Ohlin model with uniform Hicks-neutral technical differences. In this version of the Heckscher-Ohlin model, a_{it}^j in (17) is replaced by its average value over the nine sectors.[24] The second case suppresses the Heckscher-Ohlin effects by assuming a multifactor Ricardian model with the same composite-factor function for all sectors. Sectoral shares of capital and labor in this version of the Ricardian model are set equal to their average value over the nine sectors and these average shares are used to calculate c_{it}^j in (17).[25]

The results for the Heckscher-Ohlin and Ricardian models are shown in columns 3 and 4 of Table 2, respectively. We focus on the case of exogenous productivity, as this case is more appropriate for the Ricardian model and endogenous productivity is not a concern for the Heckscher-Ohlin model (because it suppresses sectoral TFP differences). The results are consistent with both models. For each model, the variable highlighted by the model has a significant effect in the predicted direction. The variable whose inter-sectoral variation is suppressed is not precisely estimated and is insignificant in each case. The relative explanatory power of the two models varies across markets and neither model dominates the other in terms of the R^2 for individual market regressions. The special models do not perform as well as the

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[23] We added interaction terms to allow time effects for Japan to differ from those of European countries in each market. These terms, however, were not significant and are not included in the regressions reported in the tables, for simplicity.

[24] Observed differences in this index are viewed as measurement errors (with a zero mean). Note that the average value of the index equals $\theta_i^Y \ln(\widetilde{A}_t^j / \widetilde{A}_t^k)$, where $\ln(\widetilde{A}_t^j / \widetilde{A}_t^k)$ is the average value of $\ln(\widetilde{A}_{it}^j / \widetilde{A}_{it}^k)$. Also, f_{it}^j in (17) is now measured as $\ln(Y_{it}^j / Y_{it}^k) - \ln(\widetilde{A}_t^j / \widetilde{A}_t^k)$.

[25] This index is the same for all sectors and equals $\theta^Y [\widetilde{\theta}^K \ln(R_t^j / R_t^k) + \widetilde{\theta}^L \ln(W_t^j / W_t^k)]$, where θ^Y, $\widetilde{\theta}^K$ and $\widetilde{\theta}^L$ are the average values of θ_i^Y, $\widetilde{\theta}_i^K$ and $\widetilde{\theta}_i^L$, respectively. Observed differences in shares are now considered measurement errors and estimates of a_{it}^j and f_{it}^j are based on the average values of the shares.

basic model. As the comparison of R^2 in columns 1, 3 and 4 of the table shows, the basic model outperforms the Ricardian model in all markets and the Heckscher-Ohlin model in all markets except Japan. The difference in R^2 between the basic and special models, however, tends to be small. It should be pointed out that our measures of the size and productivity variables are related in such a way that changes in one measure produce offsetting changes in the other. This linkage may explain why variations of the model that involve alternative measures of productivity do not lead to marked differences in R^2.[26]

Next we explore a variant of the model, which incorporates non-traded intermediate goods. For this variant, we simply consider one aggregate non-traded sector comprising of most sectors outside manufacturing, mining and agriculture. As discussed above, we simplify the empirical implementation of the model by assuming that non-traded goods are produced under perfect competition. The productivity and cost variables are redefined to include the influence of TFP and factor costs in the production of non-traded intermediate inputs.[27]

Estimates of the non-traded goods version are presented for both exogenous- and endogenous-productivity cases in columns 5 and 6 of Table 2. Inclusion of non-traded intermediates reduces the size of the productivity and cost coefficients in both cases but the reduction is substantial only in the case of endogenous productivity. The results remain consistent with the model. In particular, parameter estimates continue to satisfy the restrictions that the size coefficient equals one and the absolute values of the productivity and cost coefficients are the same. R^2 for individual markets generally improves but the improvement tends to be very modest. Thus accounting for non-traded intermediate goods adds little to the explanatory power of the model.

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[26] For example, letting a hat denote the variables for the special Heckscher-Ohlin model, our measures (as discussed in footnote 20) imply that $a_{it}^j = \hat{a}_{it}^j + \theta_i^Y \varepsilon_{it}^j$ and $f_{it}^j = \hat{f}_{it}^j - \varepsilon_{it}^j$, where ε_{it}^j [$\equiv \ln(\widetilde{A}_{it}^j / \widetilde{A}_{it}^k) - \ln(\widetilde{A}_t^j / \widetilde{A}_t^k)$] represents the difference between the value-added TFP measures for the basic and special models. If (as assumed by the models) the coefficients of the size and productivity variables equal 1 and $\sigma - 1$, respectively, then the net effect of ε_{it}^j via the two variables would equal $\theta_i^Y(\sigma - 1) - 1$ and would be small if the average value of this term is close to zero. The value-added TFP measure for the Ricardian model also differs from that of the basic model and this difference would also lead to similar offsetting effects.

[27] Letting N represent the set of non-traded sectors and ignoring purchases of intermediate goods by these sectors, the two measures are defined as $c_{it}^j \equiv \theta_i^Y \ln(\widetilde{C}_{it}^j / \widetilde{C}_{it}^k) + \theta_i^N \ln(\widetilde{C}_{Nt}^j / \widetilde{C}_{Nt}^k)$ and $a_{it}^j \equiv \theta_i^Y \ln(\widetilde{A}_{it}^j / \widetilde{A}_{it}^k) + \theta_i^N \ln(\widetilde{A}_{Nt}^j / \widetilde{A}_{Nt}^k)$, where $\widetilde{C}_{Nt}^j$ and $\widetilde{A}_{Nt}^j$ represent the cost and TFP of the non-traded sector's value-added and θ_i^N is the share of the non-traded sector in i's (gross) output.

VII. Conclusions

Traditional explanations of international trade suggested by Ricardian and Heckscher-Ohlin models emphasize international differences in technology and factor endowments as the key sources of international trade. New trade models based on differentiated products and monopolistic competition were initially viewed as providing an alternative account of international trade, but are now generally seen as suggesting explanations that are complementary to the traditional explanations. Although these new models have not had much influence on the empirical work on international trade in factor content, they help provide a theoretical basis for developing empirical relations that explain trade in goods.

The paper empirically implements a relation based on an integrated framework that combines Ricardian, Heckscher-Ohlin and monopolistic-competition models, and explains relative sectoral exports of a pair of countries to markets in the rest of the world. Controlling for size and trade barriers, relative exports of the two countries depend on their relative productivity and factor costs. Estimation of this relation requires data on productivity, factor cost and factor use only for the country pair.[28] This data is obtained from ISDB for a sample that includes five pairs of major exporting countries (with the US as the reference country), nine manufacturing sectors and four (5-6 year) time periods. The export relation is estimated using trade data that covers relative exports of our sample countries to eight multi- and single-country markets.

Estimates of the export relation conform very well to the integrated model, especially for regressions that impose cross-market restrictions implied by the model. As predicted by the Ricardo-Heckscher-Ohlin framework, the productivity variable exerts a positive effect and the cost variable a negative effect on relative exports. Also, as suggested by the monopolistic-competition model, both the size variable and the country and time fixed effects (introduced to capture trade barriers) are found to be significant determinants of relative exports. The pattern of country fixed effects across different markets suggests that geographical location is an important (but not the only) source of these effects. Thus variables suggested by both traditional and new trade models play an important role in explaining trade in goods.

It should be emphasized that although the paper's empirical model identifies size, productivity and factor costs at the sectoral level as important influences on sectoral exports, it does not provide an explanation of why these variables differ across countries. Explaining the sources of these differences would be an interesting topic for future research.

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[28] This data requirement is much less stringent than for relations determining the factor content of net exports, which (in the presence of international differences in technology or factor prices) require data on production techniques used by all trading partners that supply imports.

Table 1. Selected Characteristics of Sample Countries: 1970–90 Averages

Country	Rent	Wage	Rent/Wage Ratio	Capital-Labor Ratio in Manufacturing	TFP Ratio in Manufacturing 1/
France	10,106	16,667	0.61	71,682	0.84
West Germany	11,890	17,012	0.70	55,410	0.86
Italy	8,940	12,880	0.69	67,535	0.66
United Kingdom	10,011	12,463	0.80	56,497	0.63
Japan	11,142	13,990	0.80	49,357	0.73
USA	8,720	22,405	0.39	84,830	1

1/ The TFP ratio represents the ratio of TFP in a country relative to the US TFP.

Table 2. Basic Regressions and Variations of the Basic Regressions

	Basic Regressions		Special Models		Nontraded Goods	
	1	2	3	4	5	6
Relative Size (f^j_{it})	0.96 *	0.92 *	1.01 *	0.99 *	0.96 *	1.00 *
	(0.07)	(0.07)	(0.06)	(0.08)	(0.07)	(0.07)
Relative Productivity (a^j_{it})	1.64 *	2.57 *	1.19	1.57 *	1.54 *	1.53 *
	(0.28)	(0.32)	(2.00)	(0.29)	(0.28)	(0.28)
Relative Cost (c^j_{it})	-2.13 *	-2.57 *	-2.30 *	-0.34	-2.06 *	-2.04 *
	(0.58)	(0.59)	(0.73)	(0.67)	(0.51)	(0.51)
Endogenous Productivity	No	Yes	No	No	No	Yes
R^2 for Individual Market Regressions						
EU	0.698	0.694	0.676	0.663	0.701	0.698
Other Western Europe	0.652	0.648	0.623	0.623	0.656	0.649
East Asia	0.628	0.622	0.603	0.609	0.634	0.628
France	0.801	0.794	0.778	0.786	0.802	0.799
W. Germany	0.671	0.664	0.652	0.645	0.672	0.670
UK	0.692	0.683	0.650	0.674	0.692	0.686
Japan	0.419	0.432	0.450	0.409	0.414	0.420
Italy	0.749	0.747	0.741	0.743	0.749	0.750

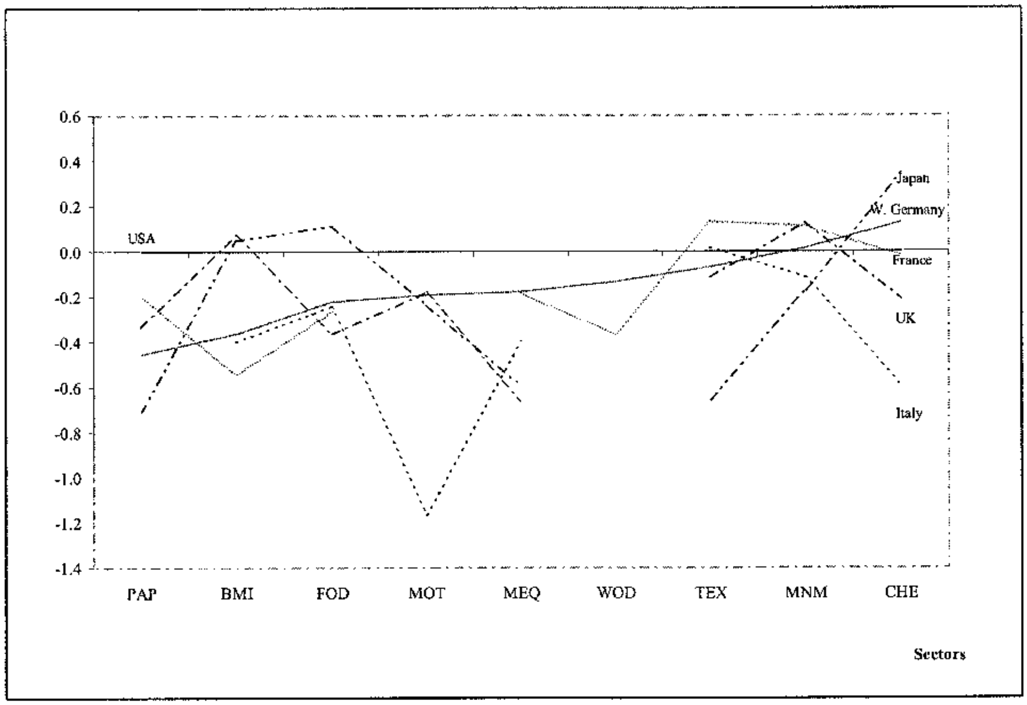
Note: The dependent variable is an index of relative exports (x^{jm}_{it}). The regression equations are estimated as a system using three stage least squares with cross-market constraints. This procedure uses instrumental variables and allows for heteroskedasticty and correlation between residuals across markets. See text for further explanation. Standard errors are shown in parentheses. * denotes significance at the 5% level. None of the coefficient estimates of f^j_{it} are significantly different from one at the 5% level. Also, the coefficients of a^j_{it} and c^j_{it} are not significantly different from each other in absolute value at the 5% level in all cases.

Table 3. Basic Regression With Unrestricted Coefficients Across Markets

	EU	Other Western Europe	East Asia	France	W. Germany	Italy	UK	Japan	Chi-square (percent)
Relative Size (f^j_{it})	0.78 *	0.44 *a	0.48 *a	0.82 *	1.09 *	1.49 *a	0.83 *a	1.73 *a	85.5 *
	(0.11)	(0.12)	(0.13)	(0.11)	(0.11)	(0.12)	(0.09)	(0.24)	
Relative Productivity (a^j_{it})	2.62 *b	2.99 *	1.82 *	2.37 *	2.08 *	1.92 *	2.26 *	3.59 *	8.77
	(0.54)	(0.55)	(0.64)	(0.55)	(0.48)	(0.64)	(0.43)	(0.81)	
Relative Cost (c^j_{it})	-4.95 *b	-4.1 *	-3.69 *	-1.81 *	-2.52 *	-0.85	-1.58	-4.13 *	16.98 *
	(0.93)	(0.94)	(1.13)	(0.92)	(0.95)	(1.02)	(0.94)	(1.44)	
<u>Fixed Effects</u>									
France	2.14 *	1.75 *	-1.23 *	na	2.91 *	3.5 *	1.7 *	-0.09	446.48 *
	(0.20)	(0.20)	(0.24)		(0.20)	(0.23)	(0.16)	(0.39)	
W. Germany	2.63 *	2.92 *	-0.77 *	3.19 *	na	3.26 *	1.81 *	0.14	612.82 *
	(0.17)	(0.17)	(0.21)	(0.16)		(0.18)	(0.13)	(0.31)	
Italy	0.92 *	1.37 *	-1.75 *	2.79 *	2.62 *	na	1.4 *	-0.45	263.5 *
	(0.23)	(0.23)	(0.27)	(0.22)	(0.22)		(0.19)	(0.40)	
UK	1.45 *	1.29 *	-0.9 *	1.75 *	1.48 *	2.15 *	na	0.1	159.82 *
	(0.23)	(0.23)	(0.27)	(0.22)	(0.22)	0.24		(0.41)	
Japan	-0.95 *	-0.57 *	0.86 *	-0.59 *	-0.17	-0.48 *	-0.39 *	na	155.39 *
	(0.17)	(0.17)	(0.20)	(0.17)	(0.17)	0.19	(0.14)		
Period 1	-0.88 *	-0.82 *	-0.46 *	-0.92 *	-0.75 *	-0.75 *	-1.11 *	-0.66 *	9.54
	(0.17)	(0.17)	(0.22)	(0.18)	(0.19)	(0.19)	(0.17)	(0.26)	
Period 2	-0.36 *	-0.35 *	-0.19	-0.59 *	-0.37 *	-0.47 *	-0.58 *	-0.25	7.31
	(0.14)	(0.14)	(0.17)	(0.14)	(0.14)	(0.16)	(0.12)	(0.21)	
Period 3	-0.73 *	-0.64 *	-0.51 *	-0.52 *	-0.53 *	-0.4 *	-0.58 *	-0.67 *	4.9
	(0.15)	(0.15)	(0.18)	(0.15)	(0.15)	(0.16)	(0.14)	(0.22)	
No. of Observations	160	160	160	128	124	132	128	128	
R-squared	0.712	0.687	0.664	0.802	0.661	0.736	0.700	0.510	

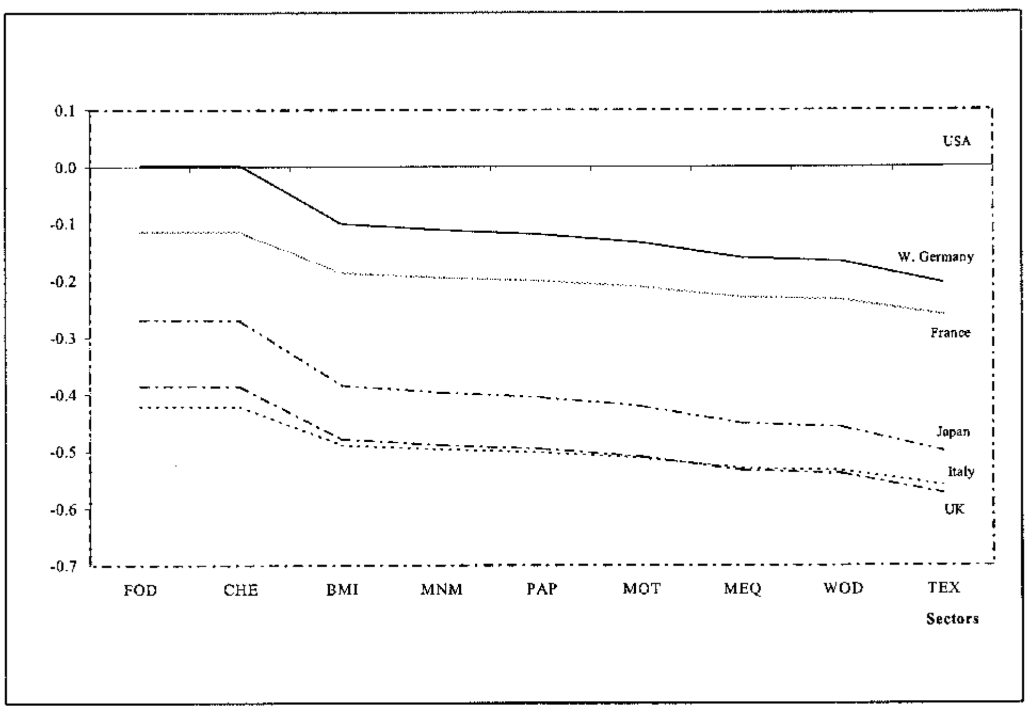
Note: The dependent variable is an index of relative exports (x^{jm}_{it}). The regression equations are estimated as a system using three-stage least squares which allows for heteroskedasticity and correlation between residuals across markets, and uses instrumental variables for f^j_{it} and a^j_{it} (see text for further explanation). Standard errors are shown in parentheses. The chi-square statistics represent the Wald test for the restriction that the coefficient is the same across markets. * denotes significance at the 5% level. a indicates that the coefficient of f^j_{it} is significantly different from one at the 5% level while b indicates that coefficients of a^j_{it} and c^j_{it} are significantly different from each other in absolute value at the 5% level.

Figure 1. TFP Indexes Across Sectors



Note: Gaps represent missing data for TFP. See text for further explanation. Also, see Appendix Table for definitions of sectoral codes.

Figure 2. Composite Factor Cost Indexes Across Sectors



Note: See text for further explanation. Also, see Appendix Table for definitions of sectoral codes.

Data Appendix

This Appendix provides the definition and data sources for each variable used in the study. Unless indicated otherwise, the data are obtained from the OECD's International Sectoral Database (ISDB), and ISDB variable codes are given when this source is used.

1. X_{it}^{jm} is the value of exports of sector *i*'s products from country *j* to market *m* in thousands of US$. Source: OECD's Bilateral trade database (1998 edition).

2. Y_{it}^{j} is value added at 1990 prices and 1990 purchasing power parities (PPPs, in US dollars). ISDB variable code: GDPD.

3. K_{it}^{j} is the gross capital stock at 1990 prices and 1990 PPPs (US dollars). ISDB variable code: KTVD.

4. L_{it}^{j} is total employment. ISDB variable code: ET.

5. $\tilde{\theta}_{i}^{L}$ is the share of labor in value added for each sector, *i*. This was calculated as follows (using ISDB variable codes):
   $\tilde{\theta}_{it}^{L} = (WSSS * ET / EE) / GDP$
   where,
   WSSS represents the compensation of employees in current prices and national currency,
   ET represents total employed,
   EE is the number of employees, and
   GDP is value added in current prices and national currency.
   Total compensation is re-scaled by the ratio of total employment to total employees in order to include self-employed in the weighting scheme. $\tilde{\theta}_{i}^{L}$ represents the average value of $\tilde{\theta}_{it}^{L}$ over the 1970–1990 period and the six countries in the sample.

6. $\tilde{\theta}_{i}^{K}$ is the share of capital in value added for each sector, *i*. This was calculated as:

   $$\tilde{\theta}_{i}^{K} = (1 - \tilde{\theta}_{i}^{L}).$$

7. θ_{i}^{Y} is the ratio of value added in current prices to gross output in current prices for each sector, *i* (for each sector, six-country averages of the ratios over the 1970–1990 period were used). Source: OECD's STAN Database for Industrial Analysis, 1998.

8. W_{t}^{j} is the manufacturing wage rate in country *j* at time *t* expressed in US dollars. This was calculated from the ISDB as: $W_{t}^{j} = (WSSS / EE) / ER$
   where,
   ER is the exchange rate defined as national currency per US dollar (from the IMF's IFS).

9. R_t^j is the rental rate in country *j* at time *t* expressed in US dollars. This was constructed using the method originated by Hall and Jorgenson (1967) as:
$R_t^j = (r + \delta) P_I$
where *r* is the real interest rate, δ is the depreciation rate, and P_I is the price of the investment good. The real interest rate was calculated as the government bond yield minus the percentage change in the CPI index (both from various issues of the International Monetary Fund's *International Financial Statistics*) as in Caballero and Lyons (1990). Also, like Caballero and Lyons, a depreciation rate of 10% was used to calculate the rental price of capital. P_I is defined as the US dollar price of a comparable investment good (i.e., worth one US dollar in 1990 prices) and is measured as follows:
$$P_I = \frac{IT * PPP(GFCF)}{ITV * ER}$$
where
IT represents gross fixed capital formation in current prices and national currency, ITV represents gross fixed capital formation in 1990 prices and national currency, and PPP (GFCF) represents the purchasing power parity for investment goods (in national currencies per US dollar).
Since data on IT was not available for Japan, a proxy for P_I for Japan was obtained as follows:
$$P_I = \frac{GDP * PPP(GFCF)}{GDPV * ER}$$
where,
GDPV represents value added in 1990 prices and national currency.

10. The nontraded goods sectors are assumed to consist of: electricity, gas and water, construction, transport, storage and communication, community, social and personal services, and producers of government services. The TFP and cost ratios of the non-traded sector are calculated using the same methodology as for manufacturing sectors. θ_i^N is the average share [across (available) time periods and countries] of the nontraded sector in sector *i*'s gross output (Source: OECD's Input Output Database).

11. PS_t^j is population size in millions in country *j* at time *t*. Source: IMF's *International Financial Statistics*.

12. RD_t^j is business enterprise expenditure on R&D activities for total manufacturing in US dollar PPPs. This measure is not available for 1970–1972. Therefore, the 1973–1974 average is used for the first subperiod. Source: OECD's Anberd database

Appendix Table. Sector Codes

ISIC codes	ISDB codes	Industry
3	MAN	Manufacturing
31	FOD	Food, beverages and tobacco
32	TEX	Textiles, wearing apparel and leather industries
33	WOD	Wood and wood products, including furniture
34	PAP	Paper and paper products, printing and publishing
35	CHE	Chemicals and chemical petroleum, coal, rubber and plastic products
36	MNM	Non-metallic mineral products except products of petroleum and coal
37	BMI	Basic metal industries
38	MEQ	Fabricated metal products, machinery and equipment
39	MOT	Other manufacturing industries

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