U.S. Exports and Multinational Production

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Abstract

This paper presents a monopolistic competition model of trade and multinational production that incorporates asymmetric trade barriers and international differences in production costs. The model predicts the functional form for the dependence of U.S. exports and multinational production on tariffs, distance and production costs. To deal with the issue of simultaneity, we estimate the non-linear equations of U.S. exports and multinational production simultaneously. The estimation yields reasonable estimates of the structural parameters including the elasticity of substitution. The estimates suggest that the elimination of tariffs would increase U.S. exports by 20.7 percent, but reduce U.S. multinational production by 3.7 percent. (JEL classification: F1, Keywords: monopolistic competition, multinational production, trade barriers, elasticity of substitution)

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1 Introduction

Trade and multinational production are inseparable components of international commercial activities. New trade theory, as represented by the monopolistic competition model, has been developed to explain intra-industry trade among countries (Krugman 1980, Helpman 1981, Helpman and Krugman 1985). It has also been expanded to explain the rapid growth of multinational production over the last two decades (Helpman 1984, Markusen 1984, Brainard 1993).

In the empirical literature it is traditional to debate whether trade and multinational production are complements or substitutes. The traditional approach is to postulate a regression of trade on foreign direct investment (FDI). A positive coefficient is used as evidence of complementarity, while a negative coefficient is used as evidence of substitutability. Alternatively, FDI is regressed on trade and similar inferences are drawn. In recognition of the simultaneity between trade flows and multinational sales, Brainard (1997) uses the share of total trade accounted for by multinational sales as the dependent variable.

In this paper we incorporate trade and multinational production into a unified monopolistic competition framework (Krugman 1980, Helpman and Krugman 1985). We also allow for a rich set of international asymmetries, including asymmetric trade barriers and international differences in production costs. Tariffs, distance, and production costs enter firms' pricing and output decisions. These decisions, when set against the backdrop of CES preferences, yield precise predictions for the dependence of U.S. exports and multinational production on tariffs, distance and production costs.

In the empirical analysis, we take a radically different approach to dealing with the simultaneity between trade and FDI. Specifically, we estimate the non-linear equations of U.S. exports and multinational production simultaneously, recognizing the cross-equation restrictions on parameters and the correlation between residuals. The two equations are linked in two ways. First, both exports and multinational production depend on trade costs and production costs, and share common structural parameters (e.g., the elasticity of substitution). Second, our augmented monopolistic competition model may not capture all the determinants of trade and FDI (e.g., fixed costs of exporting, fixed costs of investment abroad, economies of scales at the plant or firm level). The omitted variables can affect both exports and multinational production. The existence of common omitted variables implies that the residuals in the two equations will be correlated.

Our empirical study draws data from various sources. The estimation yields reasonable estimates of the structural parameters. In particular, the baseline estimate of the elasticity of substitution σ is 4.72, which is slightly smaller than other estimates obtained in the literature. For example, the point estimate of σ is 6.43 in Baier and Bergstrand (2001). The smaller estimate of σ is also well in line with the estimates obtained from micro studies of demand for differentiated products (e.g., Feenstra 1994). Furthermore, our estimation reveals that the effects of excluded variables affecting U.S. exports have significantly negative correlation with those of excluded factors affecting U.S. multinational production. This is consistent with the hypothesis of the proximity-concentration trade-off between trade and FDI as examined in Brainard (1997).

The estimation of the structural parameters also allows us to evaluate the relative effects of trade policy on trade and multinational production. We find that the elimination of tariffs would substantially increase U.S. exports (20.7 percent on average with substantial country variation), but would have only a very small negative effect on U.S. production abroad (3.7 percent on average).

This paper is closely related to the large empirical literature on trade and FDI. Many papers have studied the implications of the monopolistic competition model for the relationship between trade costs and bilateral trade (e.g., Bergstrand 1985, Bairer and Bergstrand 2001, Hummels and Levinsohn 1995, Anderson and van Wincoop 2003).

Built on their insight about the importance of distance and tariffs, our model further incorporates production costs. Empirically, we identify the elasticity of substitution in consumption between varieties, the distance effect, and the production-cost effect in a highly non-linear model. This is a significant contribution to the literature since previous studies have placed zero restrictions on at least some of these structural parameters.

This paper also differs from the existing literature on FDI in several major aspects. First, most studies focus on two major reasons for multinationals to produce abroad. One is to take advantage of low production costs abroad and serve the domestic market, i.e., vertical integration (Helpman 1984, Helpman and Krugman 1985). The other is to jump over trade barriers and serve the host country, i.e., horizontal integration (Markusen 1984, Markusen and Venables 2000). However, using data on U.S. multinationals from the Bureau of Economic Analysis (BEA), Hanson, Mataloni and Slaughter (2001) find that in many cases U.S. foreign affiliates serve as export platforms, i.e., they produce in the host country and sell their products to third markets (besides the host country and the United States). Motivated by this finding, in this paper we assume that goods produced by multinationals will be sold to all markets. That is, we do not treat vertical and horizontal integration as mutually exclusive options. Our assumption is also consistent with the love-for-variety CES preferences in the standard monopolistic competition model. The model predicts that the host country's ability to access other markets plays a key role in determining the output level for multinationals. This theoretical prediction is strongly supported by our empirical findings.

Second, previous empirical studies on FDI use reduced-form specifications and yield no structural parameters. In contrast, our estimation relies on the exact functional form as predicted by the theory. Estimating structural parameters is one of our main goals. Third, recent empirical papers (e.g., Brainard 1997, Helpman, Melitz and Yeaple 2003) focus on shares rather than levels of FDI. In contrast, we are concerned with levels of

U.S. multinational production. The estimation of structural parameters and the use of level specifications allow us to assess the implications of trade liberalization for U.S. multinational production. This has not been done in previous studies.

Finally, we emphasize that our paper is *not* about the location decisions of multinationals. Instead, in order to make our current problem tractable, we take the location decisions of multinationals as given. That is, we focus on the trade and production decisions of *existing* plants. However, by taking account of the correlation between the residuals in the equations of U.S. exports and multinational production, we partially control for the effects of excluded variables determining the choice between exports and FDI.

This paper is organized as follows. Section 2 develops a simple monopolistic competition model of trade and multinational production. Section 3 presents the empirical specification and the sample. Sections 4-7 illustrate empirical results, including linear and non-linear estimates, implications of trade liberalization, and additional alternative specifications. Section 8 concludes the paper.

2 Theory

Trade and multinational production are inseparable components of international commercial activities. For example, in 1992, about one-third of U.S. exports and over 40 percent of U.S. imports consisted of intra-firm trade between U.S. firms and their foreign affiliates or between foreign multinationals and their U.S. affiliates (see Mataloni 1995).

In this section we present a simple monopolistic competition model of international trade and multinational production. Since we are interested in short-run output decisions by firms, we take location choices as given. That is, our analysis does not deal with the choice between exports and FDI.[1] Instead, we focus on trade flows and multinational

[1] The choice between exports and FDI has been considered by many authors, e.g., Helpman (1984),

production after location choices have been made. Furthermore, we assume that products are differentiated by locations. In particular, products that are made in different countries by plants of the same multinational are treated as different varieties. This may be due to consumer perceptions and quality control standards of different countries.

2.1 Set-up

The set-up closely follows the standard monopolistic competition model (e.g., Krugman 1980, Krugman and Helpman 1985). However, we extend the standard model to many countries since the multilateral relationship among countries plays an essential role in our analysis.

Consumers have identical Cobb-Douglas preferences over goods and CES preferences over varieties. With Cobb-Douglas preferences we can look at one good at a time. So we fix the good and suppress the goods index. Let k index varieties. Let i and j index user countries and producer countries, respectively. The total number of countries is N. In the first stage a representative consumer in country i allocates Y_i to the differentiated good in question. In the second stage the representative consumer maximizes the CES subutility function subject to the expenditure constraint:

$$\max_{\{q_{ij}^k\}} U_i = \left[\sum_{j=1}^{N} \int_0^{n_j} \left(q_{ij}^k\right)^{\frac{\sigma-1}{\sigma}} dk\right]^{\frac{\sigma}{\sigma-1}}$$

$$s.t. \sum_{j=1}^{N} \int_0^{n_j} p_{ij}^k q_{ij}^k dk = Y_i$$

where n_j is the number of varieties (or firms) in country j, q_{ij}^k is country i's demand for variety k produced in country j, p_{ij}^k is the consumer price associated with q_{ij}^k, and σ is the elasticity of substitution between varieties ($\sigma > 1$). In equilibrium σ is also the elasticity

Markusen (1984), Brainard (1993, 1997), and Helpman, Melitz and Yeaple (2003).

of demand for varieties. Then country i's demand for variety k produced in country j is

$$q_{ij}^{k} = \frac{\left(p_{ij}^{k}\right)^{-\sigma}}{\sum_{j'=1}^{N} \int_{0}^{n_{j'}} \left(p_{ij'}^{k'}\right)^{1-\sigma} dk'} Y_i. \tag{1}$$

In the following we will add a rich set of international asymmetries to the standard monopolistic competition model. First, firms located in different countries incur different costs of production. We take firms' location choices as given. Once a firm is set up, the firm's fixed costs are irrelevant to price and quantity decisions.[2] So these decisions will only be affected by marginal costs. Let c_j be the marginal cost for firms in country j. Second, we allow for tariffs. Let τ_{ij} be 1 plus the *ad valorem* tariff rate imposed by country i on varieties imported from country j. Finally, we take into account distance-related transportation costs. Let d_{ij} be 1 plus the per unit transportation cost that increases in the bilateral distance between countries i and j. It means that for each unit of output reaching country i, $d_{ij} - 1$ units are lost in the transaction process.

Trade barriers drive a wedge between the price paid by consumers and the price received by producers. Let p_j^k be the price charged by the producers of variety k in country j. Given that the elasticity of demand for varieties is σ, the producer price is set as $p_j^k = c_j \sigma / (\sigma - 1)$. Then consumers in country i will face the price

$$p_{ij}^{k} = d_{ij} \tau_{ij} p_j^k = \frac{\sigma}{\sigma - 1} d_{ij} \tau_{ij} c_j. \tag{2}$$

Substituting this price into equation (1) gives the quantity demanded:

$$q_{ij}^{k} = \frac{\left(d_{ij} \tau_{ij} c_j\right)^{-\sigma}}{\sum_{j'=1}^{N} n_{j'} \left(d_{ij'} \tau_{ij'} c_{j'}\right)^{1-\sigma}} \frac{\sigma - 1}{\sigma} Y_i. \tag{3}$$

Equations (2) and (3) show that p_{ij}^k and q_{ij}^k vary across country pairs i and j, but not

[2] Fixed costs are important for the choice between exports and FDI (e.g., Markusen 1984, Brainard 1993, Helpman, Melitz and Yeaple 2003).

varieties k.

In the following we use this simple framework to examine the impact of trade barriers and production costs on U.S. exports and multinational production.

2.2 U.S. Exports

Using the price equation (2) and the quantity-demanded equation (3), the value of U.S. exports to country i can be derived as the sum of country i's demand for varieties produced in the United States ($j = us$):

$$X_i^{us} = \int_0^{n_{us}} \left(p_{i,us}^k q_{i,us}^k\right) dk = n_{us} \left(d_{i,us}\tau_{i,us}c_{us}\right)^{1-\sigma} P_i^{\sigma-1} Y_i \tag{4}$$

where $P_i \equiv \left[\sum_{j=1}^{N} n_j \left(d_{ij}\tau_{ij}c_j\right)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$ is the aggregate CES price index in country i. Note that $P_i^{\sigma-1}Y_i$ measures the effective market size or 'market capacity' of country i (See Redding and Venables 2002). The number of firms (or varieties) in the United States, n_{us}, includes U.S. firms as well as foreign multinationals operating in the United States.[3]

Equation (4) specifies the impact of trade barriers and production costs on U.S. exports. The results can be summarized in theorem 1.

Theorem 1 *(1)* $\partial X_i^{us}/\partial \tau_{i,us} < 0$, $\partial X_i^{us}/\partial d_{i,us} < 0$; *(2)* $\partial X_i^{us}/\partial \tau_{ij} > 0$, $\partial X_i^{us}/\partial d_{ij} > 0$ *for* $j \neq us$; *(3)* $\partial X_i^{us}/\partial c_{us} < 0$, $\partial X_i^{us}/\partial c_j > 0$ *for* $j \neq us$.

Part (1) of theorem 1 suggests that trade barriers between country i and the United States (i.e., $\tau_{i,us}$ and $d_{i,us}$) reduce U.S. exports to country i. Part (2) captures the effects of trade barriers between country i and other trading partners. Both τ_{ij} and d_{ij} affect the effective market size of country i (i.e., $P_i^{\sigma-1}Y_i$) via the aggregate CES price index

[3] Above 10 percent of total U.S. merchandise exports in 1992 were through foreign multinationals operating in the United States.

P_i. *Ceteris paribus*, country i's trade barriers against its non-U.S. trading partners raise its effective market size, thus increasing its imports from the United States. Part (3) captures the impact of production costs. Specifically, either high production costs in the United States or low production costs elsewhere can reduce U.S. exports to country i.

2.3 U.S. Multinational Production

We are especially interested in the short-run output decisions by U.S. multinational firms. Most previous studies focus on two major reasons for multinationals to produce abroad. One is to take advantage of low production costs abroad and serve the U.S. market, i.e., vertical integration (Helpman 1984, Helpman and Krugman 1985). The other is to jump over trade barriers and serve the host country market, i.e., horizontal integration (Markusen 1984). However, Hanson, Mataloni and Slaughter (2001) find that in many cases U.S. foreign affiliates also serve as export platforms, that is, they produce in the host country and sell their products to third markets (besides the host country and the United States).

Inspired by their finding, in this paper we do not treat horizontal and vertical integration as mutually exclusive options. Instead, we combine the two types of integration by assuming that varieties produced by U.S. affiliates in country j will be sold to all markets.[4] This assumption is also consistent with the love-for-variety CES preferences in the standard monopolistic competition model. Specifically, let n_j^{us} be the number of U.S. affiliates operating in country j. Using equations (2) and (3), country i's demand for the varieties produced by U.S. affiliates in country j may be derived as $Q_{ij}^{us} = n_j^{us} \left(d_{ij}\tau_{ij}c_j\right)^{1-\sigma} P_i^{\sigma-1} Y_i$. (Recall that $P_i \equiv \left[\sum_{j'=1}^{N} n_{j'} \left(d_{ij'}\tau_{ij'}c_{j'}\right)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$ is the aggregate CES price index in country i.) Unfortunately, data on Q_{ij}^{us} are unavailable. This forces us to aggregate Q_{ij}^{us} over

[4] It is worth pointing out that the knowledge-capital model nicely combines the vertical and horizontal multinationals in a unified framework (Markusen 2002, Carr, Markusen and Maskus 2001).

i to obtain the total demand for the products made by U.S. affiliates in country j:

$$Q_j^{us} = \sum\nolimits_{i=1}^{N} Q_{ij}^{us} = n_j^{us} c_j^{1-\sigma} \sum\nolimits_{i=1}^{N} \left(d_{ij}\tau_{ij}\right)^{1-\sigma} P_i^{\sigma-1} Y_i. \tag{5}$$

Data on Q_j^{us} are available from the BEA.

Equation (5) shows that two major factors are at play. On the one hand, U.S. multinational production depends on the host country's marginal cost c_j. Specifically, $\partial Q_j^{us}/\partial c_j < 0$, i.e., U.S. multinationals will produce more in a host country where the marginal cost is lower. The production cost differences may arise from endowment differences among countries as considered in Helpman (1984).

On the other hand, U.S. multinational production also hinges on a host country's 'market access' which is captured by $\sum_{i=1}^{N} \left(d_{ij}\tau_{ij}\right)^{1-\sigma} P_i^{\sigma-1} Y_i$.[5] This can be decomposed as $P_j^{\sigma-1} Y_j + \sum_{i=1, i\neq j}^{N} \left(d_{ij}\tau_{ij}\right)^{1-\sigma} P_i^{\sigma-1} Y_i$. The first term is the effective market size of host country j. The second term measures host country j's ability to access other markets, including the United States. In particular, if host country j faces low tariff barriers to access large markets (i.e., low τ_{ij} and big $P_i^{\sigma-1} Y_i$), or country j is close to large markets (i.e., small d_{ij} and big $P_i^{\sigma-1} Y_i$), host country j has better access to other markets. From equation (5), it follows that U.S. multinational production will be higher in a host country with a larger market size of its own and/or better access to other big markets.

Therefore, in addition to factor prices, market access plays a key role in determining short-run output for multinational firms. Since market access depends on trade barriers and marginal costs, we summarize the effects of those ultimate factors on multinational production in theorem 2.

Theorem 2 *For $i \neq j$, (1) $\partial Q_j^{us}/\partial \tau_{ij} < 0$, $\partial Q_j^{us}/\partial \tau_{ji} > 0$; (2) $\partial Q_j^{us}/\partial c_j < 0$, $\partial Q_j^{us}/\partial c_i > 0$.*

[5] This definition of market access follows Redding and Venables (2002). However, in addition to transport costs, we also incorporate bilateral tariffs.

Part (1) of theorem 2 suggests that τ_{ij} and τ_{ji} have asymmetric effects on Q_j^{us}. In particular, a smaller τ_{ij} implies that country j has better access to country i's market. This provides an incentive for U.S. multinationals to expand production in country j and thus export more to country i. On the other hand, if country j erects high tariff barriers against other countries (including the United States), the effective market size of country j (i.e., $P_j^{\sigma-1}Y_j$) becomes larger. (This effect of tariff barriers on market size works through the aggregate CES price index P_j.) U.S. multinationals will produce more in country j in order to serve the local market there.[6] This represents the tariff-jumping motivation for multinationals. However, differing from the traditional tariff-jumping argument, country j's tariffs against non-U.S. trading partners also matter in our model. This is because our model emphasizes the multilateral relationship between countries by taking into account the spillover effects among them.

Part (2) of theorem 2 implies that if production costs are low in country j or high in other countries, U.S. multinationals will produce more in the low-cost country j. Cost differences have often been considered as a driving force behind vertical integration (Helpman 1984).

Finally, unlike tariffs, distance has an ambiguous effect on Q_j^{us}. This difference arises from the fact that d_{ij} (or d_{ji}) has opposing effects on country j's market access. On the one hand, a smaller d_{ij} (or d_{ji}) improves country j's access to other markets, i.e., $\sum_{i=1,i\neq j}^{N} (d_{ij}\tau_{ij})^{1-\sigma} P_i^{\sigma-1}Y_i$. On the other hand, a smaller d_{ij} (or d_{ji}) reduces the effective market size of country j, i.e., $P_j^{\sigma-1}Y_j$. Therefore, if the former effect dominates, a smaller d_{ij} (or d_{ji}) is likely to be related to a larger Q_j^{us}.

To summarize, our simple model yields precise predictions for the dependence of U.S. exports and multinational production on trade barriers and production costs. In the

[6] These implications of theorem 2 are largely consistent with the observation by Hanson, Mataloni, and Slaughter (2001) that "in larger, more-protected, and higher-tax economies, affiliates target most of their sales to the domestic market. But in smaller, less-protected, and lower-tax economies, U.S. multinationals set up export platforms that devote more of their sales to export markets in nearby regions and beyond."

following we turn to a detailed empirical analysis of theorems 1-2 and equations (4)-(5).

3 Empirical Specification and Data

In the empirical literature it is traditional to argue about whether trade and FDI are substitutes or complements. The usual method for addressing this begins with a regression of trade on FDI. A positive FDI coefficient is seen as evidence of complementarity between trade and FDI. A negative coefficient is seen as evidence of substitutability between trade and FDI. Alternatively, FDI is regressed on trade and similar conclusions are drawn. To deal with the problem of simultaneity between trade flows and multinational sales, Brainard (1997) uses the share of total trade accounted for by multinational sales as the dependent variable.

In this paper we take a radically different approach. Specifically, we use equations (4) and (5) to form the framework for evaluating the connection between trade and multinational production. We parameterize the model as follows. Following Brainard (1993), we assume $d_{ij} = e^{\mu D_{ij}}$, where D_{ij} is the distance between country i and country j. The parameter μ is interpreted as the trade cost per unit of distance. To measure marginal costs, we assume $c_j = \omega_j^{\alpha}$, where ω_j is an international comparable price index of inputs (see the data appendix for more details). The parameter α captures the effect of production costs. Then substituting d_{ij} and c_j into equations (4) and (5), taking logarithms, adding time subscripts and error terms, we have

$$\log X_{jt}^{us} = \log\left(\frac{n_{us,t}\left[e^{\mu D_{j,us}}\tau_{j,us,t}\omega_{us,t}^{\alpha}\right]^{1-\sigma}}{\sum_{j'=1}^{N} n_{j't}\left[e^{\mu D_{jj'}}\tau_{jj't}\omega_{j't}^{\alpha}\right]^{1-\sigma}}Y_{jt}\right) + \varepsilon_{jt} \tag{6}$$

$$\log Q_{jt}^{us} = \log\left(\sum_{i=1}^{N}\frac{n_{jt}^{us}\left[e^{\mu D_{ij}}\tau_{ijt}\omega_{jt}^{\alpha}\right]^{1-\sigma}}{\sum_{j'=1}^{N} n_{j't}\left[e^{\mu D_{ij'}}\tau_{ij't}\omega_{j't}^{\alpha}\right]^{1-\sigma}}Y_{it}\right) + \upsilon_{jt} \tag{7}$$

where ε_{jt} and υ_{jt} include all unobserved factors that may determine X_{jt}^{us} and Q_{jt}^{us}, respectively.

Equations (6) and (7) are linked in two ways. First, they share a common set of parameters σ, μ and α. These parameters can be estimated by relating the variation in patterns of trade and multinational production to variation in tariffs, distance, and marginal costs. Second, our augmented monopolistic competition model does not completely incorporate all the determinants of trade and multinational production. For example, it excludes fixed costs of exporting, fixed costs of investment abroad,[7] economies of scales, language and adjacency dummies, unmeasured non-tariff barriers, political risk and other omitted determinants of trade and FDI. These omitted determinants are arguably common to both equations, and implicitly contribute to the residuals ε_{jt} and υ_{jt}. The existence of common omitted variables implies that the two residuals will be correlated. We assume the structure of the error terms is

$$\begin{bmatrix} \varepsilon_{jt} \\ \upsilon_{jt} \end{bmatrix} \sim N \left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \delta_{\varepsilon}^2 & \rho\delta_{\varepsilon}\delta_{\upsilon} \\ \rho\delta_{\varepsilon}\delta_{\upsilon} & \delta_{\upsilon}^2 \end{bmatrix} \right)$$

where δ_{ε}^2 is the variance of ε_{jt}, δ_{υ}^2 is the variance of υ_{jt}, and ρ measures the correlation of the residuals ε_{jt} and υ_{jt}.

Recognizing the cross-equation restrictions on parameters and the correlation of error terms, we estimate equations (6)-(7) simultaneously using the maximum-likelihood method. The joint estimation has two advantages. First, it improves efficiency compared to single-equation estimations. Second, it allows us to test joint hypotheses involving parameters in both equations. The estimation will yield structural parameters σ, μ and α. These parameters can be used to evaluate the impact of trade liberalization on U.S. exports and multinational production. In contrast, previous studies of FDI yield no structural parameters and thus cannot assess the effects of trade liberalization on FDI. Furthermore, even though our specification does not explicitly include measures of fixed

[7] In Helpman, Melitz and Yeaple (2003), fixed costs of exporting and fixed costs of investment abroad are captured by country-specific fixed effects.

costs or economies of scales, the parameter ρ can partially imply the effects of those omitted variables on U.S. exports and multinational production. In particular, a negative ρ indicates that the effects of the omitted variables affecting U.S. exports are negatively correlated with those of the omitted variables affecting multinational production.

This empirical analysis draws data from various sources. See the data appendix for more details about data sources and measurement. The sample covers 30 U.S. trading partners. The selection of countries is dictated by the availability of data on bilateral tariffs, the number of manufacturing plants, and U.S. multinational production. The countries include most of the OECD, many of the developing countries intensively engaged in trade, and some developing countries that are not heavy traders. The years used are 1984, 1986, 1988, 1990, and 1992. The availability of multinational production data and tariff data allows us to focus only on this time period.

The analysis is done at the level of aggregate manufacturing. There are two reasons. First, the results in Lai and Trefler (2002) suggest that the σ for aggregate manufacturing is an aggregate of the σ from individual manufacturing industries. This means that the analysis of aggregate manufacturing is sufficient to explore the key economic interpretations of our model. Second, estimation at the aggregate level can better illustrate the different roles of tariffs, distance and production costs in determining trade and multinational production.

4 Preliminary Estimates

To examine theorems 1 and 2 empirically, we start with simple linear regressions. Equations (6) and (7) can be linearized as

$$\log X_{jt}^{us} = \beta_1 \log \tau_{j,us,t} + \beta_2 D_{j,us} + \beta_3 \log Y_{jt} + \beta_4 \log n_{us,t} + \varepsilon'_{jt} \quad (8)$$

$$\log Q_{jt}^{us} = \gamma_1 \log \tau_{j,us,t} + \gamma_2 D_{j,us} + \gamma_3 \sum_{i=1}^{N} Y_{it} \log \tau_{ijt} + \gamma_4 \sum_{i=1}^{N} Y_{it} D_{ij} + \gamma_5 \log \omega_{jt} + \gamma_6 \log n_{jt}^{us} + \upsilon'_{jt} \quad (9)$$

where the residual ε'_{jt} absorbs the effect of τ_{jit}, D_{ji}, and ω_{it} ($i \neq j$, $i \neq us$) on X_{jt}^{us},[8] and the residual υ'_{jt} absorbs the effect of τ_{jit} and ω_{it} ($i \neq j$, $i \neq us$) on Q_{jt}^{us}. Using theorems 1 and 2, most coefficients in equations (8) and (9) can be signed. Unfortunately, the structural parameters σ, α and μ cannot be retrieved from these linear regressions.

Consider the U.S. exports equation (8). As implied by part (1) of theorem 1, both β_1 and β_2 should be negative.[9] That is, bilateral tariffs ($\tau_{j,us,t}$) and distance ($D_{j,us}$) should reduce U.S. exports to country j. The individual demand of the importer ($\log Y_{jt}$) and the industrial supply of the exporter ($\log n_{us,t}$) control for the effect of country size. We expect that β_3 and β_4 are positive. The OLS estimate of equation (8) is reported in column 1 of Table 1. The estimates of all coefficients have expected signs. In particular, $\log \tau_{j,us,t}$ has a significantly negative effect on U.S. exports to country j.

Now consider the U.S. multinational production equation (9). As implied by part (1) of theorem 2, $\tau_{j,us,t}$ increases Q_{jt}^{us}, which reflects the tariff-jumping effect. So γ_1 is positive. On the other hand, τ_{ijt} reduces Q_{jt}^{us}, implying that $\sum_{i=1}^{N} Y_{it} \log \tau_{ijt}$ is negatively correlated with Q_{jt}^{us}. That is, γ_3 should be negative. Part (2) of theorem 2 suggests that Q_{jt}^{us} will

[8] Note that our price index of inputs uses the U.S. price index as the numeraire (see the data appendix). That is, $\omega_{us,t} \equiv 1$. Thus, the term of $\log(\omega_{us,t})$ vanishes from equation (8).

[9] Note that $d_{j,us}$ increases in $D_{j,us}$.

be higher in a host country with lower ω_{jt}.[10] Thus, γ_5 should be negative. However, the theory has no clear-cut prediction for the distance effect. As mentioned in section 2.3, a smaller D_{ij} (which implies a smaller d_{ij}) increases country j's ability to access other markets but reduces its own effective market size. If the former effect dominates, D_{ij} and Q_{jt}^{us} should be negatively correlated, implying that γ_2 and γ_4 are negative. Otherwise, γ_2 and γ_4 may be positive. Finally, equation (9) also controls for $\log n_{jt}^{us}$, the production capacity of U.S. affiliates in host country j. We expect γ_6 to be positive.

Columns 2-3 of Table 1 display the OLS estimates of the U.S. multinational production equation (9). In column 2, $\log \tau_{j,us,t}$ and $D_{j,us}$ are excluded. The estimated coefficient on $\sum_{i=1}^{N} Y_{it} \log \tau_{ijt}$ is -7.25 ($t = -2.37$) in column 2 and -8.63 ($t = -2.71$) in column 3. These results confirm our theoretical prediction. The estimate of the coefficient on $\sum_{i=1}^{N} Y_{it} D_{ij}$ is also negative. In particular, the estimated γ_4 is -0.07 ($t = -2.05$) in column 3. This suggests that with a smaller distance, the increase in country j's access to other markets likely dominates the decrease in its own effective market size. Since both $\sum_{i=1}^{N} Y_{it} \log \tau_{ijt}$ and $\sum_{i=1}^{N} Y_{it} D_{ij}$ are related to country j's ability to access other markets, these results make it clear that the host country's market access matters.

At the same time, the estimated coefficient on $\log \omega_{jt}$ is negative, indicating that U.S. multinational production will be higher in low-cost countries. However, this negative production-cost effect is not statistically significant. This is consistent with Brainard (1997) who emphasizes trade barriers rather than the factor price difference as a determinant of FDI. Furthermore, in column 3 the estimated coefficients on $\log \tau_{j,us,t}$ and $D_{j,us}$ are weakly positive, which provides some evidence of the trade-barrier-jumping effect on multinational production. Interestingly, the coefficient on $\log n_{jt}^{us}$ is estimated to be greater than one, indicating that the scale of production (i.e., Q_{jt}^{us}/n_{jt}^{us}) is larger in a host country which has more U.S. affiliates operating there.

[10] Note that c_{jt} increases in ω_{jt}.

Therefore, the estimates in columns 2-3 reveal that U.S. multinational production is influenced by various factors, including cost differentials, the trade-barrier-jumping effect, and the host country's market access in particular.

Finally, we estimate equations (8) and (9) jointly using seemingly unrelated regressions (SUR) analysis. The residuals ε'_{jt} and υ'_{jt} are likely to be correlated due to some common omitted variables, e.g., τ_{ijt}, ω_{it}, fixed costs, and economies of scales. As shown in columns 4-5, the estimates are fairly similar to those in columns 1 and 3. Notably, the estimated correlation of the two residuals is -0.22. This negative correlation suggests that the omitted variables have opposing effects on trade and multinational production. This is also in accord with the hypothesis of the proximity-concentration trade-off between multinational sales and trade as examined in Brainard (1997).

To summarize, the estimates from linear regressions are largely consistent with our theoretical predictions. On the other hand, linear regressions fail to deliver estimates of the structural parameters σ, μ and α. Thus, in the next section we turn to non-linear estimation of equations (6) and (7).

5 Maximum-Likelihood Estimates

We estimate equations (6) and (7) simultaneously using the maximum-likelihood method. The results are presented in Table 2. In all specifications we impose the cross-equation constraints that the three structural parameters σ, μ, and α are equal in the two equations. Column 1 reports the baseline estimates. The estimated elasticity of substitution $\widehat{\sigma}$ is 4.72 ($t = 4.46$). This estimate is somewhat smaller than the basic model estimate of 6.42 in Lai and Zhu (2003), but slightly larger than their fixed-effects estimate of 3.99. This difference mainly arises from the fact that the two studies use different samples and empirical specifications. First, Lai and Zhu (2003) are only concerned with the determinants of bilateral trade flows. Their analysis involves bilateral trade among 34 countries over 4 years (1980, 1984, 1988, 1992). That is, their study uses a much larger sample. In contrast, in this paper we examine the determinants of trade and multinational production *jointly*. Since detailed data on multinational production are publicly available only for the United States, we have to restrict our sample to the United States and its trading partners. As a result, the sample size is greatly reduced. Second, in Lai and Zhu (2003), the basic specification of bilateral trade flows is similar to equation (6). Their fixed-effects model adds country-pair fixed effects. In contrast, in this study we estimate the U.S. exports equation (6) and multinational production equation (7) simultaneously. By allowing for the correlation of the residuals ε'_{jt} and υ'_{jt}, we partially control for the country fixed effects that influence both U.S. exports and multinational production. This may help explain why our estimate of σ is closer to the fixed-effects estimate in Lai and Zhu (2003).

Our estimated σ of 4.72 is also slightly smaller than other estimates obtained in the literature. For example, the point estimate of σ is 6.43 in Baier and Bergstrand (2001). The average estimates in Hummels (1999) are 5.79, 6.23 and 7.04 for 1-digit, 2-digit and 3-digit manufacturing industries, respectively. The smaller estimate of σ is well in line with the estimates obtained from micro studies of demand for differentiated products (e.g., Feenstra 1994). It is also close to the elasticities that are used in the CGE literature, e.g., in Deardorff and Stern (1990), 17 out of 21 industries involving nonagricultural traded goods have elasticities of substitution that are less than 3.1.

The distance effect is significant. The estimate of μ is 0.04 ($t = 4.53$). With distance scaled to 1000 miles per unit, this estimate indicates that 4 percent of a product's value is lost per thousand miles of shipping. This estimate is slightly smaller than the fixed-effects estimate of 0.05 in Lai and Zhu (2003). Finally, the effect of marginal costs is also significantly positive, as expected. The estimated α is 0.36 ($t = 2.68$).[11]

[11] In Lai and Zhu (2003), the measure of marginal costs is based on wages adjusted by labor productivity or TFP. In this paper, the measure of marginal costs consists of both wages and prices of capital

The parameter ρ captures the correlation of residuals. The estimation yields $\widehat{\rho} = -0.50$ ($t = -7.29$). This means that the effects of excluded factors affecting U.S. exports are significantly and negatively correlated with those of excluded factors affecting U.S. multinational production. This negative correlation is even stronger than that obtained from the linear regressions.

In order to examine how well the model predicts U.S. exports and multinational production, we calculate the correlation between actual $\log X_{jt}^{us}$ and its prediction $\log \widehat{X_{jt}^{us}}$, as well as that between $\log Q_{jt}^{us}$ and $\log \widehat{Q_{jt}^{us}}$. As shown at the bottom of column 1 in Table 2, the correlation is 0.74 between $\log X_{jt}^{us}$ and $\log \widehat{X_{jt}^{us}}$, and is 0.91 between $\log Q_{jt}^{us}$ and $\log \widehat{Q_{jt}^{us}}$. These high correlations are also reflected in Figure 1. The left panel displays the plot of $\log X_{jt}^{us}$ against its prediction. The smaller $\log X_{jt}^{us}$ fits well, but the larger $\log X_{jt}^{us}$ is slightly over-predicted. That is, the model over-predicts U.S. exports to Japan, Germany, Canada and the United Kingdom. The right panel presents the plot of $\log Q_{jt}^{us}$ against its prediction. It shows that the smaller $\log Q_{jt}^{us}$ fits well, but the larger $\log Q_{jt}^{us}$ is slightly under-predicted. This means that the model under-predicts U.S. multinational production in the United Kingdom, Canada, Germany and the Netherlands.

Using equation (7) and the baseline estimates of σ, μ and α in column 1 of Table 2, we further decompose the predicted $\log Q_{jt}^{us}$ into three components: (1) production capacity of U.S. affiliates in country j, $\log n_{jt}^{us}$; (2) production costs in country j, $\log \omega_{jt}^{\widehat{\alpha}(1-\widehat{\sigma})}$; and (3) country j's market access, $\log \left[\sum_{i=1}^{N} \left(e^{\widehat{\mu} D_{ij}} \tau_{ijt}\right)^{1-\widehat{\sigma}} \widehat{P}_{it}^{\widehat{\sigma}-1} Y_{it}\right]$, where $\widehat{P}_{it} = \left[\sum_{j=1}^{N} n_{jt} \left(e^{\widehat{\mu} D_{ij}} \tau_{ijt} w_{jt}^{\widehat{\alpha}}\right)^{1-\widehat{\sigma}}\right]^{\frac{1}{1-\widehat{\sigma}}}$. The correlations between $\log Q_{jt}^{us}$ and the three components are 0.93 (p-value < 0.01), 0.14 (p-value $= 0.10$), and 0.36 (p-value < 0.01), respectively. Thus, controlling for n_{jt}^{us}, the variation in $\log Q_{jt}^{us}$ appears to be more significantly related to variation in market access than variation in production costs. Figure

investment. (See the data appendix for more details.) We have also estimated the equations using the measure in Lai and Zhu (2003). However, similar to the fixed-effects estimate in Lai and Zhu (2003), the estimated effect of production costs is not statistically significant. Thus, we do not report those results in this paper.

2 displays the plots of actual $\log Q_{jt}^{us}$ against predicted production costs and market access.[12] It is apparent that the correlation is stronger for market access than for production costs.

Columns 2-4 of Table 2 report the results when constraints are imposed on α or μ. We first assume $\alpha = 1$. Compared to the baseline estimate $\widehat{\alpha} = 0.36$, this specification allows for a stronger effect of marginal costs on trade flows and multinational production. As shown in column 2, the estimate of μ rises to 0.11, indicating a larger distance effect. On the other hand, the elasticity of substitution $\widehat{\sigma}$ becomes 2.81, which is smaller than the baseline estimate of 4.72 in column 1. Thus, in the presence of stronger effects of marginal costs and distance, a smaller σ is required to explain the variation in trade flows and multinational production.

In column 3, we omit marginal costs by restricting α to be 0. The estimated μ drops slightly to 0.03, while the estimated σ rises to 5.30. In column 4 we exclude distance by assuming $\mu = 0$. As it is apparent, the omission of distance has a substantial impact on the estimates. The estimated α becomes much smaller and statistically insignificant. The estimate of σ increases to 7.20, which is much larger than the baseline estimate of 4.72. The results in columns 3-4 suggest that with smaller effects of marginal costs and distance, a bigger σ is needed to account for the variation in trade and multinational production.

In columns 2-4, the estimates of ρ are very stable. They are between -0.48 and -0.52. The correlation between $\log X_{jt}^{us}$ and $\log \widehat{X_{jt}^{us}}$ or between $\log Q_{jt}^{us}$ and $\log \widehat{Q_{jt}^{us}}$ remains high in all specifications. Finally, the χ^2-statistics of the likelihood ratio tests in columns 2-4 are 5.20, 5.58 and 12.98, respectively. (The 5% critical value is 3.84.) Thus, all restrictions in columns 2-4 are rejected at the 5% significance level. This indicates that

[12]Note that the price index of inputs ω_{jt} is expressed relative to the United States. From equations (6) and (7), it is easy to see that using this normalization does not affect the estimation of structural parameters.

tariffs, distance, and production costs have separate effects on trade and multinational production.

6 Implications for Trade Liberalization

The estimation of structural parameters allows us to evaluate the relative effects of trade policy on trade and multinational production. In the non-linear setting, the percentage change in U.S. exports due to the hypothetical elimination of existing tariffs in any time period t is calculated as:

$$\textit{Tariff effect on } X_{jt}^{us} = \left(E\left[X_{jt}^{us}|\tau_{ij't}=0\right] - E\left[X_{jt}^{us}|\tau_{ij't}>0\right]\right)/E\left[X_{jt}^{us}|\tau_{ij't}>0\right].$$

The tariff effect on Q_{jt}^{us} is defined similarly. We calculate the tariff effects for the year 1992 using the baseline estimates in column 1 of Table 2.

When we do this calculation, we need to keep two caveats in mind. First, since factor prices are exogenous in our model, our estimates may not fully capture general equilibrium effects via changes in factor prices.[13] Second, the number of multinational plants is assumed to be exogenous. The model does not account for the multinationals' long-run decision to reallocate plants in response to trade policy change. Therefore the trade policy effects we obtained are short-run effects.

As trade policy variables vary, U.S. overseas multinational production can either substitute, complement, or be independent of U.S. exports to its trading partners. The substitution scenario is consistent with the hypothesis that multinational production arises from incentives for firms to jump over trade barriers; the complementary scenario implies that both trade and multinational production increase when trade barriers fall, and the

[13]The determinants of wages in a general equilibrium model are carefully examined in Redding and Venables (2002). However, their estimation does not yield the structural parameters that we focus on.

world becomes more integrated as trade becomes more liberalized.

Table 3 shows the tariff effects on U.S. exports and multinational production. In 1992, if bilateral tariffs between the United States and its trading partners were eliminated, U.S. total exports would increase by 20.7 percent. In particular, non-OECD countries would experience a jump in U.S. exports by 58.6 percent, which is much higher than the increase for OECD countries (12.2 percent).[14] Table 3 also reveals substantial variation in the tariff effects at the national level. This large variation in the tariff effects largely reflects the differences among countries in tariff rates and trade volumes prior to trade liberalization. With the elimination of tariffs, the effect on U.S. exports to host countries is straightforward: the higher is the pre-liberalization tariff and the lower is the pre-liberalization trade, the higher is the percentage increase in post-liberalization exports. Within non-OECD countries, India would have the highest growth in U.S. exports (658.7 percent). In contrast, there would be little change in U.S. exports to Hong Kong (0 percent) and Singapore (1.0 percent). Furthermore, compared to non-OECD countries, variation in the tariff effects is relatively small among OECD countries. Austria would experience the highest growth in U.S. exports (22.1 percent) while Canada would see the smallest increase (3.2 percent).

On the other hand, with trade liberalization, total U.S. overseas multinational production would decrease by 3.7 percent. This is consistent with the argument that multinational production arises from incentives for firms to jump over trade barriers. That is, as tariffs fall, multinational production also falls.[15] The percentage decrease would be 5.3 percent for non-OECD countries and 3.3 percent for OECD countries. The larger decrease for non-OECD countries is likely linked to their higher tariffs and lower U.S.

[14] During the period of our analysis (1984–1992), both Mexico and Korea had not joined the OECD. Thus, in our empirical study they are classified as non-OECD member countries.

[15] As implied by part (1) of theorem 2, if country j eliminated tariffs against imports from the United States (i.e., $\tau_{j,us}$ falls to 0), Q_j^{us} would fall. On the other hand, if the United States removed tariffs against country j (i.e., $\tau_{us,j}$ falls to 0), Q_j^{us} would rise. Thus, the negative effect of trade liberalization on U.S. multinational production may reflect that the reduction of $\tau_{j,us}$ has a dominant impact on Q_j^{us}.

multinational production prior to trade liberalization. In addition, the effects on U.S. multinational production in host countries are all negative. However, variation in the tariff effects among countries is small. Within OECD countries, Canada would experience the largest reduction in U.S. multinational production (-5.4 percent) while Portugal would have the smallest decrease (-1.4 percent). Among non-OECD countries, India would see the highest reduction (-8.2 percent) while Korea would have the lowest decrease (-2.9 percent).

In short, these results suggest that trade liberalization may increase U.S. exports significantly, but have only a very small negative effect on U.S. overseas multinational production.

7 Additional Specifications

Table 4 presents the results of sensitivity analysis for the baseline model. The 'baseline' row reports the baseline estimates in column 1 of Table 2. The remaining rows omit all observations related to the indicating country. In general, the estimates are robust when any single country is dropped. The only exception is Indonesia. When Indonesia is excluded, $\widehat{\mu}$ falls to its minimum and becomes statistically insignificant. On the other hand, $\widehat{\sigma}$ rises to its maximum. Notably, the estimates of ρ remain very stable.

In the baseline model the structural parameters σ, μ and α are assumed to be equal in equations (6) and (7). This is because both trade and multinational production are derived from a unified framework as the one illustrated in section 2. As implied by the theory, the two equations should share common structural parameters. In order to examine whether these cross-equation constraints are robust, we relax some or all of the constraints imposed on σ, μ, and α across the two equations. Estimates of these additional specifications are provided in Table 5. We use subscripts 'X' to indicate parameters estimated from the export equation (6), and subscripts 'Q' to indicate parameters from the multinational production equation (7). For comparison, column 1 of Table 5 reports the baseline estimates from column 1 of Table 2.

We first allow all three structural parameters to differ in the two equations (i.e., 'no constraints'). As shown in column 2, both $\widehat{\alpha_X}$ and $\widehat{\mu_X}$ are larger than their counterparts, indicating that marginal costs and distance have bigger impacts on U.S. exports than on U.S. multinational production. On the other hand, $\widehat{\sigma_X}$ (4.06) is much smaller than $\widehat{\sigma_Q}$ (12.49), which indicates that the elasticity of substitution is smaller for U.S. exports than for U.S. multinational production. This difference may arise from the fact that the same U.S. multinational firms are likely to operate in different countries and produce fairly similar or related products in different locations. Thus we expect that the elasticity of substitution would be larger for the varieties produced by multinationals.

The bottom of Table 5 shows that without any parameter constraints, the value of log-likelihood increases to -194.62 (column 2) from -209.91 (column 1). Thus, the likelihood ratio test ($\chi^2 = 30.58$ in column 1) rejects the cross-equation constraints at the 5% significance level. In order to understand the sources of this rejection, we impose only one cross-equation constraint each time. Specifically, in columns 3-5 we assume $\alpha_X = \alpha_Q$, $\sigma_X = \sigma_Q$, and $\mu_X = \mu_Q$, respectively.

Compared to the case of 'no constraints' in column 2, the values of log-likelihood drop slightly in columns 3 and 4. As shown at the bottom of Table 5, the χ^2-statistics of the likelihood ratio tests are 0.48 in column 3 and 3.24 in column 4. Thus, the hypotheses of $\alpha_X = \alpha_Q$, and $\sigma_X = \sigma_Q$ cannot be rejected at the 5% significance level. In contrast, when the restriction of $\mu_X = \mu_Q$ is imposed, the value of log-likelihood is reduced dramatically to -206.77 (see column 5). The likelihood ratio test ($\chi^2 = 24.3$) indicates that this constraint can be rejected at the 5% significance level. Therefore, the rejection of the joint cross-equation constraints on σ, μ and α is likely caused by the significantly different distance effects on U.S. exports and multinational production.

To investigate further, in column 6 we assume $\alpha_X = \alpha_Q$ and $\sigma_X = \sigma_Q$. The value of log-likelihood is reduced to -196.64. The χ^2-statistic of the likelihood ratio test is 4.04, suggesting that the joint restrictions of $\alpha_X = \alpha_Q$ and $\sigma_X = \sigma_Q$ cannot be rejected at the 5% significance level. This result provides further evidence that the different distance effects on U.S. exports and multinational production likely lead to the rejection of joint cross-equation constraints on all three structural parameters.

8 Conclusions

In this paper, we developed a monopolistic competition model of trade and multinational production that allows for asymmetric trade barriers and international differences in production costs. The model implies highly non-linear equations of U.S. exports and multinational production. To deal with the simultaneity between trade flows and multinational production, we took a radically different approach from the existing literature. Specifically, within a unified framework, we explored the interdependence between trade and multinational production in the form of cross-equation restrictions in parameters and the correlation of error terms. Our estimation yields reasonable estimates of the structural parameters including the elasticity of substitution. This is a significant contribution to the literature: previous studies placed zero restrictions on at least some of these parameters. Further, we find that the error terms in the equations of U.S. exports and multinational production are significantly and negatively correlated, suggesting that the omitted variables have opposing effects on U.S. exports and multinational production.

The estimates also allowed us to assess the impact of trade liberalization on U.S. exports and overseas multinational production. We find that the elimination of tariffs would significantly increase U.S. exports but insignificantly reduce its overseas multinational production. We must emphasize that we focused on firms' short-run production decisions. We did not model firms' long-run plant allocation decisions. As a result, the estimated effects of trade liberalization are short-run effects.

A Appendix: Data Sources and Measurement

This empirical study draws data from various sources. Trade flows (including the U.S. exports X_{jt}^{us}) come from the World Trade Database (Feenstra *et al.* 1997). Data on U.S. multinational production (Q_{jt}^{us}) are from the BEA published tables *Gross Product of Affiliates, Country by Industry.* Differing from affiliate sales, gross product excludes the costs of materials and other intermediate inputs.

The expenditures on manufacturing goods (Y_{jt}) are calculated as manufacturing output plus net imports. Gross manufacturing output is from UNIDO's INDSTAT database. Bilateral tariff data (τ_{ijt}) have been carefully compiled from different sources and took 8 months of full-time work to construct. See Lai and Trefler (2002) for more details. Distance (D_{ij}) comes from Antweiler (1996). The data are compiled using information on longitudes and latitudes of major cities in the world along with their population figures. Great-circle distances between major cities are computed. An 'average' distance between big countries such as Canada and the United States is calculated as a weighted average of all pairwise distances between major cities in the two countries. The weights are chosen such that a greater weight is given to populous cities and to neighboring cities.

In order to obtain an international comparable price index of inputs (ω_{jt}), we adjust prices of inputs by TFP. Let w_{jt} denote the wage rate in country j and year t. Let e_{jt} be the number of employment, and v_{jt} the value added. Define the labor share as $\theta_{jt} \equiv w_{jt}e_{jt}/v_{jt}$. Let r_{jt} be the price of capital investment, and k_{jt} the capital stock. Let N be the total number of countries. Under the assumption of Cobb-Douglas production function, ω_{jt} is calculated using the following formula:

$$\omega_{jt} \equiv \frac{\left(w_{jt}/w_{US,t}\right)^{\widetilde{\theta}_{jt}} \left(r_{jt}/r_{US,t}\right)^{1-\widetilde{\theta}_{jt}}}{TFP_{jt}/TFP_{US,t}},$$

where $\widetilde{\theta}_{jt} \equiv \left(\theta_{jt} + \frac{1}{N}\sum_{j=1}^{N}\theta_{jt}\right)/2$, and $\ln\left(TFP_{jt}\right) \equiv \ln\left(v_{jt}\right) - \widetilde{\theta}_{jt}\ln\left(e_{jt}\right) - \left(1-\widetilde{\theta}_{jt}\right)\ln\left(k_{jt}\right)$.

The calculation of the multilateral TFP index follows Caves, Christensen and Dieweit (1982) and Keller (2002). The price of capital investment is measured by the price index of investment from the Penn World Table 6.1. This price index of investment (in international dollars) is expressed relative to the United States. That is, we only have data on the ratio of $r_{jt}/r_{US,t}$ available.[16] This forces us to define ω_{jt} in terms of the U.S. price index. Wages (w_{jt}), employment (e_{jt}), value added (v_{jt}), and gross fixed capital formation series come from UNIDO. The gross fixed capital formation series are converted into U.S. dollars and deflated using the price levels of investment from the Penn World Table 6.1. The deflated gross fixed capital formation series are then used to construct capital stocks (k_{jt}) using a 10-year double declining balance method. The depreciation rate is 0.20. Note that $\omega_{us,t} \equiv 1$ by definition. As mentioned in footnote 12, this normalization does not affect the estimation of structural parameters.

We draw data on the number of establishments (n_{jt}) from the *Yearbook of Industrial Statistics* and the *International Yearbook of Industrial Statistics*. An 'establishment' is a unit which engages, under a single ownership or control, in one, or predominantly one, kind of activity at a single location, for example, workshop or factory. Finally, the number of U.S. overseas plants (n_{jt}^{us}) comes from BEA published tables *Selected Data for Foreign Affiliates in All Countries in Which Investment Was Reported.*

[16] Under the assumption that interest rates and depreciation rates are equalized across countries, the ratio of $r_{jt}/r_{US,t}$ is equal to the ratio of user costs of capital in country j relative to the United States.

References

Anderson, James E. and Eric van Wincoop (2003) "Gravity with Gravitas: A Solution to the Border Puzzle," *American Economic Review*, *93* (1), 170–92.

Antweiler, Werner (1996) "Three Essays on International Trade: Hysteresis, R&D, and the Environment." Ph.D. Thesis, University of Toronto.

Baier, Scott L. and Jeffrey H. Bergstrand (2001) "The Growth of World Trade: Tariffs, Transport Costs, and Income Similarity," *Journal of International Economics*, *53* (1), 1–27.

Bergstrand, Jeffrey H. (1985) "The Gravity Equation in International Trade: Some Microeconomic Foundations and Empirical Evidence," *Review of Economics and Statistics*, *67* (3), 474–81.

Brainard, S. Lael (1993) "A Simple Theory of Multinational Corporations and Trade with a Trade-off between Proximity and Concentration." NBER Working Paper #4269.

—— (1997) "An Empirical Assessment of the Proximity-Concentration Trade-off Between Multinational Sales and Trade," *American Economic Review*, *87* (4), 520–44.

Carr, David L., James R. Markusen, and Keith E. Maskus (2001) "Estimating the Knowledge-Capital Model of the Multinational Enterprise," *American Economic Review*, *91* (3), 693–708.

Caves, Douglas W., Laurits R. Christensen, and Erwin W. Diewert (1982) "The Economic Theory of Index Numbers and the Measurement of Input, Output, and Productivity," *Econometrica*, *50* (6), 1393–1414.

Deardorff, Alan V. and Robert M. Stern (1990) *Computational Analysis of Global Trading Arrangements*, Ann Arbor, MI: University of Michigan Press.

Feenstra, Robert C. (1994) "New Product Varieties and the Measurement of International Prices," *American Economic Review*, *84* (1), 157–77.

—— **, Robert E. Lipsey, and Harry P. Bowen** (1997) "World Trade Flows, 1970-1992, With Production and Tariff Data." NBER Working Paper #5910.

Hanson, Gordon H., Raymond J. Mataloni, and Matthew J. Slaughter (2001) "Expansion Strategies of U.S. Multinational Firms," in Dani Rodrik and Susan Collins, eds., *Brookings Trade Forum 2001*, Washington. D.C.: Brookings.

Helpman, Elhanan (1981) "International Trade in the Presence of Product Differentiation, Economies of Scale and Monopolistic Competition: A Chamberlin-Heckscher-Ohlin Approach," *Journal of International Economics*, *11* (3), 305–40.

—— (1984) "A Simple Theory of International Trade with Multinational Corporations," *Journal of Political Economy*, *92* (3), 451–71.

—— **and Paul R. Krugman** (1985) *Market Structure and Foreign Trade: Increasing Returns, Imperfect Competition, and the International Economy*, Cambridge MA: The MIT Press.

—— **, Marc J. Melitz, and Stephen R. Yeaple** (2003) "Export versus FDI with Heterogenous Firms," *American Economic Review, forthcoming.*

Hummels, David (1999) "Toward a Geography of Trade Costs." Mimeo, University of Chicago.

—— **and James Levinsohn** (1995) "Monopolistic Competition and International Trade: Reconsidering the Evidence," *Quarterly Journal of Economics*, *110* (3), 799–836.

Keller, Wolfgang (2002) "Geographic Localization of International Technology Diffusion," *American Economic Review*, *92* (1), 120–42.

Krugman, Paul R. (1980) "Scale Economies, Product Differentiation, and the Pattern of Trade," *American Economic Review*, *70* (5), 950–59.

Lai, Huiwen and Daniel Trefler (2002) "The Gains From Trade with Monopolistic Competition: Specification, Estimation, and Mis-specification." NBER Working Paper #9169.

—— **and Susan Chun Zhu** (2003) "The Determinants of Bilateral Trade," *Canadian Journal of Economics, forthcoming.*

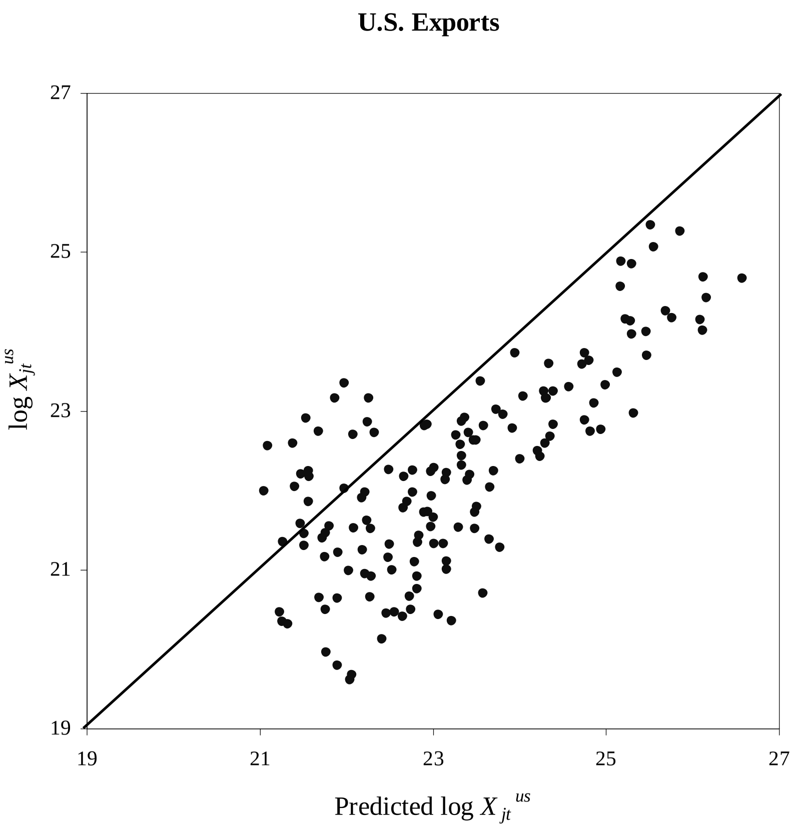
Markusen, James R. (1984) "Multinationals, Multi-plant Economies, and the Gains from Trade," *Journal of International Economics*, *16* (3-4), 205–26.

—— (2002) *Multinational Firms and the Theory of International Trade*, Cambridge, MA: The MIT Press.

—— **and Anthony J. Venables** (2000) "The Theory of Endowment, Intra-industry and Multi-national Trade," *Journal of International Economics*, *52* (2), 209–34.

Mataloni, Raymond (1995) "A Guide to BEA Statistics on U.S. Multinational Companies," *Survey of Current Business.*

Redding, Stephen and Anthony J. Venables (2002) "Economic Geography and International Inequality." Mimeo, London School of Economics.



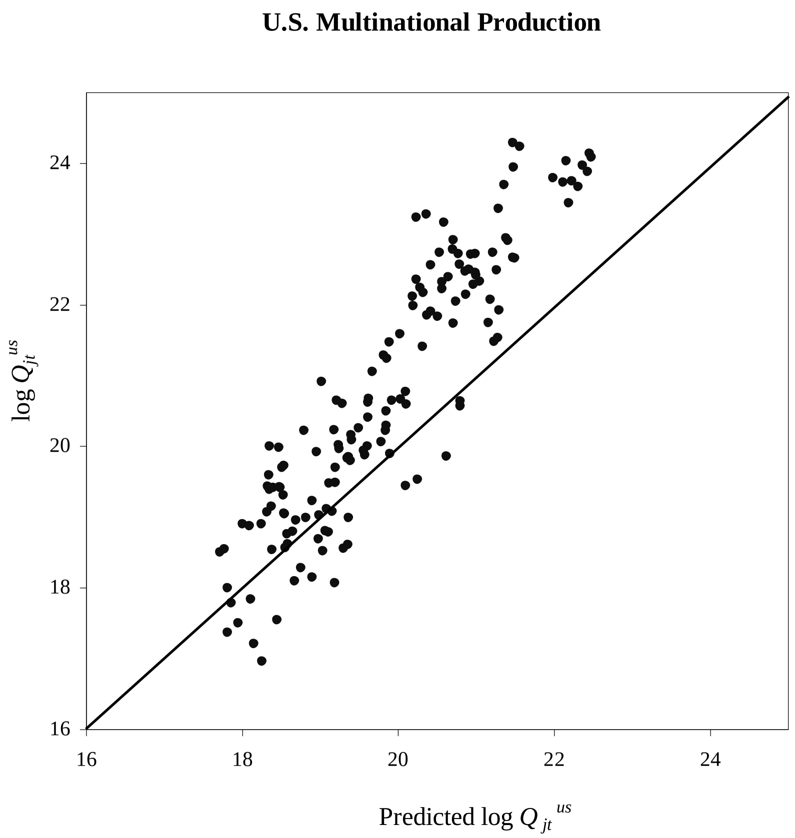


Figure 1. Model Fit: U.S. Exports Equation and Multinational Production Equation

Notes: The predicted values of $\log X_{jt}^{us}$ and $\log Q_{jt}^{us}$ are calculated using the baseline estimates in column 1 of Table 2.

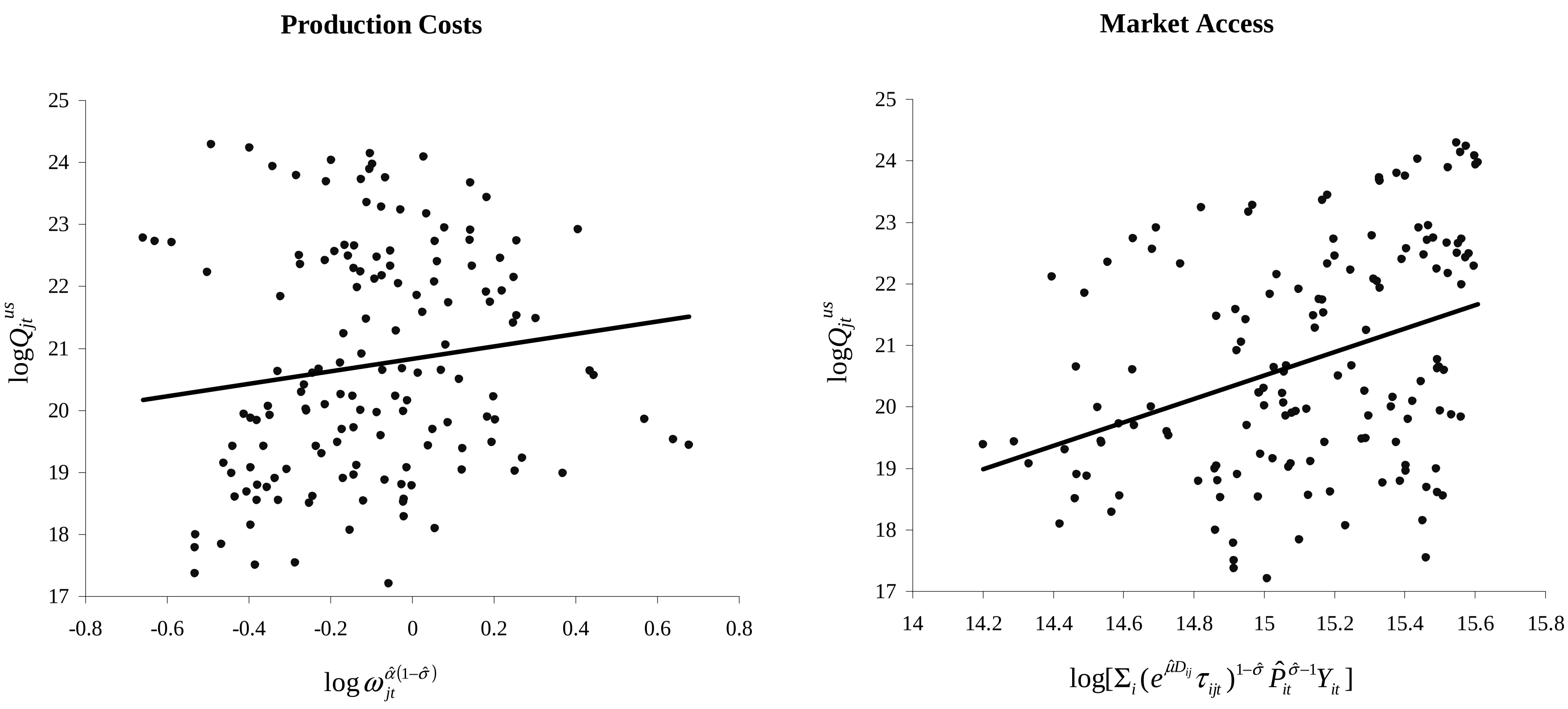


Figure 2. Determinants of U.S. Multinational Production: Production Costs vs. Market Access

Notes: The predicted values of production costs and market access are calculated using the baseline estimates in column 1 of Table 2. See section 5 for more details.

Table 1. Linear Regressions

	OLS			SUR[a]	
Dependent Variable	$\log X_{jt}^{us}$	$\log Q_{jt}^{us}$		$\log X_{jt}^{us}$	$\log Q_{jt}^{us}$
	(1)	(2)	(3)	(4)	(5)
$\log \tau_{j,us,t}$	-0.80		0.57	-0.80	0.57
	(-1.94)		(1.69)	(-1.95)	(1.68)
$D_{j,us}$	-0.01		0.04	-0.01	0.03
	(-0.28)		(1.69)	(-0.25)	(1.56)
$\Sigma_i Y_{it} \log \tau_{ijt}$		-7.25	-8.63		-9.11
		(-2.37)	(-2.71)		(-2.89)
$\Sigma_i Y_{it} D_{ij}$		-0.02	-0.07		-0.05
		(-0.86)	(-2.05)		(-1.48)
$\log \omega_{jt}$		-0.32	-0.44		-0.49
		(-0.95)	(-1.32)		(-1.49)
$\log Y_{jt}$	0.73			0.74	
	(12.28)			(12.52)	
$\log n_{us,t}$	0.57			0.56	
	(1.88)			(1.85)	
$\log n_{jt}^{us}$		1.64	1.67		1.72
		(27.92)	(28.04)		(28.97)
R^2	0.54	0.88	0.88	0.54	0.87

Notes: This table reports estimates of equations (8) and (9). There are 150 observations involving 30 countries and 5 years (1984, 1986, 1988, 1990 and 1992). t-statistics are in parentheses.

a) Columns 4-5 report results of seemingly unrelated regressions (SUR). The correlation of the residuals in equations (8) and (9) is -0.22.

Table 2. Maximum-Likelihood Estimates with Cross-Equation Constraints

	Baseline	α=1	α=0	μ=0
	(1)	(2)	(3)	(4)
σ	4.72	2.81	5.30	7.20
	(4.46)	(1.22)	(4.62)	(5.88)
μ	0.04	0.11	0.03	0
	(4.53)	(7.19)	(2.30)	
α	0.36	1	0	0.09
	(2.68)			(0.99)
ρ	-0.50	-0.52	-0.48	-0.50
	(-7.29)	(-7.49)	(-6.99)	(-7.72)
$corr(\log X_{jt}^{us}, \log \hat{X}_{jt}^{us})$	0.74	0.73	0.74	0.74
$corr(\log Q_{jt}^{us}, \log \hat{Q}_{jt}^{us})$	0.91	0.90	0.92	0.93
Log-likelihood	-209.91	-212.51	-212.70	-216.40
Likelihood ratio test (χ^2-statistics)[a]		5.20	5.58	12.98

Notes: This table reports the maximum-likelihood estimates of equations (6) and (7). There are 150 observations involving 30 countries and 5 years (1984, 1986, 1988, 1990 and 1992). t-statistics are in parentheses.

a) The hypotheses for the likelihood ratio tests in columns 2-4 are α=1, α=0, and μ=0, respectively. The 5% critical value is 3.84.

Table 3. The Estimated Tariff Effects on U.S. Exports and Multinational Production

	Exports		Multinational Production		Tariffs Rates	
	Tariff Effects	Volumes	Tariff Effects	Volumes	$\tau_{j,us,92}$	$\tau_{us,j,92}$
All countries	**20.7%**	**418.7**	**-3.7%**	**156.3**		
OECD countries	**12.2%**	**285.1**	**-3.3%**	**129.6**		
Australia	16.2%	11.7	-4.0%	5.1	5.4%	3.9%
Austria	22.1%	2.2	-2.9%	0.9	7.1%	4.7%
Belgium	21.1%	7.4	-5.0%	5.9	6.8%	0.0%
Canada	3.2%	101.3	-5.4%	21.6	2.1%	0.4%
Denmark	14.5%	1.8	-2.9%	0.5	4.9%	4.1%
Finland	7.1%	1.2	-3.1%	0.1	2.5%	3.6%
Germany	14.5%	30.8	-2.9%	35.6	4.9%	4.0%
Greece	15.1%	1.0	-3.2%	0.3	4.9%	5.2%
Ireland	14.0%	3.7	-2.6%	4.6	4.9%	4.2%
Italy	15.3%	10.7	-2.4%	8.9	5.0%	5.4%
Japan	10.7%	52.4	-2.7%	7.9	3.4%	3.4%
Netherlands	15.3%	12.5	-3.0%	7.0	5.1%	3.8%
New Zealand	11.6%	2.2	-3.0%	0.2	4.5%	4.5%
Norway	8.5%	2.6	-3.2%	0.1	3.0%	3.4%
Portugal	14.0%	1.0	-1.4%	0.6	5.0%	6.9%
Spain	14.9%	7.6	-2.7%	5.8	5.1%	4.3%
Sweden	9.3%	4.6	-3.1%	0.9	3.2%	3.4%
United Kingdom	13.8%	30.3	-2.8%	23.7	4.8%	3.8%
Non-OECD countries	**58.6%**	**133.6**	**-5.3%**	**26.7**		
Brazil	85.9%	8.2	-7.7%	12.4	26.6%	2.7%
Chile	43.3%	3.1	-8.0%	0.4	19.3%	1.2%
Ecuador	14.1%	1.3	-8.0%	0.0	7.9%	0.1%
Hong Kong	0.0%	11.5	-2.0%	0.9	0.0%	7.9%
India	658.7%	2.2	-8.2%	0.1	91.6%	5.0%
Indonesia	48.7%	4.3	-3.5%	0.1	13.3%	7.2%
Korea[a]	53.1%	20.2	-2.9%	0.6	15.4%	5.2%
Malaysia	54.8%	7.6	-5.0%	1.2	14.7%	1.5%
Mexico[a]	28.4%	46.8	-7.3%	7.4	21.8%	1.7%
Singapore	1.0%	13.8	-4.2%	2.1	0.3%	3.3%
Thailand	96.0%	6.4	-5.5%	0.5	24.6%	3.4%
Venezuela	31.9%	8.1	-6.4%	1.0	18.2%	1.1%

Notes: This table reports the estimated effects on U.S. exports and multinational production if 1992 bilateral tariffs between the United States and its trading partners were eliminated. The tariff effects are calculated using the estimates in column 1 of Table 2. The volumes of 1992 U.S. exports and multinational production are expressed in billions of U.S. dollars. $\tau_{j,us,92}$ represents country *j*'s tariffs against imports from the United States for the year 1992. $\tau_{us,j,92}$ represents 1992 U.S. tariffs against imports from country *j*.

a) Since both Mexico and Korea had not joined the OECD during the period of our analysis, they are classified as non-OECD countries.

Table 4. Sensitivity Analysis

	σ	t-stat	μ	t-stat	α	t-stat	ρ	t-stat
Baseline	4.72	4.46	0.04	4.53	0.36	2.68	-0.50	-7.29
Omitting the Indicated Country								
Indonesia	5.54	3.65	0.02	1.47	0.25	1.91	-0.54	-8.23
Finland	5.30	5.20	0.03	3.25	0.24	2.01	-0.59	-10.11
Greece	5.02	4.73	0.03	3.54	0.30	2.16	-0.53	-8.00
Norway	5.02	4.68	0.03	3.53	0.27	2.01	-0.55	-8.66
Japan	4.93	4.57	0.04	4.11	0.41	3.20	-0.49	-7.09
Italy	4.93	4.48	0.04	3.59	0.28	1.91	-0.47	-6.54
Germany	4.92	4.48	0.04	3.81	0.38	3.03	-0.47	-6.54
Spain	4.90	4.57	0.04	3.91	0.30	2.15	-0.47	-6.45
Canada	4.89	4.22	0.04	3.24	0.35	2.54	-0.52	-7.57
Denmark	4.87	4.55	0.04	3.96	0.32	2.28	-0.51	-7.57
United Kingdom	4.84	4.38	0.04	3.76	0.31	2.21	-0.48	-6.84
India	4.82	2.49	0.04	4.53	0.37	2.82	-0.50	-6.70
Ireland	4.80	4.33	0.04	3.86	0.35	2.57	-0.52	-7.60
Netherlands	4.78	4.38	0.04	4.10	0.35	2.50	-0.50	-7.11
Sweden	4.78	4.48	0.04	4.33	0.35	2.59	-0.50	-7.16
Austria	4.78	4.41	0.04	4.02	0.33	2.32	-0.50	-7.32
Ecuador	4.77	4.36	0.04	3.83	0.30	1.93	-0.51	-7.44
Thailand	4.77	4.36	0.04	4.09	0.38	2.71	-0.50	-7.13
Belgium	4.76	4.36	0.04	4.15	0.36	2.60	-0.50	-7.14
Korea	4.74	4.36	0.04	4.19	0.35	2.52	-0.50	-7.16
New Zealand	4.73	4.55	0.04	4.87	0.38	2.82	-0.49	-6.70
Portugal	4.72	4.77	0.04	5.24	0.41	3.05	-0.52	-7.43
Brazil	4.72	4.11	0.05	5.16	0.34	2.64	-0.47	-6.23
Venezuela	4.60	4.35	0.05	4.91	0.39	2.77	-0.49	-6.87
Australia	4.57	4.32	0.05	5.58	0.39	2.89	-0.49	-6.93
Mexico	4.55	4.33	0.05	5.04	0.38	2.71	-0.49	-6.61
Chile	4.54	4.38	0.05	5.29	0.40	2.92	-0.48	-6.34
Hong Kong	4.43	4.14	0.05	4.97	0.43	2.82	-0.48	-6.23
Malaysia	3.94	3.82	0.07	8.01	0.55	3.70	-0.52	-7.40
Singapore	3.79	3.87	0.08	9.04	0.54	3.50	-0.52	-7.45

Notes: 'Baseline' refers to the estimates in column 1 of Table 2. Each of the other rows refers to the case when observations related to the indicated country are excluded. The countries are sorted by σ in a descending order.

Table 5. Additional Specifications

	Baseline	no constraints	$\alpha_X = \alpha_Q$	$\sigma_X = \sigma_Q$	$\mu_X = \mu_Q$	$\alpha_X = \alpha_Q$ $\sigma_X = \sigma_Q$
	(1)	(2)	(3)	(4)	(5)	(6)
U.S. Exports						
σ_X	4.72	4.06	4.01	4.07	5.04	4.02
	(4.46)	(4.24)	(4.26)	(4.21)	(4.91)	(4.23)
μ_X	0.04	0.08	0.08	0.08	0.03	0.08
	(4.53)	(10.81)	(10.48)	(11.24)	(4.67)	(10.67)
α_X	0.36	0.65	0.48	0.65	-0.00	0.42
	(2.68)	(2.32)	(3.57)	(2.33)	(-0.00)	(3.24)
U.S. Multinational Production						
σ_Q	4.72	12.49	12.63	4.07	11.45	4.02
	(4.46)	(3.19)	(3.26)	(4.21)	(3.04)	(4.23)
μ_Q	0.04	-0.00	-0.00	-0.01	0.03	-0.01
	(4.53)	(-1.35)	(-1.35)	(-0.81)	(4.67)	(-0.80)
α_Q	0.36	0.12	0.48	0.37	0.15	0.42
	(2.68)	(3.11)	(3.57)	(2.79)	(3.40)	(3.24)
ρ	-0.50	-0.43	-0.44	-0.42	-0.53	-0.43
	(-7.29)	(-5.65)	(-5.76)	(-5.52)	(-8.18)	(-5.64)
$corr(\log X_{jt}^{us}, \log \hat{X}_{jt}^{us})$	0.74	0.71	0.71	0.71	0.75	0.70
$corr(\log Q_{jt}^{us}, \log \hat{Q}_{jt}^{us})$	0.91	0.93	0.91	0.92	0.92	0.92
Log-likelihood	-209.91	-194.62	-194.86	-196.24	-206.77	-196.64
Likelihood ratio test (χ^2-statistics)[a]	30.58		0.48	3.24	24.30	4.04

Notes: This table reports estimates of equations (6) and (7) when some or all of the cross-equation constraints are relaxed. For comparison, column 1 of this table reports the baseline estimates from column 1 of Table 2. There are 150 observations involving 30 countries and 5 years (1984, 1986, 1988, 1990 and 1992). t-statistics are in parentheses.

a) The hypothesis for the likelihood ratio test in column 1 is $\alpha_X = \alpha_Q$, $\mu_X = \mu_Q$, and $\sigma_X = \sigma_Q$ jointly. The 5% critical value from a χ^2 distribution with 3 degrees of freedom is 7.82. The hypotheses in columns 3-5 are $\alpha_X = \alpha_Q$, $\mu_X = \mu_Q$, and $\sigma_X = \sigma_Q$, respectively. The 5% critical value from a χ^2 distribution with 1 degree of freedom is 3.84. Finally, the hypothesis in column 6 is $\alpha_X = \alpha_Q$ and $\sigma_X = \sigma_Q$ jointly. The 5% critical value from a χ^2 distribution with 2 degrees of freedom is 5.99.