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Eliminating Excessive Tariffs on Exports of Least Developed Countries

Bernard Hoekman, Francis Ng, and Marcelo Olarreaga

Although average OECD tariffs on imports from the least developed countries are very low; tariffs above 15 percent (peaks) have a disproportional effect on their exports. Products subject to tariff peaks tend to be heavily concentrated in agriculture and food products and labor-intensive sectors, such as apparel and footwear. Although the least developed countries benefit from preferential access, preferences tend to be smallest for tariff peak products. A major exception is the European Union, so that the recent European initiative to grant full duty-free and quota-free access for the least developed countries (the so-called Everything But Arms initiative) will result in only a small increase in their exports of tariff peak items (less than 1 percent of total exports). However, as preferences are less significant in other major OECD countries, a more general emulation of the European Union initiative would increase the least developed countries' total exports of peak products by US$2.5 billion (11 percent of total exports). Although almost half of this increase is at the expense of other developing country exports, this represents less than 0.05 percent of their total exports. This trade diversion can be avoided by reducing tariff peaks to a uniform 5 percent applied on a nondiscriminatory basis. However, such a reform would imply no gains for the least developed countries, suggesting that the globally welfare-superior policy of nondiscriminatory elimination of tariff peaks should be complemented by greater direct assistance to poor countries.

Despite generally low average tariffs, the structure of protection in the Quad (Canada, the European Union, Japan, and the United States) is characterized by many tariffs above 15 percent. Such tariff peaks are often concentrated in products that developing countries export. They include major agricultural staple food products such as sugar, cereals, and fish; tobacco; certain alcoholic beverages, fruits, and vegetables; food industry products with a high sugar content; clothing; and footwear.

The existence of these peaks is a reflection of the political economy of trade policy. Powerful groups in the Organisation for Economic Co-operation and

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Development (OECD) countries have been able to keep barriers high partly because of the strategy followed by many developing countries of not participating in the reciprocal exchange of liberalization commitments under the General Agreement on Tariffs and Trade (GATT). The last set of GATT negotiations, the Uruguay Round, actually led to an increase in peaks, as tariffication of nontariff barriers in agriculture led to the imposition of very high duties on products that had previously been constrained by quotas. As a result, tariffs that are more than three times higher than the average tariff are not uncommon in OECD countries.

To some extent, the negative impact of excessive tariffs is offset by preferential access schemes such as the Generalized System of Preferences (GSP) and related programs. These should, in principle, help developing country exporters overcome the high tariff hurdles. In practice, preferences tend to be limited. The programs often exclude "sensitive" products or subject them to some type of quantitative limitation, either in terms of the amount that can be imported under the preferential rates (a tariff rate quota) or in terms of the countries that are eligible (Michalopoulos 1999, Hallaert 2000).

This article assesses the potential effects on least developed country (LDC) exports of duty-free access to the Quad markets on tariff peak items.[1] The article is motivated by the European Union's Everything But Arms initiative, which offers duty-free and quota-free access to the European Union for all LDC exports except arms.[2] We also assess the impact of extending duty-free access for tariff peak items to non-LDC developing country exporters and compare this to a reduction in peaks to a uniform 5 percent tariff that applies on a nondiscriminatory basis.

The article focuses on tariff peaks only—tariffs above 15 percent—for a number of reasons. First, peaks affect commodities that are of relatively greater importance to LDCs—that is, they account for a larger share of total LDC exports. Second, from a political economy point of view, these products face the highest protection in the Quad. Third, moving beyond tariff peaks to consider elimination of all tariffs on all imports from LDCs requires the use of a computable general equilibrium model of the world economy. Although such tools are readily available, they do not allow a detailed and disaggregated analysis of the effect of liberalization at the product and country level. Because we are interested in determining the likely impact of duty-free access on the pattern and composition of trade, we use a partial equilibrium approach. By limiting our focus to tariff peaks—which account for only a small share of the total

1. We use the U.N. definition of LDCs comprising 49 countries. See Hoekman, Ng, and Olarreaga (2001) for a list of these countries. The 49th member of this group, Senegal, is not included in the calculations that follow as it became an LDC after this article was completed.

2. The initiative entered into force in 2001, with the exception of exports of bananas (excluded until 2006) and sugar and rice (excluded until 2009). For a comprehensive description of the Everything But Arms proposal, see UNCTAD (2001).

trade of most countries—such an approach is unlikely to lead to misleading conclusions.[3] The exaggeration of resource shifts associated with a partial equilibrium approach is also likely to be mitigated by restricting attention to tariff peaks only. Finally, the partial equilibrium approach has an important advantage in that the effects of peaks are not diluted through aggregation into a small number of sectors, as is required if a general equilibrium simulation approach is used.

We do not quantify the effect of remaining nontariff barriers—the focus is solely on tariff peaks and preferences. Nontariff barriers are relatively unimportant in the Quad—only 1.2 percent of tariff lines are subject to such barriers in Canada, 4.2 percent in Europe, 2.6 percent in Japan, and 2.9 percent in the United States (see OECD 1997). However, nontariff barriers apply to a sector that is of great interest to developing countries—clothing, which is still constrained by quotas. By ignoring quotas and the associated rents, our analysis underestimates the potential export response by LDCs following duty-free and quota-free access for tariff peak products. However, given that the quotas restrict the most efficient developing country exporters and that all remaining quotas must be eliminated by the end of 2004 under the World Trade Organization (WTO) Agreement on Textiles and Clothing, any effect of preferential elimination of prevailing quotas for LDC exporters will be of short duration.

In the case of agriculture, the Uruguay Round led to tariffication of all nontariff barriers (with the exception of rice in Japan).[4] Tariff rate quotas are often used. These involve two-tier tariff systems: A limited quantity enters subject to a low tariff, and all imports that exceed this tariff quota are subject to higher tariffs. In this article, we use out-of-quota tariffs as the appropriate measure of protection. This may lead to an overestimation of the effects of duty-free access. Note that the net effect of ignoring textile quotas and using out-of-quota tariffs for agricultural products is unclear, but that the two potential sources of bias are offsetting. In general, the approach used is conservative in that we assume only limited supply responsiveness to the changed incentives in LDCs.

Section I describes the extent and importance of existing tariff peaks in the Quad. The article discusses the preferential treatment granted to developing countries for these tariff peak products (section II) and the prevailing pattern of developing country exports (section III).[5] Section IV presents a simple partial equilibrium model. Section V uses this model to estimate the impact of duty-free

3. See Ianchovichina, Mattoo, and Olarreaga (2001) and UNCTAD (2001) for general-equilibrium-based estimates of the gains for developing countries of unrestricted market access for all goods in the Quad. Although the methodology, product, and country coverage in these two studies differ from the present article, they all provide similar estimates of the increase in LDC export revenue and the export displacement for other developing countries.

4. Specific tariffs—frequently used for agricultural products in the Quad—have been converted into ad valorem equivalents using OECD (1997, 2000).

5. Under preferential treatment, we include both unilateral schemes, such as GSP, and those granted under bilateral free trade agreements.

access for LDCs to the Quad and compares this with a nondiscriminatory reduction of peaks to 5 percent. Section VI concludes the article.

I. Tariff Peaks and Imports in the Quad

Between 6 and 14 percent of Quad tariff categories at the six-digit level of aggregation of the Harmonized System classification are above 15 percent (table 1). The United States, the European Union, and Japan have 200 to 300 such lines, whereas Canada has more than 700 tariff peaks. The average unweighted tariff in the Quad over all tariff peak products is 28 percent, more than four times the unweighted total average tariff of 6.2 percent. In the United States and Canada, most tariff peaks affect industrial products (more than 85 percent); in the European Union and Japan, most peaks affect agricultural products (91 and 77 percent, respectively). The maximum tariff rate at this level of aggregation is 340 percent for butter in Canada, 250 percent for edible bovine offal in the European Union, 170 percent for raw cane sugar in Japan, and 120 percent for ground nuts in shell in the United States.

In 1999, imports of products subject to tariff peaks in at least one member were US$92.8 billion. More than 60 percent of these imports originated in developing countries (US$55.2 billion) and *potentially* faced an average tariff of 28 percent.[6] This represented around 5 percent of total developing country exports to these high-income country markets. LDCs are more specialized in products subject to peaks, which affect 11 percent of their total exports to the Quad.

II. Tariff Peaks and Developing Country Preferences

Most developing countries enjoy preferential access to Quad markets, either through unilateral schemes such as the GSP, or through free trade agreements. In the case of Canada, Japan and the European Union, some 170 developing countries benefit from GSP (or better) preferences.[7] In the case of the United States, 140 developing countries benefit from some type of preferential access. As shown in table 2, preferences are of a cascading nature; countries with free trade agreements generally get the best treatment, followed by LDCs and other developing countries. The United States grants preferences to the Andean Pact, the Caribbean, and Mexico (under the North American Free Trade Agreement).[8] The European Union provides preferences for a large group of African, Caribbean,

6. Tariff preferences granted to developing countries through bilateral or unilateral schemes will bring down the tariff faced by these exporters.

7. The European Union was the first customs territory to grant GSP preferences to developing countries in 1971. See Kennan and Stevens (1997) or Hallaert (2000) for a detailed description of the European GSP.

8. In the simulations discussed in the following material, we also include preferences for developed countries that benefit from preferences in other Quad markets (see the notes in table 2).

TABLE 1. Tariff Peaks and Imports by the Quad, 1999

Tariff peak product (at HS six-digit level)	Canada	European Union 15[a]	Japan	United States	All Quad
Number of tariff peak products (MFN≥15%)[b]	732	317	233	307	1,077[c]
Agriculture products[b]	85	290	178	48	364[c]
Industrial products[b]	647	27	55	263	713[c]
Tariff peak products as % of all tariff lines	14.3	6.2	4.6	6.1	7.8[d]
Average MFN tariff rates (unweighted in %)					
Tariff peak products	30.5	40.3	27.8	20.8	28.0
All products	8.3	7.4	4.3	5.0	6.2
Maximum rate	342.7	251.9	170.5	121.0	221.5
Total imports of tariff peak products (US$ billions)	8.7	27.1	15.8	41.2	92.8
All preferential and GSP countries	7.6	16.5	4.8	26.3	55.2
Least developed countries[e]	0.09	0.3	0.03	0.9	1.3
Share of tariff peak products in total imports (%)	4.6	3.4	4.9	4.6	4.2
of which: All preferential & GSP countries (%)	4.8	4.9	2.8	6.6	5.2
Least developed countries[e]	30.2	2.8	2.6	15.0	11.4
Import revenue collection in tariff peak products from world (in US$ billions)	1.6	8.9	6.3	5.4	22.2
All preferential and GSP developing countries	0.7	4.3	1.4	4.6	11.0
Least developed countries[e]	0.02	0.03	0.001	0.2	0.2

[a]Excludes all European Union intra-trade in world totals.

[b]There are no overlapping items in the Quad aggregates.

[c]Number of nonoverlapping categories.

[d]This is the simple (unweighted) average across Quad countries. Note that of the 5,032 tariff lines at the six-digit level of the Harmonized System, 21 percent (1,077/5,032) include a tariff peak item in at least one Quad member.

[e]Based on the United Nations classification of 48 countries.

Source: OECD for MFN tariff, WTO tariff files for preferences, and U.N. Comtrade Statistics for trade.

TABLE 2. Tariff Peaks and Preferential Duty Rates in the Quad, 1999

Preferential trade agreements/GSP	Number of countries	Average preference rate (unweighted percent): Tariff peak products	All goods at HS-6
Canada			
United States	1	7.1	1.6
Australia	1	28.2	7.8
New Zealand	1	28.2	7.8
Mexico	1	15.9	3.1
Chile	1	12.2	2.4
Israel	1	11.8	2.5
Caribbean countries[a]	18	23.3	4.3
GSP-only beneficiaries[b]	108	28.2	6.2
Least developed countries[c]	47	22.8	4.4
Other countries (MFN rate)		(30.5)	(8.3)
European Union	15		
Eastern Europe and Middle East[d]	30	20.1	1.8
GSP-only beneficiaries[e]	42	19.8	3.6
Least developed ACP countries[f]	37	11.9	0.8
Other ACP countries[g]	32	12.4	0.9
Other least developed countries[h]	11	12.6	0.9
Other countries (MFN rate)[i]		(40.3)	(7.4)
Japan			
GSP-only beneficiaries[j]	127	22.7	2.3
Least developed countries[k]	42	19.0	1.7
Other countries (MFN rate)		(27.8)	(4.3)
United States			
Canada	1	0.6	0.1
Mexico	1	1.6	0.3
Israel	1	0.6	0.1
ANDEAN[l]	4	14	1.7
Caribbean community[m]	22	13.5	1.6
GSP-only beneficiaries[n]	80	16	2.4
Least developed countries[o]	38	14.4	1.8
Other countries (MFN rate)		(20.8)	(5.0)

[a]Includes 18 Caribbean countries or territories under Commonwealth Caribbean Countries Tariff.

[b]Excludes eight developing countries: Albania, Aruba, Bosnia and Herzegovina, Macedonia, Mongolia, Oman, Saudi Arabia, and the former Yugoslavia.

[c]Excludes Myanmar.

[d]Includes countries with reciprocal and nonreciprocal trade agreements with the European Union.

[e]Most developing countries in Latin America and Asia; excludes Hong Kong, Rep. of Korea, and Singapore (non-GSP nations).

[f]Includes 37 ACP and least developed countries under the Lomé Convention.

[g]Includes 32 ACP countries under the Lomé Convention but not under the group of least developed countries.

[h]Includes 11 least developed countries but not under ACP countries.

[i]Includes all industrial countries, Hong Kong, Korea, Singapore, and 14 transition countries.

[j]127 countries; excludes Albania, Bosnia, Estonia, Latvia, Lebanon, Lithuania, Macedonia, Moldova, Vietnam, and the former Yugoslavia.

[k]Excludes three least developed countries: Comoros, Djibouti, and Tuvalu. Three others (Rep. of Congo, Kiribati, and Zambia) are included in the GSP group.

[l]Includes Bolivia, Colombia, Ecuador, and Peru under the Andean Trade Preference Act.

[m]Based on 20 Caribbean countries under the Caribbean Basin Economic Recovery Act and the Bahamas and Nicaragua.

[n]Includes 80 developing countries or territories under the GSP scheme but excludes 29 other developing economies.

[o]Based on the United Nations 48 least developed countries but excludes 10 countries.

Source: World Trade Organization files.

and Pacific countries—mostly former colonies of European states—and free trade agreement preferences. In the case of the European Union, two different groups of LDCs are constructed for purposes of analysis: African, Caribbean, and Pacific countries and others. In the case of Canada, developing countries are grouped into those benefiting from LDC, GSP, or Caribbean preferences and Mexico and Chile, which benefit from free trade agreements. Finally, in the case of Japan, developing countries are split into GSP and LDC beneficiaries.

On average, the preferential schemes are quite generous. In the European Union, the average tariff faced by LDCs is less than 1 percent, compared with the 7.4 percent average most-favored-nation (MFN) tariff. GSP preferences in the European Union are less generous but still imply a margin of more than 50 percent. Japan and the United States grant a 50 percent preference margin under their GSP regime and an average 60 percent preference for LDCs. Canada gives a 25 percent preference to GSP countries and 45 percent to LDCs.

Preferences are much less generous for tariff peak products. Preference margins for GSP beneficiaries in Canada, Japan, and the United States on tariff peak items are only 8, 18, and 23 percent, respectively. For LDCs, the margins increase to 25 percent in Canada and 30 percent in the United States and Japan. The exception is the European Union, with a 50 percent margin for GSP beneficiaries and a 70 percent margin for LDCs in tariff peak items.

Thus, although existing preferential schemes grant significant preferences to developing countries, these are concentrated in products with low tariffs (between 0 and 15 percent) rather than on tariff peaks. In other words, preferential schemes offer little protection against tariff peaks in the Quad, except for the European Union. Hoekman, Ng, and Olarreaga (2001) and the U.N. Conference on Trade and Development (UNCTAD 2001) provide more detailed data on the average MFN import duties on tariff peak products and preference margins granted by the Quad to different groups of developing countries.

III. Tariff Peaks and LDC Exports

Simulation of the possible effects of duty-free access to the Quad requires data on global LDC exports of products that are subject to tariff peaks in these markets, because duty-free access to the Quad can result in redirection of exports to these markets and an increase in world prices (see section IV). During 1996–99, total LDC exports averaged $22.7 billion, of which $17 billion went to the Quad.[9] More than $5.5 billion of LDC exports to the world—25 percent of their total exports—were *potentially* affected by tariff peaks in Canada. Most of these affected exports are in apparel and clothing. More than 99 percent of LDC exports of apparel to the world are affected by an average tariff peak of 22 percent in

9. The simulations reported below use export data for 1996–99 as the base period. For a description of the product breakdown of LDC exports, see appendix B in Hoekman, Ng, and Olarreaga (2001). Data are drawn from the U.N. Comtrade database.

Canada. There is almost no preferential access for LDCs in these items (the preference margin is only 8 percent), implying that tariff peak elimination by Canada is likely to have a significant effect on LDC exports. Exports of other developing countries (non-LDCs) *potentially* affected by Canadian tariff peaks are also concentrated in apparel, with even smaller preference margins (around 3 percent). However, Mexico and Chile benefit from a 66 percent preference margin in these items under their respective free trade agreements with Canada, bringing the tariff they face to around 10 percent.

Similarly, more than $3 billion of LDC exports to the world, or 14 percent of all exports, are potentially affected by tariff peaks in the United States. Most LDC exports subject to tariff peaks in the United States are concentrated in apparel ($2.6 billion), facing an average tariff of 19 percent. Tobacco is another tariff peak item that is an important export item for developing countries. In the case of LDCs, more than 95 percent of their total exports of tobacco to the world potentially face a tariff peak in the United States of 63 percent (the MFN rate on these products averages 73 percent, but there is a 14 percent preference margin for LDCs).

The numbers are smaller in the case of Japan and the European Union, with tariff peaks in each market affecting some $500 and $800 million of LDC exports to the world, respectively. Although these numbers are small in absolute terms, the effect of peaks in these markets on specific LDCs may be quite large. For example, in the 1996–99 period, Djibouti, Kiribati, Somalia, and Tuvalu together exported less than $50 million to the world.

LDC exports affected by European Union tariff peaks are concentrated in meat and fish products, crustaceans, sugar, tobacco, and footwear. With the exception of meat, fish, or mollusk products and sugar, all of these exports benefit from full duty-free access into the European Union. In the case of preparations of meat, the 68 percent preference margin brought the tariff faced by LDC exporters down to around 10 percent. In the case of sugar, however, the preference margin granted to LDCs is quite small; their exports faced an average tariff of 29 percent.[10]

LDC exports to the world that are affected by Japanese tariff peaks include sugar, raw hides and skins, and footwear. Of these three products, the preference margin for sugar is only 5 percent, bringing the tariff faced by LDC exporters to 66 percent. By contrast, full preferences (duty-free access) are granted for raw hides and skins, whereas in the case of footwear an 80-percent preference margin applies to LDCs.

IV. A Simple Partial Equilibrium Model

To estimate the impact that the elimination of tariff peaks may have on LDC exporters, we use a simple partial equilibrium model. World markets are assumed

10. Note that the Everything But Arms initiative excludes sugar until 2009.

to be perfectly competitive and integrated in the sense that there is no further scope for arbitrage across countries. Products traded in world markets under the same six-digit Harmonized System classification are considered to be perfectly homogenous. Each six-digit Harmonized System product category represents only a small share of the economy, so that the effect on other product markets of changes in a particular category is negligible.

Import demand M_i for each Harmonized System six-digit product of country i = United States, European Union, Canada, Japan is given by:

$$M_i = A_i / [P_W (1 + T_i)]^E, \tag{1}$$

where E is the import demand elasticity (common to all countries in our simulations),[11] P_W is the world price; T_i is the MFN tariff in country i; and A_i is a demand parameter in country i. We assume throughout that tariffs are kept constant in the rest of the world. Rest-of-the-world import demand M_{ROW} is therefore

$$M_{ROW} = A_{ROW} / [P_W]^E. \tag{2}$$

Export supply $X_{j \to i}$ from country j to country i is given by

$$X_{j \to i} = B_j [P_W (1 + T_i \Pi_{i \to j})]^{\theta}, \tag{3}$$

where θ is the export supply elasticity (common to all countries), B_j is a supply parameter, and $\Pi_{i \to j}$ is the level of tariff preference granted by country i to exports from j. Thus, if $\Pi_{i \to j} = 0$, imports of i from j have to pay country i's MFN tariff. Similarly if $\Pi_{i \to j} = 2$, exports from j receive the domestic price in i.

The equilibrium world price, P_W^E, is obtained by solving for the world price in the world market-clearing condition, that is, the price for which

$$[\sum_k M_k - \sum_j X_j = 0] \Leftrightarrow P_W^E = [A_i / (1 + T_i)^E + A_{ROW}] / [\sum_j B_j (1 + T_i \Pi_{i \to j})^{\theta}]. \tag{4}$$

All demand and supply parameters are calibrated using U.N. trade data (value and quantities) at the six-digit level of the Harmonized System classification, MFN tariffs and preference margins of country i (see the appendix for data sources):[12]

$$B_j = X_j / [1 + T_i \Pi_{i \to j}]^{\theta}; \; A_{ROW} = M_{ROW} [P_W]^E; \; A_i = M_i [P_W (1 + T_i)]^E. \tag{5}$$

Using the calibrated parameters in expression (5) and replacing them in the right-hand side of equation (4) allows us to simulate the effect on world prices (and developing countries' export revenue) of changes in country i's tariff peaks on either a preferential or an MFN basis. Once the new world price is obtained, we substitute it into equations (1) and (3). Using the new tariff or preference margin, we then obtain the new import demand and export supply quantities for each country.

11. The six-digit Harmonized System import demand elasticities are derived from Stern, Francis, and Schumacher (1976) and Shiells, Stern, and Deardorff (1986).

12. Given that goods are perfectly substitutable, exports of j to the rest of the world need to receive the same price as exports to country i.

To determine the effect of a reduction in tariff peaks on world prices, we differentiate equation (4) with respect to t_i. This yields

$$\frac{\partial p_W^E}{\partial t_i} = \frac{1}{E+\theta}\left[\frac{A_{ROW} + A_i/(1+T_i)^E}{\sum_j B_j(1+T_i\Pi_{i\to j})^\theta}\right]^{1/(\theta+E)-1}\left(-E\frac{A_i/(1+T_i)^{E-1}}{\sum_j B_j(1+T_i\Pi_{i\to j})^\theta} - \left\{\sum_j B_j\theta(1+T_i\Pi_{i\to j})^{\theta-1}\Pi_{i\to j}\right\}\frac{A_{ROW} + A_i/(1+T_i)^E}{\left[\sum_j B_j(1+T_i\Pi_{i\to j})^\theta\right]^2}\right) < 0 \tag{6}$$

Thus, a reduction of country i's tariff peaks will necessarily lead to an increase in world prices. This does not necessarily lead to an increase of country j's export revenue, because some countries benefit from preferential access so that their export price is partly determined by the tariff. The export revenue of country j is given by

$$ER_j = P_W(1 + T_i\Pi_{i\to j})X_j = B_j(P_W[1 + T_i\Pi_{i\to j}])^{\theta+1}. \tag{7}$$

The change in export revenue following a change in country i's tariff is obtained by differentiating the right-hand side of equation (7) with respect to

$$[(\partial ER_m)/(\partial T_i)] = B_j(\theta + 1)(P_W[1 + T_i\Pi_{i\to j}])^\theta([(\partial P_W)/(\partial T_i)] + P_W\Pi_{i\to j}). \tag{8}$$

If country j has no preferential access to country i's market (that is, $\Pi_{i\to j} = 0$), a tariff cut will necessarily increase the export revenue of country j.[13] Similarly, if country j has full preferential access to country i's market (that is, $\Pi_{i\to j} = 1$), a tariff cut will reduce the export revenue of country i.[14] More generally, the export revenue of country j will increase following a tariff reduction in country i if

$$|(\partial P_W/\partial T_i)(T_i/P_W)| = T_i\Pi_{i\to j}. \tag{9}$$

That is, the elasticity of the world price with respect to the tariff must be smaller than the tariff faced by exporter j in country i. Thus, a crucial element for the analysis of the effects of tariff reductions in the Quad on the export revenue of developing countries is the degree of preferential access that developing countries initially enjoy in Quad markets.

Changes in export revenue are a function of current export levels. This implies that estimated export growth will be modest for countries that do not export much in the base period. This problem is attenuated by calibrating the model using global export supply and not bilateral export flows. However, calibration

13. To see this, note that the term in brackets on the right-hand side of equation (8) will have the same sign as the change in world prices (which is negative).

14. The term in brackets on the right-hand side of equation (8) will now necessarily be greater than zero because the elasticity of world price with respect to the tariff change in country i is smaller than the initial tariff in absolute value (unless we are in the presence of the Metzler paradox, that is, when a reduction in the tariff increases domestic prices).

ignores potential trade ("production") deflection that could generate large export growth in countries with important domestic production, but no exports in the base period. In principle, countries that are given large preferences have an incentive to redirect their whole domestic production to the export market. Not allowing for this is a shortcoming of the methodology that will tend to underestimate the potential for export growth.[15]

To determine the effect on world prices of an increase in preferential access for a subset of countries, it is necessary to also determine the impact on exporters in the rest of the world. The derivative of the world price, given in equilibrium by equation (4), with respect to the degree of preference ($\Pi_{i\to j}$) is clearly negative, suggesting that any increase in the tariff preferences that country i grants to country j will reduce world prices. This in turn will reduce the export revenue of exporters in the rest of the world.

Finally, we can measure the change in welfare in the exporting country associated with a change in preferential access or tariffs in the importing country by looking at the exporters' producer surplus. The change in welfare is

$$(10) \qquad \Delta W_j = \int_{P_j^{X,0}}^{P_j^{X,1}} B_j \left[P_j^X\right]^{\theta} dP_j^X = \left[(B_j)/(\theta+1)\right]\left[\left(P_j^{X,0}\right)^{\theta+1} - \left(P_j^{X,0}\right)^{\theta+1}\right],$$

where ΔW_j is the change in welfare in exporting country j and $P_j^{X,T}$ is the export price faced by exporters in country j at time T (where $T = 0$ for the pre-tariff-change period and $T = 1$ after the tariff change in the importing country).

V. Elimination of Tariff Peaks in the Quad

This section estimates how LDC exports would change if Quad members were to grant duty-free access to all LDC exports of tariff peak items. It also calculates the impact of a nondiscriminatory (MFN) reduction in all tariff peaks to a level of 5 percent (the Quad average). Each case groups developing countries according to the type of preference that they receive, distinguishing between LDCs, GSP beneficiaries, and free trade agreement partners.[16] In all simulations, we also take into account the existence of free trade agreements between industrial countries, although we do not report results for changes in exports of industrial countries.[17] The numbers that are reported are aggregations of all affected 6-digit tariff peak items.

Table 3 summarizes the expected changes in export revenue and welfare for LDCs and other developing countries if LDC exporters were to obtain full duty-

15. In other words, the export supply elasticity may be much larger than the 0.5 assumed in the analysis. To partially correct for this, we also run some simulations with an elasticity of export supply equal to 2 for products in which the preference margin is larger than 30 percent.

16. Stevens and Kennan (2000) have identified more than 30 tariff regimes in the European Union. We follow them in working with only the major aggregate categories/groups.

17. These are available from the authors on request.

TABLE 3. Impact of Duty-Free Access to Quad Markets for Least Developed Country Exporters
(millions of dollars)

Indicator	Canada	European Union	Japan	United States	Quad
Change in LDC exports	1,602	185	496	1,107	2,497
	(7.20)	(0.83)	(2.23)	(4.97)	(11.22)
Change in GSP beneficiary exports	−558	−100	−292	−387	−929
	(−0.03)	(−0.01)	(−0.02)	(−0.04)	(−0.05)
Change in all developing country exports	1,013	72	204	654	1,362
	(0.03)	(0.00)	(0.01)	(0.02)	(0.04)
Change in imports in the Quad	15	2	3	108	117
	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)
Change in LDC welfare	1,159	122	332	915	1,694
	(0.67)	(0.07)	(0.19)	(0.53)	(0.99)

Note: Figures in parentheses are percentages of values in the base year (1996–98 averages).
Source: Authors' calculations.

free access for tariff peak items in each Quad market. The European Everything But Arms initiative would increase LDC tariff peak exports by only $185 million (less than 1 percent of total LDC exports). This partly reflects the fact that LDCs already enjoy relatively good access to the European Union. If all the Quad members granted unrestricted access to LDC exports of tariff peak products, the increase in export revenue could be as large as $2.5 billion (or 11 percent).[18] Most of this would be due to better access to Canada and the United States.[19]

Exports from other developing countries would fall by as much as $1.1 billion.[20] This is equivalent to one-third of the total increase in LDC exports, but represents only 0.05 percent of total developing country exports. Thus, although trade diversion against other developing country sources would occur, the relatively small magnitude of LDC exports implies that in relative terms this would have negligible effects on the affected countries. Welfare changes for non-LDCs would be close to zero, whereas LDCs would see their welfare increase by 1 percent of GDP (table 3).

18. Note that the export changes across Quad markets cannot be simply added to obtain the total change due to all Quad members granting duty-free access to LDCs on tariff peak items. This is because in some cases tariff peaks on a six-digit item are found in more than one Quad market. It is therefore necessary to correct for double counting.

19. As a robustness check, we also performed some simulations for which the elasticity of export supply was increased to 2 whenever the preference margin was larger than 30 percent. This increased LDC exports by almost an extra $1 billion. Most of this increase was generated in Japan, where preferences are large on products that LDCs export in small quantities but where there is potential for supply expansion.

20. This is made up of a loss of $929 million incurred by GSP beneficiaries and a loss of $206 million incurred by other developing countries that enjoy preferential access to the Quad.

Total imports into the Quad associated with duty-free access for peak products would expand by only a modest $117 million (0.01 percent).[21] The reason for this very small increase is that imports from other sources (industrial and developing countries) would fall and tariff revenue would be transferred. This suggests there is not a compelling reason to be concerned with possible adjustment costs for domestic import-competing industries located in the Quad. However, it also implies that a MFN reduction in tariff peaks would result in a significant expansion of exports by OECD countries.

The distribution of changes in export revenue across products and countries would vary across Quad members (figure 1). In the case of the European Union, two-thirds of the $185 million increase in LDC export revenue would be concentrated in sugar and confectionery. The main beneficiaries would be Malawi, Zambia, and Mozambique, with 27, 19, and 15 percent of the total increase in LDC sugar exports, respectively. (Note, however, that the Everything But Arms initiative delays liberalization of LDC exports of sugar until 2009.) Some 10 percent of the total increase in exports to the European Union would occur in meat products. Because these are subject to phytosanitary standards that may be difficult for LDCs to satisfy, the estimated export increase might be too optimistic.[22] The main beneficiary in terms of the absolute increase in exports to the European Union would be Madagascar. Its exports would increase by $26 million or about 4 percent of total exports in 1999. In relative terms, the LDC that would gain the most from duty-free access for tariff peak items in the European Union would be the Maldives, with a 19 percent increase in exports ($14 million).[23]

In the case of Japan, most of the increase (90 percent) would be concentrated in sugar and confectionery. Malawi, Zambia, and Mozambique again are predicted to capture most of this increase. Bangladesh would benefit the most in absolute terms, with an export increase of $229 million (47 percent of the total increase in LDC exports to Japan). This represents around 5 percent of Bangladesh exports in 1999. In relative terms, the main beneficiaries would be Somalia (a 43 percent increase in exports or $13 million) and Cape Verde (a 23 percent increase in exports or $4.4 million).

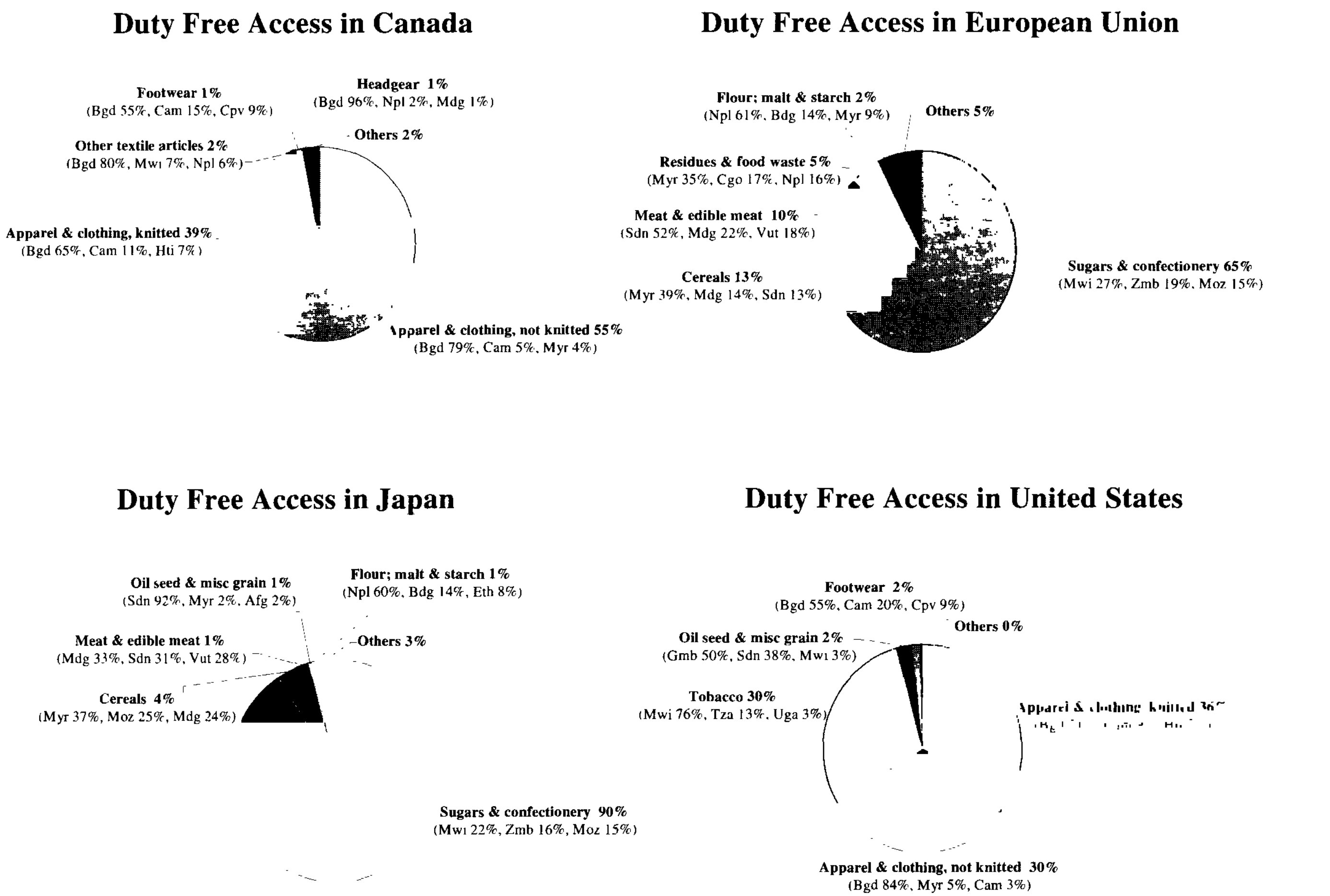
In the case of Canada and the United States, most of the expansion in exports would occur in apparel and clothing and footwear. The main beneficiary would be Bangladesh, with an increase of more than $1 billion in exports, more than

21. The increase in total imports in each Quad member is measured as the difference between actual imports and estimated imports at the new domestic price; that is, it excludes the increase in LDC imports that is simply explained by tariff revenue transfers.

22. The main beneficiary is predicted to be Sudan, which is unlikely to be able to export meat to the European Union due to the existence of foot and mouth disease. This is an example of the overestimation of changes in exports that can arise due to the assumption of product homogeneity across markets.

23. For more details in terms of the increase in exports by country associated with each Quad member granting duty-free access, see table 7 in Hoekman, Ng, and Olarreaga (2001).

FIGURE 1. Product and Country Decomposition of Changes in the Export Revenue of Least Developed Countries



20 percent of its total exports in 1999. In relative terms, Liberia, Haiti, Laos, and Cambodia would gain substantially from the elimination of tariff peaks in Canada, with export increases of more than 20 percent. In the U.S. market, tobacco is also an important tariff peak item. Elimination of tariffs would benefit such producers as Malawi, which would be expected to experience a 25 percent increase in exports.

In the aggregate, the losses associated with the displacement of exports from other developing countries would be small and not concentrated. In principle, it would be expected that the major losers from preferential elimination of tariff peaks for LDCs would be developing countries that currently benefit from preferences. These include African, Caribbean, and Pacific countries that are not LDCs and countries with free trade agreement status, such as Mexico. The African, Caribbean, and Pacific countries that do not benefit from the Everything But Arms initiative would lose only $21 million from its implementation. However, more than 90 percent of the contraction would occur in sugar exports. Because sugar is excluded from the initiative until 2009, the loss for the non-LDC African, Caribbean, and Pacific members would fall to just $1 million.[24] More than 60 percent of the potential loss for African, Caribbean, and Pacific countries would be concentrated in Mauritius, Fiji, Guyana, and Jamaica. None of these countries would lose more than 0.1 percent in terms of export revenue.

In the case of Canada and the United States, the effect on Mexico and the Caribbean would be negligible: a decline of $20 million on a base of total exports of more than $150 billion. Generally, there is not a single developing country for which the loss in export revenue would represent more than 0.7 percent of its total exports. If Canada granted duty-free access to LDCs for tariff peak products, Jamaica would incur the largest relative loss, a decline of 0.63 percent in its global exports.

Nondiscriminatory Reduction of Tariff Peaks

Preferential liberalization of trade is inferior to nondiscriminatory liberalization in welfare (efficiency) terms. The reason is that trade diversion can easily occur, whereby less efficient suppliers that are granted preferential access are able to force more efficient ones out of a market. Consumers then end up paying too much for the products concerned, with associated efficiency losses.

A reduction in all tariff peaks in the Quad to 5 percent, applied on a nondiscriminatory basis to all exporters, would eliminate all the gains incurred by LDCs under the preferential scenario (table 4). LDCs would not only need to compete with other developing and industrial countries in Quad markets but the value of their current preferential access under GSP or LDC preferential schemes would

24. Note that for the other two products for which implementation of duty-free access has been delayed, the loss in terms of export revenue for non-LDC African, Caribbean, and Pacific countries would be negligible: around $0.3 million for bananas and $0.2 million for rice.

TABLE 4. Impact of Reducing Quad Tariff Peaks to 5 Percent on a Most-Favored-Nation Basis
(millions of dollars)

Indicator	Canada	European Union	Japan	United States	Quad
Change in LDC exports	–116	–47	–94	128	–71
	(–0.52)	(–0.21)	(–0.42)	(0.57)	(–0.32)
Change in GSP beneficiary exports	1,512	797	–3,932	2,949	423
	(0.09)	(0.08)	(–0.22)	(0.33)	(0.02)
Change in all developing country exports	1,294	645	–4,126	2,659	–110
	(0.04)	(0.02)	(–0.14)	(0.09)	(0.00)
Change in imports in the Quad	1,223	628	826	5,862	7,343
	(0.64)	(0.08)	(0.26)	(0.66)	(0.34)
Change in LDC welfare	–78	–32	–62	86	–45
	(–0.05)	(–0.02)	(–0.04)	(0.05)	(–0.03)

Note: Figures in parentheses are percentages of values in the base year (1996–98 averages).
Source: Authors' calculations.

erode.[25] The same would be true for non-LDC developing country exporters. As a result, under an MFN scenario, aggregate exports of developing countries as a group would actually fall slightly. Total imports by Quad members would expand by some $7.3 billion, reflecting greater exports by OECD countries.

Two implications emerge from this analysis. First, the net gain to developing countries of a nondiscriminatory elimination of tariff peaks would be essentially zero. However, this would be due to an expected decline in exports to Japan; exports to the other Quad members would increase significantly, reflecting the importance of textiles and clothing for developing countries. However, LDCs would not see export expansion of tariff peak items in any Quad market. Second, a unilateral MFN elimination of all tariff peaks in the Quad could be difficult to realize in political terms because it would lead to a nonnegligible increase in import penetration in the Quad. Thus, an MFN reduction of tariff peak items would likely require a broader context that would allow for reciprocal concessions to be offered by countries that would see their exports expand. That is, MFN elimination would likely require a WTO round of trade negotiations.

Caveats

Although we have made a number of assumptions that limit the supply response to tariff peak elimination, the estimated gains from preferential access may nonetheless be too high.[26] Though expanding exports to a particular market by redi-

25. The decline in Japan would be concentrated in leather footwear, where the MFN tariff is 23–30 percent and GSP preferences were around 60 percent in the late 1990s. If this preference margin were eliminated, OECD exporters would increase their market share significantly.

26. We could argue that the static nature of the simulations underestimates the potential export gains for LDCs. Once we allow for investment (foreign direct investment), the supply response in LDCs to large tariff preference margins in the Quad may be much higher than that assumed in the simulations.

recting exports from other regions would not require an increase in total supply, it would require the establishment of strong business relationships and a good reputation as a supplier in the new market. This might limit the gains from these preferential initiatives.

The estimates might also be overly rosy in that they assume that access is truly free. In practice, any type of preference will be accompanied by rules of origin and may remain subject to the threat of contingent protection—antidumping, countervailing duties, and safeguard actions. These types of policy instruments can be used to make duty-free access irrelevant in practice. Examples abound of protectionist lobbying in Quad members to tighten GSP rules of origin to restrict the ability of beneficiaries to significantly expand exports (see Bovard 1991 for examples in the United States). Rules of origin are also costly to administer—the tariff equivalent of the associated red tape can be significant. Herin (1986) estimates that the ad valorem cost of fulfilling rules of origin in trade between the European Union and other European countries (that in principle benefited from free trade status) was high enough for some 25 percent of all trade to pay the MFN tariff rather than document origin. Similarly, Sapir (1997) shows that in 1994, only half of total European imports that could potentially benefit from the GSP entered under this preferential regime. The other half entered on an MFN basis, reflecting the combined effect of rules of origin and tariff quotas.

The WTO includes an Agreement on Rules of Origin that aims to foster the harmonization of the rules used by members. The agreement calls for a work program to be undertaken by a technical committee, in conjunction with the World Customs Organization, to develop a classification system regarding the changes in tariff subheadings based on the Harmonized System that constitute a substantial transformation (Hoekman and Kostecki 2001).[27] The harmonization program provides a potential solution to problems of rules of origin. The rules of origin are intended to be applied for nonpreferential commercial policy instruments—tariffs, import licensing, antidumping, and so forth—but there is no reason why they could not be applied to preferential trade as well.

The threat of antidumping and similar instruments of contingent protection can also make duty-free access redundant if there is a probability that once exports have expanded they will be targeted by such mechanisms. It is therefore

27. Rules of origin are intended to prevent trade deflection and to determine where a good originates for duty purposes when two or more countries are involved in the production of a product. The general rule is that the origin of a product is the one in which the last substantial transformation took place, that is, the country in which significant manufacturing or processing occurred most recently. Significant or substantial is defined as sufficient to give the product its essential character. Various criteria can be used to determine whether a substantial transformation occurred. These include a change in tariff heading—as a result of whatever processing was performed, the good is classified in another category of the Harmonized System—the use of specific processing operations that do (or do not) imply substantial transformation, a test based on the value of additional materials embodied in the transformed product, or the amount of value added in the last country where the good was transformed. See Hoekman and Kostecki (2001).

important that duty-free access schemes exempt LDCs from the application of antidumping and safeguard actions. Although this may be politically difficult to achieve, the small trade flows concerned should make such a promise relatively painless in practice.

Finally, it should be noted that the above analysis completely ignores the fact that trade barriers faced by developing countries include policies imposed by other developing countries. Almost 40 percent of developing country exports were imported by other developing countries in 1998, and increasingly this trade comprises manufactured products (Hertel and Martin 2000). The analysis also ignores the effect of own liberalization, which can be expected to be a major precondition for benefiting from duty-free access in the Quad.

VI. Concluding Remarks

Although average tariffs confronting LDCs in Quad markets are very low, tariff peaks have a disproportional effect on LDC exports. Goods that are subject to MFN tariffs of 15 percent or more account for 11 percent of LDC exports to the Quad, although these types of products represent only 4 percent of total Quad imports ($93 billion). Of this small amount, LDCs account for less than 4 percent of total Quad imports of tariff peak items—they are very small players.

Products that are subject to tariff peaks, especially in Canada and the United States, tend to benefit from only limited preferential access. The impact of tariff peaks is therefore disproportionately greater for LDCs. Tariff peak products tend to be heavily concentrated in agriculture (sugar, cereals, and meat) and in labor-intensive sectors, such as apparel and footwear.

The impact on LDC exports of tariff peak items following the Everything But Arms initiative is likely to be quite small given that preferences were already generous. The estimated increase in exports of tariff peak products is around $185 million, less than 1 percent of total LDC exports. Excluding sugar, rice, and bananas from the analysis, duty-free access in the European Union is worth only a modest $60 million increase in exports. However, if *all* Quad members were to grant duty-free access for tariff peak items, this would have a significant effect on LDC exports. The increase could be as large as an extra $2.5 billion of LDC exports, which represents an increase of 11 percent in their total exports to the world. This would constitute a major improvement in terms of export performance.

The impact of elimination of peaks for LDCs on domestic producers in the Quad would be very small. Total import demand in the Quad would increase by a negligible $117 million. Most of the increase in LDC exports would be explained by either displacement of exports from other sources, or tariff revenue transfers from Quad members as they grant preferential access to LDCs. There would be trade diversion: other developing countries would see their exports fall by as much as $1.135 billion. Although this would represent 45 percent of the total increase in LDC exports, it would be negligible in terms of other developing countries' global exports—around 0.05 percent. Moreover, developing countries as a group (including LDCs) would see their exports increase by over $1.3 billion, as LDC exports would expand more than other developing countries' trade contracts. The rest of the increase in LDC exports would be explained by displacement of exports from industrial countries or a decline in Quad tariff revenue.

The distribution of export increases across products and countries reflects differences in both the export bundle of LDCs and the tariff peaks in Quad countries. In terms of specific product categories and countries, the impact of abolishing tariff peaks for LDCs would be relatively concentrated. In the United States and Canada, most of the action would be in apparel. In the European Union and Japan, the action would be primarily in sugar and related products and cereals. In absolute terms, Bangladesh would be the big beneficiary, being the largest LDC exporter of apparel, footwear, and fish to the European Union, the United States, and Canada. Cambodia, Cape Verde, Haiti, Laos, Liberia, Malawi, Maldives, and Somalia would also benefit significantly, seeing their exports increase by 20 percent or more. Given that tariff peaks across Quad countries occur in different products and that LDC export bundles are quite diverse, if all Quad members were to eliminate tariff peaks, it would help ensure that a larger number of LDCs would benefit.

It is well known that protectionist trade regimes in industrial countries are not the most important factors constraining LDC export growth. Of greater importance are domestic distortions and institutional weaknesses that create high transactions costs and bias investment incentives (Ng and Yeats 1997; World Bank 2001). Elimination of tariff peaks would not solve the problem of the marginalization of LDCs in global trade. However, the Quad could offer to eliminate tariff peaks and thereby help offset (to some extent) the major domestic challenges and transactions costs that confront domestic entrepreneurs in LDCs. In the process, by mobilizing export-oriented groups that would benefit from improved access to the Quad, this action might help alter the domestic political economy forces that constrain the adoption and implementation of better policies.

In principle, nondiscriminatory liberalization is superior to granting preferential access in welfare terms. However, such an approach toward dealing with tariff peaks would not enhance the exports of LDCs. Any effort to reduce tariff peaks on a nondiscriminatory basis—which is the preferred option from a global efficiency perspective—should be complemented by efforts targeted at assisting poor countries to improve their capacity to use trade as part of a propoor growth strategy. Expansion of "aid for trade" should also be an element of preferential access schemes. It is generally recognized that market access without the ability to produce profitably for export is of limited value. There is a large complementary agenda that must be pursued to enhance the ability of many low-income countries to participate in the global economy.[28]

28. See Hoekman (2001) and World Bank (2001) for more detailed discussions and proposals.

Appendix: Data Sources

All trade data are from the U.N. Comtrade Database (value and unit prices). MFN tariff schedules for Quad members are from the OECD compendium of tariffs, 2000. Tariff preferences have been calculated using Quad members' tariff schedules reported in the WTO-IDB database and preference data provided by the WTO's Trade Policy Review division (when data were available at the eight- or ten-digit level, simple averages were taken). In instances where specific tariffs are applied, we used ad valorem equivalents calculated by the OECD and the WTO Trade Policy Review division. Elasticities of import demand are assumed to be equal across countries and are constructed using data reported in Shiells, Stern, and Deardorff (1986) and Stern, Francis, and Schumacher (1976). (An Excel file is available from the authors.) Export supply elasticities are also assumed constant across countries, and, due to the lack of information at this level of disaggregation, we set its value to 0.5 (alternatively, we provide estimates with the elasticity of export supply set equal to 0 in Hoekman et al. [2001]).

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Imported Machinery for Export Competitiveness

Ashoka Mody and Kamil Yilmaz

This article analyzes the relationship between export competitiveness and investment in machinery, allowing for imperfect substitution between domestically produced and imported machinery. A translog export price function is estimated for developed, export-oriented developing, and import-substituting developing economies in a panel data setting. Between 1967 and 1990 imported machinery helped lower export prices for export-oriented developing economies. Moreover, throughout the period imported machinery was not a substitute for domestic machinery. Import-substituting developing economies were unable to harness imported machinery to reduce costs early in the period, but from about the early 1980s, with the opening of their trade regimes, they were able to benefit from the cost-reducing effect. The results imply that innovative effort based on imported technologies can be a precursor to the development of domestic innovation capabilities.

In this article we build on two recent lines of research: investment in equipment as a source of economic growth and imported goods as conduits for the international diffusion of technology. We combine these two themes to assess the effectiveness of imported machinery in increasing export competitiveness and thus in stimulating growth.[1]

Underlining the importance of machinery in the development process, De Long and Summers (1991, 1992a, 1992b, 1993) find strong empirical support for a causal relationship between equipment investment and economic growth in a cross-section of developing and developed countries. In particular, they find that a 1% increase in the share of equipment investment in gross domestic product

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1. Several studies suggest that greater trade is associated with faster productivity growth (for example, Pack and Page 1994, Srinivasan 1995, 1999). However, Rodriguez and Rodrik (1999) and Rodrik (1999) are skeptical of such results when they are based on cross-sectional growth regressions. Srinivasan and Bhagwati (1999) express general concern about cross-sectional growth analyses and conclude that "nuanced and in-depth studies" of individual countries over a period of time provide the clearest evidence in favor of the beneficial effects of greater trade orientation.

(GDP) raises the GDP growth rate by 0.34%.[2] They infer that the domestic research and development (R&D) and learning activities associated with producing and installing equipment create generalized benefits for the economy.

The De Long–Summers results are also consistent with the possibility that the foreign knowledge embodied in imported equipment is of significant value to the economy buying the equipment. Studying the spillovers of knowledge across national boundaries, Coe and Helpman (1995) and Engelbrecht (1997) find that such international spillovers are mediated through imported goods: the greater the imports, the greater the benefit from the stock of foreign knowledge. Engelbrecht notes that these two articles do not distinguish between different types of imports; such a distinction is likely to be important because consumer goods, intermediate inputs, and equipment are likely to convey spillovers to differing degrees. Extending these earlier studies, Coe, Helpman, and Hoffmaister (1997) find that imported capital goods are critical conduits of international knowledge.

In this article we examine empirically the different efficiency of domestically produced and imported machinery. With information freely accessible to all there would be no difference in efficiency. In practice, however, information is not freely available. Even in the absence of formal protection of intellectual property rights, domestic producers can be at a disadvantage relative to international producers because knowledge is tacit (for a review of the economic and management literature on tacitness see, for example, Mody 1989). As a consequence, domestic and imported machinery trigger different forms of learning in the domestic economy. Domestic production and installation of machinery may in some instances be associated with considerable innovative activity in a developing economy. More often, however, the domestic production of machines is associated with adaptive R&D—that is, tailoring foreign machinery to local requirements and upgrading domestic equipment of earlier vintages.

In contrast, imported machinery is bundled with "knowledge" in various forms: blueprints, installation support, quality control software, and services of trained engineers and supervisors. Absorbing knowledge through these means is less glamorous than developing or even adapting machines. However, because imported machinery forms a more comprehensive package, it can potentially lead to greater efficiency in the short run and stronger absorptive capacity in the long run. Imported machinery will also be more efficient because it is typically of newer vintage than domestically produced machinery.[3]

2. In the traditional neoclassical model an increase in the investment rate raises output but has no long-run effect on growth rates. The endogenous growth literature identifies conditions under which increased investment has external effects and thus raises growth rates. De Long and Summers go further and find evidence that the external effects are strongest when the investment is in machinery rather than in buildings and structures.

3. Other mechanisms for knowledge transmission can also be important, including the provision of technical and marketing support by foreign buyers of exported goods in the context of long-term relationships (Westphal, Rhee, and Pursell 1981, Egan and Mody 1992, Mody and Yilmaz 1997).

We use the country's trade regime to proxy the incentives to deploy knowledge (table 1). Not all economies (or firms) are able to take advantage of bundled software and training or the greater efficiency built into the new vintages of imported machinery. In countries with a strong export orientation, however, firms are likely to have strong incentives—driven by the need to stay competitive—to exploit the knowledge flows associated with imported machinery. Import substitution was based on the premise that, with temporary protection, domestic producers would have the incentive to tap into and internalize internationally available knowledge. The extent to which this actually occurred, however, is an empirical issue. Incentives in more protected "import-substituting" economies were likely to be weaker because of the relatively small size of domestic markets and the less demanding domestic users of that machinery (see Srinivasan and Bhagwati 1999 for a review of how incentives are blunted in an import substituting regime). Thus, although alternative explanations are possible, the results of our analysis are consistent with the proposition that import-substituting regimes create weaker incentives to invest in technological improvements that can help expand their export presence in international markets.

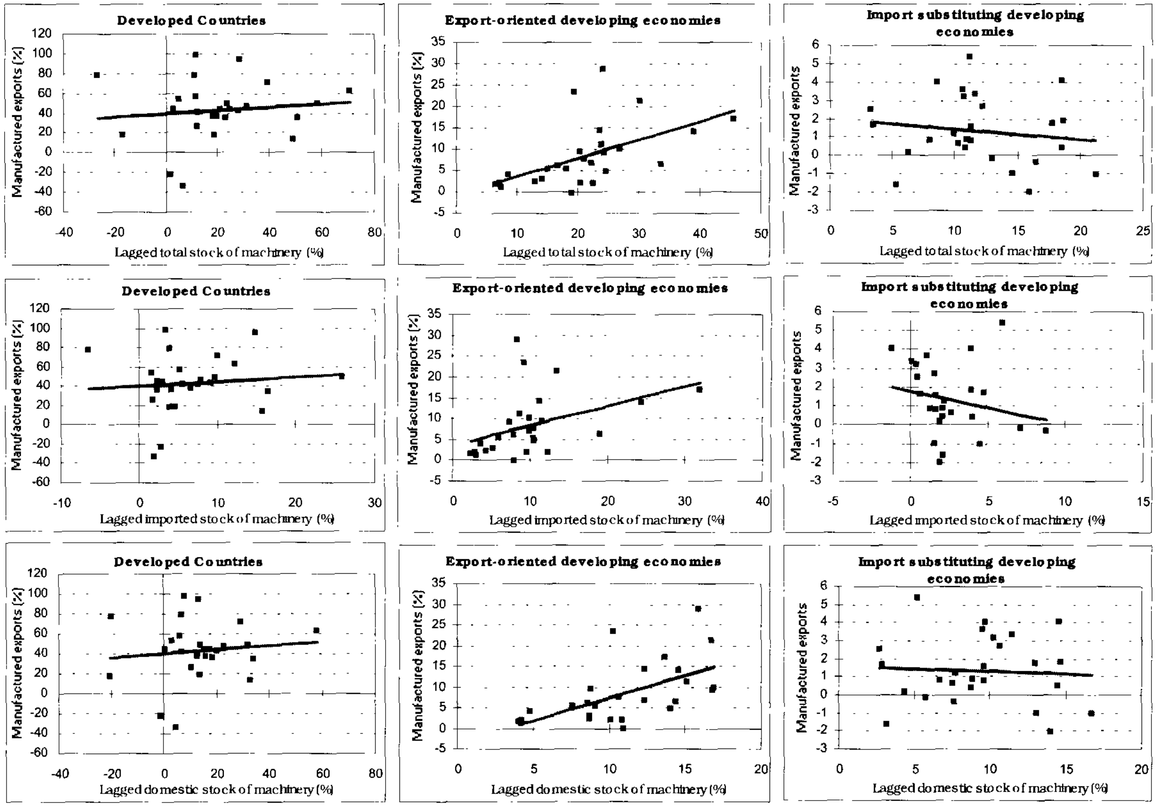
Plotting the change in the volume of exports against the change in the capital stock in the previous year, we find that in export-oriented developing economies an increase in exports is strongly associated with an increase in the stock of imported machinery (figure 1). A positive relation also exists between export growth and an increase in the stock of domestic machinery. A similar set of relationships is found for developed countries. In contrast, for import-substituting developing economies imported equipment and export growth are negatively related.

Our empirical analysis focuses on the price of exports rather than on their volume, specifying a link from imported machinery to reduced costs and prices, which in turn leads to greater exports. We analyze the relationship between machinery investment and export competitiveness using a model of imperfect competition in international markets, allowing for imperfect substitutability between domestically produced and imported machinery. We specify an export

TABLE 1. Trade Regimes and the Effects of Imported and Domestic Machinery

	Trade regime	
	Export-oriented	Import substituting
Imported machinery	Has access to the pool of international knowledge and the incentives to exploit it.	Can access the pool of international knowledge, but small domestic markets may blunt incentives.
Domestic machinery	With strong incentives, can overcome the limits of the domestic knowledge pool as domestic capacity improves.	Both the knowledge pool and incentives may be limited.

FIGURE 1. Growth of Manufactured Exports on Growth of Total, Imported, and Domestic Machinery Stock



Source: Authors' calculations, based on data from U.N. Commodity Trade (Comtrade) database.

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price function based on the demand for exports and the costs of producing the exported goods. We use a short-run cost function with variable labor and materials costs and fixed stocks of imported and domestically produced machinery. Larger stocks of capital are expected to lower short-run production costs and thus export prices. Higher productivity of imported machinery would be reflected in the greater cost reduction than can be achieved using domestic machinery.

We estimate the export price equation for developed countries, for developing economies, and, within the second group, for export-oriented and import-substituting economies. For each group we use the fixed-effects procedure on panel data. But first we use the t-bar test recently developed by Im, Pesaran, and Shin (1996) to identify (in a panel data context) the presence of stochastic trends in export price and explanatory variables. We find that the variables do have stochastic trends (unit roots). Thus, because of the potential colinearity in the movement of variables of interest, we estimate the export price function in first differences.

In the empirical application, to allow for lagged effects, we distinguish between the effects due to the past year's new investment and those due to the stock of capital at the start of the previous year. For the entire sample period, 1967–90, the flow of new imported equipment in the previous year is seen to be associated with a decline in export prices in developed and export-oriented developing economies, but not in import-substituting developing economies. The results also show that the relation between imported equipment and export prices has evolved over time. Throughout the period, we find that domestic machinery is not a substitute for imported machinery.

I. A Model of Imperfect Competition in Export Markets

In setting up the model we are guided by the following intuition: the significance that some developing countries, especially the newly industrializing countries of East Asia, attached to investment in machinery was not accidental. Instead, it was dictated by the adoption of an export-oriented strategy and the resulting discipline of international competition. To maintain market presence, exporters had to reduce production costs continually or enter into the production of higher-quality goods. Both strategies required substantial investment in new vintages of machinery and equipment. Initially, domestic machinery had lower productivity, so the scope for substituting domestic for imported machinery was small. Over time the more advanced developing countries progressed to the point where their technological capability to produce machinery could compete with imports from developed countries.

This intuition can be tested by estimating a cost function that includes machinery stock as an explanatory variable. However, data on production costs are difficult to obtain. For this reason we estimate an export price function based on both demand and cost function parameters. In a model of imperfect competition, manufacturers arrive at their export price given demand and cost conditions. Though demand depends on competitors' prices and on incomes in target markets, production costs depend on input prices, output levels, and other variables that shift the cost function, such as the stocks of imported and domestic machinery.

We assume that production for domestic and export markets are two independent decisions and focus on exports.[4] Firms produce export goods through a homothetic production function using two variable inputs, labor and materials (including fuel, electricity, and raw materials) and a quasi-fixed input, the capital stock. Firms are assumed to be price takers in input markets. Consequently, the short-run cost function can be separated into variable input prices on the one hand and the quasi-fixed input and output on the other.

We assume that each firm exports a differentiated product and chooses its export price to maximize its profit at a point in time, given the demand curve for its product and the cost of production.[5] When the second-order condition for maximizing profit is satisfied, it is possible to solve for the profit maximizing price, by inverting the first-order condition. The profit maximizing price is a function of all variables that enter the cost function—the wage rate (w), price of materials (p_m), and capital stock (K)—plus variables that shift the demand function, the competitors' average price (p_c) and world income (Y). The export price may also be a function of the exchange rate (e), as will be discussed.

$$p = p(p_c, Y, w, p_m, e, K) \tag{1}$$

The elasticities of the export price with respect to variable input prices, the prices of competing products, and the capital stock depend on the parameters of the cost and demand functions. When the second-order condition for maximizing profit is satisfied, a positive elasticity of the marginal cost with respect to input prices is sufficient to generate a positive elasticity of the export price with respect to input prices. In other words, the exporter will increase its price following an increase in input prices.

With the second-order condition satisfied, a decreasing marginal cost with an increasing machinery stock is both a necessary and a sufficient condition for the

4. For a similar assumption and empirical implementation see Feenstra (1989). If the marginal costs of production for the domestic market and export markets are not flat, influences in one market will influence the other. Essentially, an omitted variable bias would arise where the omitted variables refer to demand influences in the domestic economy. If domestic demand were to shift exogenously, the marginal costs of production would change, leading to a change in prices charged in both the domestic and the international markets. We believe that these exogenous shifts would be reflected in the prices of domestic inputs (wages and materials costs). A bias may still remain, however, though the direction of it is unclear. If increased domestic activity leads to more investment but also higher marginal costs, a larger stock of private capital would be associated with higher export prices—the opposite of the relationship that we are hypothesizing.

5. Because the analysis is restricted to the cost-reducing effect of the technology embodied in existing machinery, the model is static and does not incorporate investment demand for domestic and imported machinery. Analytically it is not difficult to incorporate the demand for machinery through a dynamic model. However, because of a lack of data on production costs and the rental price of capital stock, it is not possible to estimate factor demand functions of the long-run model.

price to be a decreasing function of the machinery stock. Consequently, if the estimated price elasticity with respect to the machinery stock is negative, it follows that the technology embodied in new machinery has a cost-reducing effect.

For the purpose of empirical estimation and following Mann (1986, 1989), we simplify the demand function by substituting the world price (p_w) for the competitors' price (p_c) and world income (Y). The world price variable reflects the influence of the pricing decisions of all competitors and of changes in world income. Thus, using the reduced-form price equation, we analyze the elasticity of the export price with respect to the world price, two input prices, and the two kinds of machinery stock.[6]

In the empirical analysis we assume that the export price decision is best summarized by the translog price function

$$\text{(2)} \qquad \log p = \lambda + \sum_i \beta_i \log X_i + 0.5 \sum_i \psi_i (\log X_i)^2, + \sum_i \sum_{j>i} \psi_{i,j} (\log X_i \log X_j),$$

where $X_i = p_w, w, p_m, e, I^m_{-1}, I^d_{-1}, K^m_{-2}, K^d_{-2}$. I^d is the investment flow for domestic capital goods, I^m is the investment flow for imported capital goods, and K^d and K^m are the corresponding stocks. A variable with the subscript −1 is lagged one period; the subscript −2 implies a two-period lag. By considering the past year's investment and the stock prior to that, we are able to obtain some sense of the lags with which the effects operate.[7]

Previous studies analyzing export price behavior under imperfect competition have noted that exchange rates often exercise an independent influence on the price of traded goods. In other words, even if all variables on both sides of the equation are measured in the same currency, exchange rate movements seem to have a significant effect on the price of exports (Feenstra 1989, Ohno 1989, Mann 1986). By representing the input prices in local currency terms and including the exchange rate as a separate variable, we allow for the possibility that changes in exchange rates are not perfectly passed through to export prices. Our primary results remain unchanged if we instead measure the input prices in dollars and drop the exchange rate variable.

II. Empirical Specification and the Data

We estimate the export price equation for a cross-section of 14 developed countries and 25 developing economies. (For the definitions of variables and the data sources, see appendix table A-1; for the descriptive statistics, see appendix table A-2.) Because we derive the model based on profit maximizing assumptions for an individual firm, it would be best to use firm- or industry-level data to estimate

6. Local currency wages and the price of material inputs were obtained by dividing the corresponding variables denominated in U.S. dollars by the annual average exchange rate.

7. Of course, we are not decomposing the stock of capital in a strict sense, with this year's stock equal to new investment plus the previous stock. That simple identity does not carry forward when we take logs.

the price function in equation (2). However, data constraints preclude that route. Though data on export prices, input prices, and investment can be found for some manufacturing subsectors in some countries, it is not possible to obtain data on domestic and imported components of investment by each industry. We are therefore forced to aggregate all manufactured exports from a country.

Aggregation can be justified by assuming either a representative firm (as in Feenstra 1989, Ohno 1989) or a translog aggregate production technology for manufacturing exporters (Pindyck and Rotemberg 1983). Aggregation presents its own problems, however. The higher the level of aggregation, the more difficult it becomes to obtain price indexes that reflect firm-level pricing decisions. An aggregate price measure incorporates changes in the composition of the commodity basket as well as in the market price of each commodity in the basket. Is this a problem for our proposed empirical analysis? No, because our focus is on cost reduction. To the extent that changes in the composition of exports from one year to the next are important, the cost-reduction effect will be blurred. Indeed, if products were moving up the quality ladder, we would expect to find no cost-reduction effect. Thus a finding of cost reduction despite that possibility provides somewhat greater confidence in our results.

Because our sample of developing economies represents substantially different development strategies, we divide the economies into two groups, export-oriented and import-substituting, based on the World Bank classification (World Bank 1986; see also Balasubramanyam, Salisu, and Sapsford 1996). Between 1967–73 and 1973–85 no major shift occurred in the outward orientation of the developing economies in our sample (appendix table A-3). However, although the economies remained differentiated in their broad policy stance, their trade policy regimes did not remain fixed. The reduction in trade barriers continued apace, with several of the import-substituting economies adopting more export-oriented policies in the 1980s. Thus the differences in policy regimes narrowed.

Before estimating the export price function, we test for nonstationarity of the variables using the t-bar statistic proposed by Im, Pesaran, and Shin (1996) for heterogeneous panels. This is a well-known crucial first step in time-series models. When a time-series equation contains a nonstationary variable, the results based on this estimation will be spurious. Im, Pesaran, and Shin (1996) recently extended the stationarity tests to cross-section, time-series models.

The test procedure is simple. It extends the widely used augmented Dickey-Fuller (ADF) test to a panel data framework and allows for heterogeneity across groups in the panel. First, the average ADF unit root test statistic for the panel is obtained as the mean of individual ADF unit root statistics. Next, the expected value and the standard error of the average ADF test statistic under the null hypothesis of a unit root are obtained through Monte Carlo simulation. The t-bar statistic is calculated as the average ADF test statistic minus its expected value divided by its standard error. Im, Pesaran, and Shin (1996) show that under the null hypothesis of a unit root, the t-bar statistic has a standard normal distribution for a sufficiently large number of countries, N, and number of periods, T, while

$\sqrt{N/T}$ goes to zero. Using the Monte Carlo method, they show that the t-bar test has more power than ADF tests applied separately to each individual in the panel.

Based on the results of the Im-Pesaran-Shin test, we cannot reject the null hypothesis of a unit root for all variables of the price function for all country groups (appendix table A-4). Consequently, estimating the export price function in levels (equation [2]) would generate spurious results. Next, we test for unit roots in the first-differenced variables and reject nonstationarity. This allows us to estimate the equation in first differences.

III. Empirical Results

We estimate the first-differenced export price equation using the fixed-effects procedure. This amounts to assuming that countries do differ in terms of the trend coefficient, which could be interpreted as disembodied technical change.[8]

For all country groups in our analysis we estimate the translog parameters using the data for 1967–90 (table 2). The specification test for functional form indicates that the translog function provides a better approximation of the export price decision than does the Cobb-Douglas function.

However, the parameters of the translog function cannot be interpreted directly. Instead, one needs to derive the elasticity estimates of the export price function with respect to input prices, the exchange rate, and imported and domestic machinery using the underlying parameters of the translog function. These elasticities take the following form:

$$E_i = \beta_i + \psi_i \overline{\log X_i} + \sum_{j \neq i} \psi_{i,j} \overline{\log X_j}, \qquad (3)$$

where $X_i = p_w, w, p_m, e, I^m_{-1}, I^d_{-1}, K^m_{-2}, K^d_{-2}$ and a bar over a variable denotes its average value for the country group throughout the sample period.

The standard error of each elasticity is estimated using the δ method (for a more detailed treatment see Rao 1973, p. 388–90). One can write the elasticities in the following matrix notation: $E = Z\Psi$, where Ψ is the 44×1 vector of translog function parameters and Z is an 8×44 matrix of zeros, ones, and the means of log variables, as given in equation (3). Using this matrix notation, we obtain the variance-covariance matrix of the elasticity matrix E, $\Sigma_E = Z\Sigma_\Psi Z'$, where Σ_Ψ is the variance-covariance matrix of the parameter estimates, excluding the intercept.

In the rest of the article the results focus on the elasticities. We present the results in two parts. First, we discuss the results for the full sample period, 1967–90, the period for which we have complete data for the variables of interest.[9]

8. Alternatively, one could assume that individual effects occur randomly rather than being fixed. This implies that the individual effect is part of the random disturbance rather than the constant term specific to a country. However, this assumption is not justified here because we did not sample countries randomly.

9. The binding data constraint is imposed by the use of machinery investment data from the Penn World Tables, which end in 1990. (See Heston and Summers, 1991.)

TABLE 2. Export Price Equation: Translog Estimates, 1967–90

Parameter	Developed Countries		Developing economies		Export-oriented developing economies		Import-substituting developing economies	
β_{Pw}	2.418	(1.62)	–2.595	(1.39)*	–1.951	(2.31)	–4.187	(2.02)**
β_{w}	–0.517	(0.89)	–0.003	(0.52)	–1.536	(1.23)	0.286	(0.78)
β_{Pm}	1.608	(1.15)	1.878	(0.35)***	2.451	(0.59)***	1.798	(0.56)***
β_{e}	–0.340	(1.71)	–2.140	(0.64)***	–1.221	(1.52)	–2.170	(0.97)**
β_{Im}	–1.699	(0.73)**	1.156	(0.35)***	0.990	(0.78)	1.936	(0.49)***
β_{Id}	–0.124	(0.46)	0.474	(0.32)	0.676	(0.45)	0.414	(0.48)
β_{Km}	2.782	(1.33)**	–1.446	(0.79)*	–2.389	(1.47)	–3.146	(1.43)**
β_{Kd}	0.428	(0.68)	–0.652	(0.98)	0.696	(1.53)	0.554	(1.57)
Ψ_{Pw}	–0.217	(0.27)	0.319	(0.31)	0.352	(0.46)	0.589	(0.47)
Ψ_{w}	–0.138	(0.14)	–0.116	(0.05)**	–0.255	(0.13)**	–0.098	(0.08)
Ψ_{Pm}	0.226	(0.18)	0.175	(0.04)	0.259	(0.05)***	0.161	(0.07)**
Ψ_{e}	0.392	(0.31)	0.135	(0.07)*	–0.039	(0.20)	0.203	(0.11)**
Ψ_{Im}	0.072	(0.08)	0.026	(0.05)	–0.059	(0.06)	0.201	(0.09)**
Ψ_{Id}	–0.010	(0.02)	0.011	(0.01)	0.004	(0.02)	0.010	(0.02)
Ψ_{Km}	–0.176	(0.18)	0.099	(0.12)	0.037	(0.20)	0.593	(0.25)**
Ψ_{Kd}	–0.060	(0.07)	0.076	(0.08)	–0.060	(0.12)	0.234	(0.13)*
$\Psi_{Pw,w}$	–0.027	(0.17)	0.037	(0.10)	0.161	(0.17)	0.032	(0.16)
$\Psi\eth_{Pw,Pm}$	–0.222	(0.18)	–0.393	(0.08)***	–0.502	(0.12)***	–0.405	(0.14)***
$\Psi_{Pw,e}$	0.094	(0.26)	0.423	(0.14)***	0.418	(0.24)*	0.424	(0.21)**
$\Psi_{Pw,Im}$	0.163	(0.10)*	–0.172	(0.07)**	–0.167	(0.11)	–0.306	(0.10)***
$\Psi_{Pw,Id}$	0.118	(0.07)*	–0.044	(0.06)	0.018	(0.07)	–0.011	(0.10)
$\Psi_{Pw,Km}$	–0.012	(0.16)	0.295	(0.11)***	0.170	(0.19)	0.515	(0.16)***
$\Psi_{Pw,Kd}$	–0.259	(0.08)***	0.077	(0.08)	0.061	(0.17)	–0.009	(0.12)
$\Psi_{w,Pm}$	0.199	(0.09)**	0.033	(0.03)	–0.011	(0.06)	0.071	(0.04)*
$\Psi_{w,e}$	–0.138	(0.13)	0.084	(0.05)*	0.284	(0.15)**	0.024	(0.08)
$\Psi_{w,Im}$	–0.066	(0.06)	0.085	(0.03)***	0.118	(0.06)**	0.049	(0.06)
$\Psi_{w,Id}$	–0.058	(0.05)	–0.028	(0.03)	–0.031	(0.05)	0.000	(0.05)
$\Psi_{w,Km}$	0.095	(0.11)	–0.088	(0.04)**	–0.097	(0.10)	–0.037	(0.10)
$\Psi_{w,Kd}$	0.117	(0.08)	0.026	(0.04)	0.113	(0.13)	–0.052	(0.09)
$\Psi_{Pm,e}$	–0.302	(0.23)	–0.216	(0.05)***	–0.255	(0.08)***	–0.231	(0.08)***

$\Psi_{Pm,Im}$	–0.053	(0.09)	0.009	(0.03)	0.036	(0.03)	0.018	(0.04)
$\Psi_{Pm,Id}$	–0.062	(0.07)	0.049	(0.02)**	0.052	(0.03)*	0.032	(0.03)
$\Psi_{Pm,Km}$	–0.109	(0.11)	0.005	(0.03)	–0.064	(0.05)	0.018	(0.05)
$\Psi_{Pm,Kd}$	0.111	(0.08)	–0.056	(0.02)**	–0.026	(0.04)	–0.050	(0.03)
$\Psi_{e,Im}$	0.249	(0.11)**	–0.100	(0.03)***	–0.164	(0.06)***	–0.059	(0.06)
$\Psi_{e,Id}$	0.083	(0.07)	–0.030	(0.03)	–0.033	(0.05)	–0.043	(0.06)
$\Psi_{e,Km}$	–0.209	(0.15)	0.069	(0.05)	0.172	(0.11)	0.010	(0.09)
$\Psi_{e,Kd}$	–0.122	(0.09)	0.055	(0.04)	–0.085	(0.13)	0.098	(0.08)
$\Psi_{Im,Id}$	0.013	(0.04)	0.012	(0.02)	–0.012	(0.03)	0.003	(0.04)
$\Psi_{Im,Km}$	–0.009	(0.09)	–0.049	(0.06)	0.109	(0.09)	–0.184	(0.09)
$\Psi_{Im,Kd}$	–0.027	(0.05)	–0.027	(0.03)	–0.081	(0.06)	–0.041	(0.06)
$\Psi_{Id,Km}$	–0.056	(0.05)	–0.013	(0.02)	–0.011	(0.04)	–0.032	(0.04)
$\Psi_{Id,Kd}$	0.029	(0.03)	–0.037	(0.02)*	–0.065	(0.04)	–0.021	(0.03)
$\Psi_{Km,Kd}$	0.074	(0.09)	–0.021	(0.08)	0.058	(0.12)	–0.287	(0.17)*
Adjusted R^2		0.87		0.35		0.49		0.22
Durbin-Watson statistic		1.95		2.21		2.22		2.21
Sum of squared residuals		0.332		5.044		1.365		3.397
Degrees of freedom	268		502		218		240	
H1	87.9	[<0.001]	2.13	[0.15]	0.65	[0.42]	0.29	[0.59]
H2	2.81	[0.20]	0.03	[0.87]	1.30	[0.26]	–0.15	[0.70]
H3	34.9	[<0.001]	18.4	[<0.001]	15.5	[<0.001]	12.9	[<0.001]
H4	43.1	[<0.001]	11.9	[0.98]	6.4	[0.85]	8.4	[0.76]
H5	154.7	[<0.001]	96.2	[<0.001]	101.7	[<0.001]	69.2	[0.001]

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: Figures in parentheses are heteroskedasticity-consistent standard errors (White 1980). Figures in square brackets are the marginal significance levels for the corresponding hypothesis.

H1: World price elasticity is equal to one (joint test for market power and economies of scale).
H2: Davidson and MacKinnon *J*-test: imported and domestic machinery are imperfect substitutes.
H3: Davidson and MacKinnon *J*-test: imported and domestic machinery are perfect substitutes.
H4: Country fixed effects do not differ from each other.
H5: Cobb-Douglas and translog functional forms do not differ.

Source: Authors' calculations (see appendix table A-1 for data sources).

Next, to undertake a more detailed analysis of the data, we repeat the estimations of the translog function for subsample windows, dropping one observation from the beginning of the sample period each time.

Full Sample Period: 1967–90

World price elasticity is high when own-price elasticity of demand is high, when there are significant diseconomies of scale in production, or when both conditions exist. Indeed, world price elasticity approaches one as own-price elasticity approaches infinity, that is, when the demand curve for the country's products is infinitely elastic. As expected, the world price elasticity estimate is lowest for developed countries (0.36), which face the least elastic demand curve and where diseconomies of scale are likely to be weakest (table 3). The test result supports the hypothesis that world price elasticity differs significantly from that for developed countries. World price elasticity for developing economies is 0.94, quite close to 1. World price elasticity for export-oriented economies is about the same as that for import-substituting economies and in both cases does not differ statistically from one.

The lower wage and material price elasticities for developing economies are consistent with their price-taking role in the world market. A price-taking firm cannot increase its prices to fully reflect increases in unit costs. In contrast, for a firm with market power, which can influence the export price of its products, wage and material price elasticity would differ significantly from zero. Wage elasticity is highest for developed countries, at 0.24. Wage elasticity for developing economies is 0.02 and does not differ significantly from zero. This result is driven mainly by import-substituting developing economies. Though their wage elasticity is –0.07 and not significantly different from 0, wage elasticity for export-oriented economies is 0.11 and statistically significant. Mate-

TABLE 3. Export Price Equation: Elasticity Estimates, 1967–90

Elasticity	Developed countries		Developing economies		Export-oriented developing economies		Import-substituting developing economies	
E_{P_w}	0.355***	(0.069)	0.942***	(0.079)	0.923***	(0.095)	0.925***	(0.139)
E_w	0.240***	(0.054)	0.020	(0.030)	0.112***	(0.041)	–0.070	(0.066)
E_{P_m}	0.187***	(0.032)	0.062***	(0.013)	0.062***	(0.016)	0.056**	(0.022)
E_e	–0.712***	(0.050)	–0.092***	(0.032)	–0.213***	(0.052)	–0.022	(0.056)
E_I^m	–0.052***	(0.019)	–0.024	(0.022)	–0.072**	(0.033)	0.001	(0.010)
E_I^d	–0.016	(0.998)	–0.041	(0.480)	–0.042	(0.860)	–0.058	(0.760)
E_K^m	0.097	(0.115)	–0.017	(0.084)	–0.029	(0.130)	0.068	(0.185)
E_K^d	–0.078	(0.089)	–0.170	(0.120)	–0.171	(0.150)	–0.188	(0.220)

**Significant at the 5 percent level.
***Significant at the 1 percent level.
Note: Figures in parentheses are heteroskedasticity-consistent standard errors (White 1980).
Source: Authors' calculations (see appendix table A-1 for data sources).

rials price elasticity differs significantly from zero for all groups. It is highest for developed countries, at 0.19, and about 0.06 for all three groups of developing economies.

What is the evidence for a cost-reducing role for the stock of machinery? Elasticity estimates for the entire period show that imported machinery has a cost-reducing effect for developed countries and export-oriented developing economies. But this effect is significant for the imports of equipment in the past year, not for the stock of imported equipment at the start of the previous year. Thus the evidence suggests that the technology embodied in new imported equipment helps competitiveness and, moreover, acts relatively quickly. For export-oriented developing economies the coefficient on the lagged capital stock term is negative but statistically insignificant. This implies that the gains from new investment in imported capital goods are not reversed. For developed countries the coefficient on the lagged capital stock term is positive but never statistically different from zero, implying some persistence in the cost-reducing effects.

Do domestic and imported machinery substitute for each other? We use the nonnested Davidson and MacKinnon (1981) *J*-test to determine whether imported and domestically produced machinery are perfect or imperfect substitutes in terms of their cost-reducing effect (for a description of the test see also Greene 1997). If they are imperfect substitutes, we need to consider their cost-reducing effects separately, and the price equation with imported and domestic machinery as separate right-hand-side variables (equation [2]) is appropriate. However, if they are perfect substitutes, we need to include their sum, the total stock of machinery, as a right-hand-side variable. The usual nested test does not apply here because an alternative to the null hypotheses cannot be constructed by restricting the parameters implied by the null. Because of this property of the model, imperfect and perfect substitution are nonnested hypotheses.

The *J*-test is used in such situations, but because it is a two-way test its use may lead to inconclusive results. In the first stage (hypothesis test H2) imperfect substitution is the null hypothesis and perfect substitution is the alternative hypothesis.[10] The procedure works as follows. First we obtain the predicted export price under the assumption of perfect substitution (combining domestic and imported machinery to form one capital stock variable). Then we include this predicted export price as an additional variable in the export price estimation under the assumption of imperfect substitution. If the coefficient on the predicted export price variable differs significantly from zero, we can reject the hypothesis of imperfect substitution. The *J*-test amounts to testing whether the estimate of the dependent variable obtained under the alternative specification of perfect substitution has any explanatory power in the null specification of imperfect substitution for the export price function. If it does, we can reject the hypothesis of imperfect substitution. The *p*-values in tables 2 and 4–6 refer to the statistical

10. We thank an anonymous referee for suggesting the use of nonnested hypothesis tests.

significance of the coefficient on the predicted price estimated from the alternative hypothesis.

Next we take perfect substitution as the null hypothesis and imperfect substitution as the alternative and again conduct the *J*-test (H3). If the test fails to reject the null hypothesis of perfect substitution, we can conclude that the two types of machinery are perfect substitutes. If instead the test rejects the null hypothesis of perfect substitution, we need to look at the result of the test in which imperfect substitution is the null hypothesis (H2). If the null hypothesis of imperfect substitution cannot be rejected, we can conclude that the two types of machinery are imperfect substitutes.

In this instance the *J*-test results are quite clear. The null hypothesis of perfect substitution is rejected, but the null hypothesis of imperfect substitution cannot be rejected even at very high levels of significance.

Subsample Windows

Considerable changes occurred in the market power and the technology absorption capacity of different countries in 1967–90. To help us study the evolution of elasticity estimates over this period, we use subsample windows regressions. We start with the full sample, 1967–90. Then we drop the observation for 1967 and estimate the model for the subsample 1968–90. Next we drop the observation for 1968 and estimate the model for 1969–90, and so on up to the subsample window 1979–90. In this fashion we obtain 13 different estimates of elasticity. As we move from the first window (1967–90) to the last (1979–90), we obtain a better fit for the regressions (the adjusted R^2 increases) for developed countries and for import-substituting developing economies, and the quality of the fit remains relatively unchanged for export-oriented developing economies. The *J*-test continues to strongly reject the null hypothesis of perfect substitution between imported and domestic machinery but not the null hypothesis of imperfect substitution.

For developed countries the cost-reducing effect of new investment in imported machinery declined quite rapidly, and although the sign continued to be negative in all but one period, by the early 1970s the effect had become statistically insignificant (table 4). Soon thereafter, by the mid-1970s, the stock of domestic machinery had a cost-reducing effect. One could interpret this shift as implying that domestic capabilities matured by the early 1970s in developed countries and thus that the leading edge of the innovation process shifted from a reliance on external sources to a locus in domestic research and adaptation. This does not necessarily mean that domestic machinery embodied more sophisticated technologies than imported machinery. Instead, it suggests that domestic machinery came to play a more central role in a broader process of technological innovation, one that had persistent effects.

For export-oriented developing economies there was a similar pattern of evolution (table 5). The cost reducing effect of new investment in imported machinery remained statistically significant throughout the period, though there is some sug-

TABLE 4. Elasticity Estimates for Developed Countries, Subsample Windows

Subsample window	Elasticity estimates								Adjusted R^2	Degrees of freedom	Hypothesis tests (marginal significance levels)		
	p_w	w	p_m	e	I^m_{-1}	I^d_{-1}	K^m_{-2}	K^d_{-2}			H1	H2	H3
1967–90	0.355 ***	0.240***	0.187***	–0.712***	–0.052***	–0.016	0.097	–0.078	0.87	268	[<0.001]	0.09	[<0.001]
1968–90	0.333 ***	0.258***	0.200***	–0.724***	–0.051***	–0.017	0.061	–0.067	0.87	255	[<0.001]	0.10	[<0.001]
1969–90	0.280 ***	0.269***	0.223***	–0.766***	–0.053***	–0.019	–0.015	–0.108	0.88	242	[<0.001]	0.97	[<0.001]
1970–90	0.277 ***	0.254***	0.218***	–0.761***	–0.039**	–0.015	0.054	–0.115	0.88	229	[<0.001]	0.93	[<0.001]
1971–90	0.254 ***	0.263***	0.222***	–0.774***	–0.035*	–0.017	0.128	–0.113	0.89	216	[<0.001]	0.19	[<0.001]
1972–90	0.214 ***	0.288***	0.238***	–0.795***	–0.019	–0.023	0.142	–0.124	0.89	202	[<0.001]	0.04	[<0.001]
1973–90	0.180 **	0.244***	0.239***	–0.812***	–0.017	–0.017	0.129	–0.136	0.90	188	[<0.001]	0.10	[<0.001]
1974–90	0.147**	0.187***	0.250***	–0.831***	–0.009	–0.018	0.216	–0.112	0.90	174	[<0.001]	0.48	[<0.001]
1975–90	0.125*	0.134***	0.212***	–0.830***	–0.020	–0.017	0.100	–0.231**	0.90	160	[<0.001]	0.48	[<0.001]
1976–90	0.194**	0.079	0.185***	–0.770***	0.004	–0.027	0.283	–0.201**	0.92	146	[<0.001]	0.23	[<0.001]
1977–90	0.162*	0.084	0.184***	–0.786***	–0.012	–0.033	0.171	–0.347***	0.92	132	[<0.001]	0.65	[<0.001]
1978–90	0.154*	0.100	0.204***	–0.800***	–0.018	–0.035	0.193	–0.346***	0.93	118	[<0.001]	0.12	[<0.001]
1979–90	0.147	0.079	0.157***	–0.806***	–0.020	–0.023	0.147	–0.233**	0.93	104	[<0.001]	0.03	[<0.001]

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: p_w is the world price, w is the wage rate, p_m is the price of raw materials, e is the exchange rate, I^m_{-1} is investment in imported machinery in $t - 1$, I^d_{-1} is investment in domestic machinery in $t - 1$, and K^m_{-2} and K^d_{-2} are the stocks of imported and domestic machinery at the end of $t - 2$.

H1: World price elasticity is equal to one (joint test for market power and economies of scale).
H2: Davidson and MacKinnon *J*-test: imported and domestic machinery are imperfect substitutes.
H3: Davidson and MacKinnon *J*-test: imported and domestic machinery are perfect substitutes.
Source: Authors' calculations (see appendix table A-1 for data sources).

TABLE 5. Elasticity Estimates for Export-Oriented Developing Economies, Subsample Windows

Subsample window	Elasticity estimates p_w	w	p_m	e	I^m_{-1}	I^d_{-1}	K^m_{-2}	K^d_{-2}	Adjusted R^2	Degrees of freedom	Hypothesis tests (marginal significance levels) H1	H2	H3
1967–90	0.923***	0.112***	0.062***	−0.213***	−0.072**	−0.042	−0.029	−0.171	0.49	218	0.42	0.26	[<0.001]
1968–90	0.882***	0.108***	0.068***	−0.220***	−0.082**	−0.044	−0.093	−0.177	0.50	209	0.20	0.60	[<0.001]
1969–90	0.925***	0.102***	0.066***	−0.201***	−0.071**	−0.038	−0.006	−0.154	0.52	199	0.42	0.68	[<0.001]
1970–90	0.854***	0.114***	0.070***	−0.221***	−0.074**	−0.037	−0.045	−0.209	0.52	188	0.11	0.71	[<0.001]
1971–90	0.844***	0.109***	0.072***	−0.207***	−0.079**	−0.042	−0.072	−0.188	0.52	177	0.09	0.59	[<0.001]
1972–90	0.857***	0.108***	0.068***	−0.191***	−0.094***	−0.025	−0.081	−0.159	0.55	165	0.10	0.43	[<0.001]
1973–90	0.829***	0.099***	0.065***	−0.192***	−0.11***	−0.003	−0.243*	−0.149	0.58	153	0.05	0.91	[<0.001]
1974–90	0.800***	0.085***	0.081***	−0.163***	−0.102***	−0.002	−0.116	−0.189	0.56	141	0.02	0.92	[<0.001]
1975–90	0.811***	0.087***	0.058***	−0.138***	−0.099***	0.001	−0.138	−0.160	0.39	129	0.03	0.73	[<0.001]
1976–90	0.862***	0.064*	0.088***	−0.134***	−0.055**	−0.023	−0.095	−0.376***	0.51	117	0.10	0.74	[<0.001]
1977–90	0.875***	0.088***	0.083***	−0.093**	−0.067**	−0.018	−0.189	−0.44***	0.52	105	0.21	0.87	[<0.001]
1978–90	0.915***	0.067*	0.061***	−0.050	−0.063**	−0.006	−0.267	−0.261**	0.52	93	0.43	0.86	[<0.001]
1979–90	0.909***	0.072*	0.056***	−0.072	−0.033*	−0.002	−0.179	−0.207*	0.55	81	0.46	0.19	[<0.001]

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: p_w is the world price, w is the wage rate, p_m is the price of raw materials, e is the exchange rate, I^m_{-1} is investment in imported machinery in $t - 1$, I^d_{-1} is investment in domestic machinery in $t - 1$, and K^m_{-2} and K^d_{-2} are the stocks of imported and domestic machinery at the end of $t - 2$.

H1: World price elasticity is equal to one (joint test for market power and economies of scale).

H2: Davidson and MacKinnon *J*-test: imported and domestic machinery are imperfect substitutes.

H3: Davidson and MacKinnon *J*-test: imported and domestic machinery are perfect substitutes.

Source: Authors' calculations (see appendix table A-1 for data sources).

gestion that the size of the effect declined in the 1980s. Also in the 1980s, with the development of domestic capabilities (not always in advanced research but typically in rapid reverse engineering and adaptation of technologies), investment in domestic capital began to play a greater part in technological advance.

Finally, for import substituting economies the effect of imported equipment was negligible until the late 1970s (table 6). The inevitable opening of markets began to occur in these economies in the early 1980s, accompanied by domestic deregulation and thus greater competition from both domestic and foreign sources. During this period investment in imported goods began to have a greater cost-reducing effect. However, the results also suggest that the period of technological advance based on imported goods has not yet been followed by a shift to domestic sources of innovation.

IV. Conclusions

We have provided empirical evidence on the relationship between export competitiveness and the flows and stock of machinery, allowing for the possibility of imperfect substitution between domestically produced and imported machinery. Our results show that imported machinery has had an important cost-reducing effect in developed and export-oriented developing economies. This effect acted quickly and typically was not reversed. For developed countries the cost-reducing effect of imported capital goods faded by the early 1970s, presumably because the locus of innovation shifted increasingly to domestic sources. For export-oriented developing economies imports of capital goods continued to have an effect throughout the period, though the benefits from domestic innovation also became tangible in the early 1980s.

In contrast, in developing economies where import substitution had been the dominant trade strategy, exporters were unable or lacked the incentive to use imported machinery to improve their competitiveness until the late 1970s. Thereafter, as some of these economies became more open to international trade and less constrained by domestic regulation, imported capital goods began to play a greater role in innovation and thus to have a cost-reducing effect. Domestically produced machinery does not appear to have provided sustained aid to international competitiveness in such economies.

One interpretation of De Long and Summers (1991, 1992a, 1992b, 1993) is that because additions to the stock of machinery spur growth, government policies should support rapid increases in this stock. The authors themselves were cautious about drawing such a conclusion, however, and were more inclined to favor a liberal import regime. While rewarding entrepreneurial behavior, a liberal regime would also facilitate the inflow of imported equipment and thus foster growth.

Our results certainly support that view. But they also suggest a possibility for sequencing in innovative activities. Early innovation is most quickly achieved by importing technology. Domestic innovation capability can be built in parallel, however, ultimately becoming the principal locus of investment in innova-

TABLE 6. Elasticity Estimates for Import Substituting Developing Economies, Subsample Windows

Subsample window	Elasticity estimates								Adjusted R^2	Degrees of freedom	Hypothesis tests (marginal significance levels)		
	p_w	w	p_m	e	I^m_{-1}	I^d_{-1}	K^m_{-2}	K^d_{-2}			H1	H2	H3
1967–90	0.925***	–0.07	0.056***	–0.022	0.0	–0.058	0.068	–0.188	0.22	240	0.59	0.70	[<0.001]
1968–90	0.853***	–0.058	0.067***	–0.030	0.013	–0.073	–0.028	–0.246	0.26	228	0.26	0.54	[<0.001]
1969–90	0.952***	–0.012	0.061***	–0.044	–0.016	–0.052	0.118	0.018	0.38	216	0.68	0.38	[<0.001]
1970–90	0.946***	–0.006	0.056***	–0.046	–0.023	–0.044	0.207	0.016	0.40	203	0.62	0.64	[<0.001]
1971–90	0.956***	0.002	0.055***	–0.029	–0.018	–0.038	0.371*	–0.076	0.44	191	0.68	0.43	[<0.001]
1972–90	0.867***	–0.02	0.052***	–0.045	–0.028	–0.029	0.276	–0.033	0.44	179	0.17	0.59	[<0.001]
1973–90	0.850***	–0.041	0.049***	–0.014	–0.020	–0.039	0.276	–0.125	0.45	166	0.11	0.65	[<0.001]
1974–90	0.818***	0.080	0.049***	–0.063*	–0.020	–0.007	0.367*	0.003	0.46	153	0.03	0.83	[<0.001]
1975–90	0.781***	0.079	0.034***	–0.063*	–0.026	–0.015	0.308	0.012	0.28	140	0.01	0.90	[<0.001]
1976–90	0.885***	0.031	0.047***	–0.070**	–0.018	–0.035	0.276	–0.086	0.37	127	0.17	0.82	[<0.001]
1977–90	0.832***	0.052	0.047***	–0.084**	–0.055**	–0.013	0.105	0.150	0.45	114	0.05	0.35	[<0.001]
1978–90	0.904***	0.042	0.029***	–0.059*	–0.056**	–0.044	0.139	0.196	0.53	101	0.25	0.46	[<0.001]
1979–90	0.918***	0.001	0.036***	–0.028	–0.078**	–0.016	–0.138	0.330	0.58	88	0.39	0.60	[<0.001]

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: p_w is the world price, w is the wage rate, p_m is the price of raw materials, e is the exchange rate, I^m_{-1} is investment in imported machinery in $t - 1$, I^d_{-1} is investment in domestic machinery in $t - 1$, and K^m_{-2} and K^d_{-2} are the stocks of imported and domestic machinery at the end of $t - 2$.

H1: World price elasticity is equal to one (joint test for market power and economies of scale).

H2: Davidson and MacKinnon J-test: imported and domestic machinery are imperfect substitutes.

H3: Davidson and MacKinnon J-test: imported and domestic machinery are perfect substitutes.

Source: Authors' calculations (see appendix table A-1 for data sources).

tion. In the wake of increasing labor costs, countries adopting an export-oriented strategy, especially the East Asian newly industrializing countries, relied heavily on machinery imports to acquire modern technology. Governments and private businesses supported the absorption and adaptation of imported technology through local R&D and engineering efforts. Over time these domestic efforts have become increasingly important.

Further analysis along these lines would benefit from disaggregated time-series data on manufacturing subsectors. In particular, sectoral data on machinery investment and machinery prices would make it possible to endogenize the use of machinery.

Our results also point to the importance of trade as a vehicle for the transfer of knowledge, identifying capital goods as the conduit. Recently, however, Keller (2000) and Branstetter (2001) have argued that knowledge spillovers within a country are quantitatively more important than international knowledge transfer. Our results suggest that the relative importance of internal and external knowledge spillovers may change as the international environment changes and as domestic incentives and absorptive capacity evolve. Thus further exploration of the determinants of internal and external knowledge spillovers is also likely to be a fruitful avenue of research.

Appendix: Data Sources, Descriptive Statistics, and Unit Root Tests

Table A-1. Definition of Variables and Data Sources

Variable	Meaning and data source
Developed and developing economies	
p_w	Unit value index for manufactured exports from the rest of the world, 1987 = 100. Calculated from the export prices of all other countries weighted by their world market shares. (*Source:* World Bank, IECTRADE database)
w	Annual wage per employee in manufacturing, in thousands of U.S. dollars (total wage bill/number of employees). (*Source:* U.N. Industrial Development Organization (UNIDO) sectoral database.) The dollar value was converted to local currency using the exchange rate e.
e	Exchange rate, units of local currency per U.S. dollar, annual average. (*Source:* International Monetary Fund, International Financial Statistics database)
K^T	Stock of total machinery, in constant 1985 U.S. dollars. Calculated from machinery investment data using the perpetual inventory method and assuming a depreciation rate of 12 percent. (*Source:* Penn World Tables, version 5.6)
K^m	Stock of imported machinery, in constant 1985 U.S. dollars. Obtained from imports of nonelectrical machinery (711, 712, 714, 715, 717, 718, 719) and electrical machinery (722, 723, 72491, 726, 7295, 7296, 7297, 7299), Standard Industrial Trade Classification (SITC), rev. 1, using the perpetual inventory method with a 12 percent depreciation rate. (*Source:* United Nations, Commodity Trade (Comtrade) Database.) *Note:* To obtain imports in constant prices, import data in current U.S. dollars were deflated by the dollar price of investment goods from the Penn World Tables, version 5.6.
K^d	Stock of domestically produced machinery ($K^T - K^m$).
Developed countries only	
p	Price index for manufactured exports of country j, U.S. dollars, 1987 = 100. (*Source:* OECD)
p_m	Price index for imported raw materials, local currency, 1987 = 100. (*Source:* OECD)
Developing economies only	
p	Price index for manufactured exports, U.S. dollars, 1987 = 100. (*Source:* World Bank, IECTRADE database)
p_m	Price index for crude petroleum imports, U.S. dollars, 1987 = 100. (*Source:* World Bank, IECTRADE database.) The U.S. dollar value was converted to local currency using the exchange rate e.

TABLE A-2. Descriptive Statistics

	Period average (standard deviation)				Capital stock (1990) (billions of 1985 U.S. dollars)		
	p (1987 = 100)	w ($1000)	p_m (1987 = 100)	e (LC/US$)	K^T	K^d	K^m
Developed countries							
Australia	82.6	10.9	78.5	1.0	95.8	56.3	39.5
	(23.9)	(5.3)	(27.8)	(0.2)			
Austria	63.4	9.8	53.9	18.3	42.2	14.3	27.9
	(26.5)	(6.5)	(32.2)	(4.9)			
Belgium-Luxembourg	65.8	10.6	67.0	42.3	52.4	6.5	45.9
	(29.0)	(6.0)	(34.8)	(8.5)			
Denmark	62.7	14.1	51.8	7.1	31.3	11.6	19.7
	(27.1)	(7.9)	(34.2)	(1.4)			
Finland	61.6	10.3	65.2	4.3	36.2	17.2	19.0
	(31.9)	(7.2)	(31.4)	(0.7)			
France	63.4	9.9	60.9	5.7	339.2	213.2	125.9
	(27.3)	(5.8)	(32.5)	(1.3)			
Germany	61.9	13.5	61.9	2.6	390.7	248.5	142.2
	(27.4)	(8.1)	(32.7)	(0.8)			
Italy	62.4	8.5	67.0	1,007.2	307.0	230.8	76.2
	(29.6)	(5.5)	(26.0)	(407.5)			
Japan	65.4	11.1	51.0	253.5	860.7	805.3	55.4
	(27.7)	(8.4)	(33.5)	(74.8)			
Netherlands	67.0	12.7	57.1	2.7	76.2	12.6	63.7
	(26.7)	(7.1)	(36.9)	(0.6)			
New Zealand	70.1	8.8	55.0	1.2	17.1	8.9	8.2
	(32.0)	(4.5)	(29.5)	(0.4)			
Sweden	66.3	11.9	79.2	5.6	54.69	13.84	40.84
	(29.1)	(5.5)	(26.1)	(1.3)			
United Kingdom	66.4	9.0	65.6	0.5	335.9	196.2	139.7
	(30.2)	(5.8)	(32.2)	(0.1)			
United States	69.8	15.6	66.6	1.0	1603.1	1203.8	399.3
	(28.9)	(6.9)	(30.9)	(0.0)			
Export-oriented developing economies							
Brazil	88.0	3.4	151.5	0.0	101.2	80.3	20.9
	(18.5)	(1.0)	(58.9)	(0.0)			
Greece	93.5	5.9	146.7	85.4	20.1	10.2	9.9
	(18.3)	(2.1)	(68.0)	(51.2)			
Hong Kong (China)	83.1	1.6	151.5	6.3	23.5	6.7	16.9
	(18.2)	(0.7)	(58.9)	(1.4)			
Indonesia	91.0	0.8	151.5	936.2	75.7	53.7	22.0
	(20.9)	(0.2)	(58.9)	(526.3)			
Israel	93.2	11.6	150.3	0.6	23.1	12.8	10.3
	(16.0)	(5.5)	(65.0)	(0.8)			
Korea, Rep. of	91.9	3.6	149.8	653.6	84.8	38.4	46.5
	(17.9)	(2.4)	(66.5)	(156.0)			
Malaysia	83.7	2.3	151.5	2.4	27.6	8.6	19.0
	(22.7)	(0.7)	(58.9)	(0.2)			

(continued)

TABLE A-2. (continued)

	Period average (standard deviation)				Capital stock (1990) (billions of 1985 U.S. dollars)		
	p (1987 = 100)	w ($1000)	p_m (1987 = 100)	e (LC/US$)	K^T	K^d	K^m
Portugal	81.0	3.5	161.0	92.1	16.9	3.4	13.4
	(18.4)	(1.1)	(66.8)	(54.3)			
Singapore	80.0	5.5	151.5	2.2	37.6	8.5	29.1
	(21.0)	(2.6)	(58.9)	(0.2)			
Thailand	86.6	1.4	152.1	22.9	39.1	19.0	20.1
	(19.3)	(0.6)	(62.5)	(2.5)			
Turkey	91.1	3.7	153.3	545.1	64.0	48.1	15.9
	(28.3)	(1.2)	(66.1)	(789.5)			
Uruguay	77.7	3.3	151.5	0.2	5.6	4.5	1.1
	(23.6)	(1.2)	(58.9)	(0.3)			
Import-substituting developing economies							
Argentina	76.2	5.5	160.5	0.03	12.0	3.7	8.3
	(25.4)	(2.3)	(56.8)	(0.1)			
Colombia	76.1	2.3	156.8	136.4	12.5	5.6	7.0
	(23.3)	(0.6)	(75.4)	(141.8)			
Guatemala	82.2	2.0	155.4	1.6	3.4	2.1	1.3
	(16.9)	(0.6)	(54.8)	(1.0)			
Honduras	83.1	2.7	155.4	2.1	2.4	1.6	0.8
	(18.3)	(0.9)	(54.8)	(0.5)			
India	91.0	1.0	162.9	10.8	165.9	150.5	15.4
	(20.9)	(0.3)	(53.3)	(3.0)			
Kenya	96.2	0.1	151.5	12.5	3.2	1.0	2.1
	(14.6)	(0.0)	(58.9)	(5.2)			
Mexico	81.0	4.3	130.7	0.6	98.6	68.6	30.1
	(22.1)	(1.1)	(50.8)	(1.0)			
Pakistan	89.7	1.1	151.5	13.4	68.6	62.2	6.5
	(17.1)	(0.4)	(58.9)	(4.1)			
Panama	89.7	4.5	173.4	1.0	1.4	0.5	0.9
	(17.1)	(1.1)	(62.9)	(0.0)			
Peru	91.0	2.9	155.8	0.01	10.1	6.3	3.9
	(15.3)	(1.0)	(72.7)	(0.0)			
Philippines	81.9	1.3	148.5	13.1	21.5	14.2	7.3
	(15.4)	(0.4)	(63.4)	(6.6)			
Sri Lanka	88.3	1.8	153.4	21.2	2.3	0.9	1.4
	(16.6)	(0.3)	(56.5)	(10.2)			
Venezuela,	93.8	7.6	151.5	10.4	28.9	11.5	17.3
R.B. de	(26.0)	(2.9)	(58.9)	(12.1)			

Source: Authors' calculations (see appendix table A-1 for data sources).

TABLE A-3. Trade Regimes of the Sample Developing Economies

1967–73				1973–85			
Outward oriented		Inward oriented		Outward oriented		Inward oriented	
Strongly	Moderately	Strongly	Moderately	Strongly	Moderately	Strongly	Moderately
Hong Kong (China)	**Brazil**	Honduras	Argentina	**Hong Kong (China)**	**Brazil**	Colombia	Argentina
Korea, Rep. of	Colombia	Kenya	India	**Korea, Rep. of**	**Israel**	Guatemala	India
Singapore	Guatemala	Mexico	Pakistan	**Singapore**	**Malaysia**	Honduras	Peru
	Indonesia	Philippines	Peru		**Thailand**	**Indonesia**	
	Israel		Sri Lanka		**Turkey**	Kenya	
	Malaysia		**Turkey**		**Uruguay**	Mexico	
	Thailand		**Uruguay**			Pakistan	
						Philippines	
						Sri Lanka	

Note: Based on World Bank classification. Not included in the classification are Greece, Panama, Portugal, and R.B. de Venezuela. Countries in bold face are treated as export oriented; the others are treated as import substituting.

Source: World Bank 1986.

DEFINITIONS

STRONGLY OUTWARD ORIENTED. Trade controls are either nonexistent or very low in the sense that any disincentives to export resulting from import barriers are more or less counterbalanced by export incentives. There is little or no use of direct controls and licensing arrangements, and the effective exchange rates for imports and exports are roughly equal.

MODERATELY OUTWARD ORIENTED. Incentives favor production for domestic rather than export markets. But the average rate of effective protection for the home market is relatively low, and the range of effective protection rates relatively narrow. The use of direct controls and licensing arrangements is limited. The effective exchange rate is higher for imports, but only slightly.

MODERATELY INWARD ORIENTED. Incentives clearly favor production for the domestic market. The average rate of effective protection for the home market is fairly high, and the range of effective protection rates relatively wide. Direct import controls are extensive. The exchange rate is somewhat overvalued.

STRONGLY INWARD ORIENTED. Incentives strongly favor production for the domestic market. The average rate of effective protection for the home market is high and the range of effective protection rates wide. Direct controls and licensing disincentives for the traditional export sector are pervasive, positive incentives for nontraditional exports are few or nonexistent, and the exchange rate is substantially overvalued.

TABLE A-4. Im-Pesaran-Shin Unit Root Test Results

	Average augmented Dickey-Fuller test statistic	Expected value[a]	Standard error[a]	Standardized average augmented Dickey-Fuller test statistic
Developed countries				
Variables in first differences				
Δp	-2.5276	-1.4290	0.1688	-6.5060**
Δw	-2.5814	-1.4492	0.1663	-6.8068**
Δp_m	-2.5718	-1.4810	0.1694	-6.4389**
Δe	-2.3588	-1.4401	0.1675	-5.4833**
ΔK^T	-2.2518	-1.6831	0.1724	-3.2990*
ΔK^m	-2.3732	-1.5619	0.1683	-4.8201**
ΔK^d	-2.6671	-1.6402	0.1680	-6.1125**
Variables in levels				
p	-1.8951	-1.4752	0.1645	-2.5530
w	-1.6833	-1.5182	0.1660	-0.9947
p_m	-1.4419	-1.6494	0.1711	1.2127
e	-1.2044	-1.5305	0.1658	1.9671
K^T	-1.9335	-1.5804	0.1671	-2.1130
K^m	-1.4902	-1.7306	0.1730	1.3891
K^d	-1.4609	-1.7087	0.1697	1.4599
Developing economies				
Variables in first differences				
Δp	-2.5336	-1.5538	0.1860	-5.2669**
Δw	-2.1085	-1.4809	0.1805	-3.4769*
Δp_m	-2.9571	-1.5463	0.1811	-7.7875**
Δe	-2.1858	-1.5420	0.1893	-3.3992*
ΔK^T	-2.3027	-1.5194	0.1749	-4.4792*
ΔK^m	-2.1896	-1.5907	0.1783	-3.3589*
ΔK^d	-2.2340	-1.4719	0.1707	-4.4646**
Variables in levels				
P	-1.1557	-1.6773	0.1928	2.7043
w	-1.1079	-1.5930	0.1836	2.6425
p_m	-1.5940	-1.5661	0.1773	-0.1572
e	-0.9703	-1.8957	0.1844	5.0184
K^T	-1.7476	-1.3673	0.1850	2.0562
K^m	-1.0986	-1.6987	0.1844	3.2539
K^d	-1.2850	-1.6709	0.1795	2.1480

*The null hypothesis of a unit root in each country's variable is rejected at the 5 percent level of significance.

**The null hypothesis of a unit root in each country's variable is rejected at the 1 percent level of significance.

[a]The expected value and standard error of the average augmented Dickey-Fuller test statistic are computed through stochastic simulations with 10,000 replications.

Source: Authors' calculations (see appendix table A-1 for data sources).

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Trade Policy Options for Chile: The Importance of Market Access

Glenn W. Harrison, Thomas F. Rutherford, and David G. Tarr

This article uses a multisector, multicountry, computable general equilibrium model to examine Chile's strategy of "additive regionalism"—negotiating bilateral free trade agreements with all of its significant trading partners. Taking Chile's regional arrangements bilaterally, only its agreements with Northern partners provide sufficient market access to overcome trade diversion costs. Due to preferential market access, however, additive regionalism is likely to provide Chile with gains that are many multiples of the static welfare gains from unilateral free trade. At least one partner country loses from each of the regional agreements considered, and excluded countries as a group always lose. Gains to the world from global free trade are estimated to be vastly larger than gains from any of the regional arrangements.

The analysis of regional trade arrangements is typically conducted in the framework of trade creation versus trade diversion, under which preferential tariff reduction is welfare inferior to nonpreferential tariff reduction. However, Wonnacott and Wonnacott (1981) show that regional trade arrangements could produce more gains due to improved market access to trading partners. The logical extension of this argument is that if a country negotiated free trade agreements with all of its trade partners, it would end up with zero effective tariffs on all imports, or free trade, despite the legal existence of positive most-favored-nation (MFN) tariffs. In the process, it would achieve preferential access to its partners' markets. Hence, without transition dynamics, this strategy may produce gains that are considerably larger than unilateral free trade.

We call the process of sequentially negotiating bilateral free trade agreements with all significant trading partners "additive regionalism." There is at least one country, Chile, that is pursuing a clearly articulated strategy of additive regionalism.[1] Does additive regionalism dominate free trade for Chile? If so, by how much?

The government of Chile has successfully concluded a free trade area (FTA) with MERCOSUR and is seeking a free trade agreement with the North American

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1. Mexico, Singapore, and, to a lesser extent, MERCOSUR may be following the same strategy.

Free Trade Agreement (NAFTA).[2] Moreover, the government of Chile is attempting to add the European Union, the rest of South America, and several other countries to its network of free trade arrangements.[3]

It is well known that most results regarding the welfare effects of regional arrangements are typically ambiguous at the theoretical level and that many questions are quantitative rather than qualitative. Thus, we employ an 11-region global computable general equilibrium (CGE) model to examine quantitatively the network of preferential arrangements that Chile is negotiating as well as unilateral trade policy options in Chile. In addition, we estimate the impact of global free trade as a reference point. Our model includes the Chilean economy as well as the economies of Argentina, Brazil, Mexico, the United States, Canada, Central America, the rest of South America, the European Union, Japan, and an aggregate for the rest of the world. Consequently, we are able to estimate the impact on partner and excluded countries from each of the agreements we evaluate.

Critics of Chile's additive regionalism strategy, such as Donoso and Hachette (1996), argue that agreements with Southern countries are unlikely to be beneficial and that it is not worth delaying the benefits of unilateral and multilateral tariff liberalization to pursue these agreements. They argue that only agreements with the European Union, the United States, or Japan offer sufficient access to be worth pursuing. Advocates for the government's strategy, however, believe that there are gains to be achieved from agreements with smaller Southern countries as well. They also argue, as in Butelmann and Meller (1995), that additive regionalism will progressively reduce trade diversion costs, lower the effective average tariff in Chile, and provide considerably improved market access. Furthermore, they note that Chile can unilaterally lower its external tariff while simultaneously pursuing additive regionalism to further reduce trade diversion costs.

We find that the results for NAFTA, MERCOSUR, and especially additive regionalism all point to the crucial importance of improved market access in preferential trading areas. Considered bilaterally, we find that trade diversion costs do indeed dominate the welfare effects of these agreements unless sufficient market access is obtained in partner countries (or third-country tariffs are lowered).

The results support the view that North–South agreements (for example, Chile with the United States or the European Union) are likely to provide sufficient market access to be beneficial, whereas the results for our South–South agreement (Chile-MERCOSUR) suggest the opposite. Agreements that include a Northern

2. MERCOSUR is a customs union between Argentina, Brazil, Paraguay, and Uruguay. Paraguay and Uruguay are too small to be included as separate countries in the dataset we employ, so our MERCOSUR region excludes them. In a FTA, partner countries eliminate tariffs and export taxes or subsidies against each other but retain separate tariffs against third countries. In a customs union, partner regions adopt a common external tariff. Chile has rejected a customs union with MERCOSUR. Although negotiations for Chile's membership in NAFTA have stalled, many commentators believe that Chile will eventually become the next member of NAFTA.

3. As of early 2001, Chile had reached preferential trade agreements with at least 15 countries.

partner increase the welfare of the members of the group in aggregate; only the Chile-MERCOSUR agreement results in net losses for the members as a group. However, Chile can unilaterally lower the external tariff, reducing trade diversion, so that even its agreement with MERCOSUR is beneficial.[4]

We find that Chile's additive regionalism strategy of combining free trade agreements with four regions—NAFTA, MERCOSUR, the European Union, and rest of South America—produces welfare gains for Chile that are many multiples of the value of unilateral free trade if it were to attain tariff-free access to all these markets. This provides support for the theoretical insight of Wonnacott and Wonnacott (1981). However, if the most highly protected sectors in the European Union and rest of South America are excluded from the agreements, the gains are dramatically reduced.[5]

We estimate that at least one of Chile's potential partners in its additive regionalism strategy will lose in all of the options we evaluate. Adding the rest of South America to its network of agreements would substantially improve Chile's preferential access and welfare but would significantly reduce the real income of the rest of South America, which would suffer large trade diversion losses with very little improved market access. Theory, intuition, and experience indicate that preferential arrangements are unlikely to be implemented if the partner countries do not also expect to gain. Nonetheless, the gains for Chile remain substantial relative to unilateral free trade if it could successfully negotiate these agreements with full market access.

Excluded regions are always estimated to lose from any of the preferential arrangements we consider. Thus, where there are gains to partner countries from preferential arrangements, they come at least partly at the expense of excluded regions.

The gains to the world from global free trade are estimated to be between US$199 billion and $456 billion per year. These gains to the world vastly exceed the gains from any of the regional arrangements. These results emphasize the continuing importance of multilateral liberalization.

Because Chile starts with a relatively efficient uniform tariff of 11 percent, we estimate that it can obtain only small additional gains from improving the efficiency of its resource allocation by further *unilateral* reduction of its tariffs.[6] We show that a country like Chile that starts with a uniform tariff will typically have

4. Chile has enacted legislation that will lower its external tariff from 11 to 6 percent in stages, as suggested by our analysis. Thus our estimates could be viewed as an ex post assessment of the policy of lowering the external tariff. In fact, the vice president of the Chilean Central Bank used estimates from an earlier version of our study in his testimony before the Chilean Parliament in favor of lowering the external tariff.

5. In fact, the experiences of some Mediterranean countries (Morocco, Tunisia, and Turkey) in their preferential trade agreements with the European Union suggest that the highly protected agricultural sectors are likely to be excluded from such an agreement.

6. This conclusion ignores the dynamic gains from trade liberalization, which could lead to much larger gains.

the gains from joining a customs union reduced if it must adopt a nonuniform structure. Conversely, if joining a customs union is a movement toward uniformity, the gains are likely to be augmented.[7] In general, this result indicates that the relative uniformity of the preexisting tariff structure for a country and the proposed common external tariff of any customs union, must be compared on a case-by-case basis to ascertain whether welfare gains will actually be achieved.

We find that the benefits of trade liberalization or regional trade arrangements are considerably reduced if tariff revenue must be replaced by distorting alternative taxes. Similarly, in our optimal tariff calculations, we find that unilateral trade liberalization can proceed to lower tariff levels if efficient replacement taxes are in place.[8]

When there is an optimal tariff, as there is in this model, the amount by which a country can reduce its tariff is limited by the distortions of the replacement tax. Consequently, we have produced an updated estimate of the collected value-added tax (VAT) rates by sector in Chile.[9] We show that Chile could reduce its legal VAT rates to about 50 percent of present levels and improve its welfare by 0.3 percent of gross domestic product (GDP) if it were able to eliminate evasion and collect the VAT uniformly.[10] These gains are significant when compared to unilateral trade liberalization options. We find that the optimal tariff in Chile is almost double with the VAT rate that is currently collected, compared with a VAT that collects taxes at equal rates across sectors.

Section I describes the model and data. Section II explains the policy results for Chile. Section III examines the impact on partner and excluded countries of Chile's agreements as well as the impact of global free trade.

I. A Multiregional Trade Model

General Features

The quantitative model developed to evaluate the trade policy options facing Chile is multi-regional and multi-sectoral. Table 1 lists the 11 regions included explicitly in the model, as well as the 24 sectors included in each region. The general

7. Two other countries with uniform tariffs that may install a nonuniform tariff of a customs union are the Kyrgyz Republic and Estonia. The Kyrgyz Republic has a uniform tariff of 10 percent and has in principle agreed to join in a customs union with Russia, Belarus, and Kazakhstan. The Kyrgyz Republic has not implemented the common external tariff, however, because of fears of the costs of the nonuniformity of the Russian tariff, which is the present common external tariff. See Michalopoulos and Tarr (1997) for details. Estonia also has a uniform tariff of zero and is one of the five transition economies the European Union has designated as candidate countries for accession. Estonian authorities have considerable concerns, however, about the costs of imposing the European Union's common external tariff, especially in the highly protected sectors.

8. However, with low elasticities, there is an adverse terms-of-trade effect that mitigates the welfare gains from reduced costs of trade diversion.

9. See Harrison, Rutherford, and Tarr (1997b).

10. In addition, we eliminate the output tax, which applies primarily to energy, beverages, and tobacco.

TABLE 1. Commodities, Regions, and Factors of Production in the Chile Model

Abbreviation	Meaning
Commodities	
WHT	Wheat
GRO	Other grains
NGC	Nongrain crops
WOL	Wool and other livestock
FRS	Forestry
FSH	Fishing
ENR	Energy products
MIN	Mineral products
MEA	Meat products
MIL	Milk products
FOO	Other food products
B_T	Beverages and tobacco
TEX	Textiles and apparel and leather products
LUM	Lumber and wood
PPP	Pulp and paper
CRP	Chemicals rubber and plastics
I_S	Primary ferrous metals
NFM	Nonferrous metals
FMP	Fabricated metal products
TRN	Transport industries
MAC	Machinery and equipment
T_T	Trade and transport
SER	Services
CGD	Savings good
Regions	
CHL	Chile
ARG	Argentina
BRA	Brazil
RSA	Rest of South America
USA	United States of America
CAN	Canada
MEX	Mexico
CAM	Central America and Caribbean
E_U	European Union 15
JPN	Japan
ROW	Rest of world
Factors	
LND	Land
LAB	Labor
CAP	Capital

specification of this model follows our earlier multiregional model of the effects of the Uruguay Round.[11] The most important differences are the inclusion of data for Chile, updated tariff rates for Argentina and Brazil, and more recent data for all other regions. We adopt a multiregion model rather than a small-open-economy model because we need to consider the possible effects on Chile of a reduction in its import tariffs on other MERCOSUR members. Crucially, we also need to account for the "market access" effects on Chilean exports of a reduction of import tariffs by MERCOSUR, NAFTA, or other regions with which Chile agrees to a free trade agreement either separately or collectively.

Although the general theory of the welfare effects of preferential trading arrangements allows for the impact of changes in partner country tariffs on the home country's terms of trade,[12] some empirical approaches to evaluating preferential trading arrangements ignore them.[13] Our framework allows us to evaluate explicitly the importance to Chile of improved market access to regions such as MERCOSUR and NAFTA as well as losses Chile may suffer as partner countries raise export prices to Chile.

An important feature of the Chilean economy is that its tariff rate is a uniform 11 percent across all traded sectors. The exception to this is the variable levy system for wheat, sugar, and edible oils. Estimates reveal that the variable levy system has resulted in an average level of protection for these three products in excess of 11 percent.[14] We ignore the variable levy system, which will slightly bias downward our estimated gains from unilateral trade liberalization. Harrison, Rutherford, and Tarr (1997b) describe the key data that are important in the analysis.

Argentine tariffs are virtually identical to Brazilian tariffs. In the case of the United States, the tariff estimates include the tariff equivalents of the nontariff barriers, which are quite important in the sectors with high tariffs. If Chile forms an FTA with MERCOSUR or NAFTA, Chilean exporters will not face these tariffs, but outside exporters to these regions will. Thus, these data are crucial in assessing the value of increased access that Chile will obtain from MERCOSUR and NAFTA, respectively.

11. Harrison, Rutherford, and Tarr (1997c). The Web site http://dmsweb.badm.sc.edu/glenn/ur_pub.htm provides access to the model and related publications.

12. See Wooton (1986) and Harrison, Rutherford, and Wooton (1989, 1993).

13. An example is the approach adopted by Bond (1996). He develops a simple general equilibrium specification of the effects on Chile of these preferential trading arrangements with an impressive level of detail with respect to tariff data. However, his results for Chile joining NAFTA differ significantly from ours because his CGE model does not incorporate the impact on Chile of access to NAFTA markets.

14. The variable levy system is applied by examining monthly prices over the previous 2.5 years for wheat and 50 months for sugar. The distribution is truncated at the top and the bottom by an equal percentage (about 15 percent). The range of the resulting truncated distribution determines the upper and lower bounds. A tariff surcharge or reduction of the tariff below the 11 percent rate is applied if the price in the present month is below or above the bounds. Because the system is not based on a domestic support price, its impact varies enormously from year to year. Valdes (1996, p. 55) estimates that between 1985 and 1995 the nominal protection rate for sugar ranged from 6 to 98 percent, and the nominal protection rate for wheat ranged from 45 to −10 percent (see also Quiroz and Valdes 1993).

We also estimate the rates of collected VAT in each industry and the tax on gross output, respectively. These rates are estimated using the procedures explained in Harrison, Rutherford, and Tarr (1997b, appendix A). The different rates of VAT across sectors arise mainly because of evasion of the VAT. The two largest sectors in Chile, trade and transport services and other services, have a combined 61 percent of value added and are the sectors with the lowest rate of collected VAT (about 3 percent as opposed to about 17 percent for most Chilean manufacturing).

Formal Specification

THE MODEL. The general specification of the model follows our earlier work on the Uruguay Round. We concentrate here on what we call our base model, which is static and assumes constant returns to scale. Except for the fact that imports and exports are distinguished by many regions, the structure of the model within any country is very close to the basic model of de Melo and Tarr (1992).

Production entails the use of intermediate inputs and primary factors (labor, capital, and land). Primary factors are mobile across sectors within a region but are internationally immobile. We assume constant elasticity of substitution (CES) production functions for value added and Leontief production functions for intermediates and the value-added composite. Output is differentiated between domestic output and exports, but exports are not differentiated by country of destination.

Each region has a single representative consumer that maximizes utility, as well as a single government agent. In Harrison, Rutherford, and Tarr (1997b, appendix C), we formally characterize the demand structure and elasticities that are critical to the results. Demand is characterized by nested CES utility functions for each agent, which allows multistage budgeting. Demand at the top level, for the composite "Armington" aggregate of each of the 24 goods in table 1, is Cobb-Douglas. Consumers first choose how much of each Armington aggregate good to consume, such as wheat, subject to aggregate incomes and composite prices of the aggregate goods. The Armington aggregate good is in turn a CES composite of domestic production and aggregate imports. Consumers decide how much to spend on aggregate imports and the domestic good subject to the prior decision of how much income will be spent on this sector, and preferences for aggregate imports and domestic goods are represented by a CES utility function. Finally, consumers decide how to allocate expenditures across imports from the 10 other regions based on their CES utility function for imports from different regions and income allocated to consumption on imports from the previous higher-level decision.

DATA AND ELASTICITIES. Except for tariff data and domestic tax data, the data employed to calibrate the model come primarily from the Global Trade Analysis Project (GTAP) database documented in Gehlhar and others (1996). We use the preliminary release of version 3 of this database, current as of May 1996.

The 11-region version of the model retains all regions of the GTAP database that are directly relevant to our policy simulations. The full GTAP database contains 37 sectors.[15]

We generally assume that the lower-level elasticity of substitution between imports from different regions, σ_{MM}, is 30 and that the higher-level elasticity between aggregate imports and domestic production, σ_{DM}, is 15. We refer to these values as our central elasticities. Econometric studies, such as those of Reinert and Roland-Holst (1992) and Shiells and Reinert (1993), suggest lower values. However, Reidel (1988) and Athukorala and Reidel (1994) argue that when the model is properly specified, the demand elasticities are not statistically different from infinity and their point estimates are close to the central elasticity values we have chosen. Moreover, elasticities would be expected to increase over time, and this model presumes an adjustment of about 10 years, a rather long period in the context of these econometric estimates.

To be clear, a value of $\sigma_{MM} = 30$ means that if Chile tried to raise its prices by 1 percent on world markets relative to an average of aggregate imports, Chilean imports would decline relative to aggregate imports by 30 percent. Given that there may be some economists who would prefer lower elasticity estimates, we also perform most of our important policy simulations with $\sigma_{MM} = 8$ and $\sigma_{DM} = 4$. We refer to these as our low elasticities. A high-elasticity scenario for a small open economy such as Chile would be a specification with still less market power for exports, such as would occur in the popular theoretical models of international trade where goods are homogeneous.

The elasticity of transformation between exports and domestic production is assumed to be about four for each sector. Elasticities of substitution between primary factors of production are taken from Harrison and others (1993) and generally reflect econometric estimates for the United States. These estimates are relatively low for primary goods, around unity for manufacturing goods, and elastic for tertiary goods. We assume fixed coefficients between all intermediates and value added.

DISTORTIONS. All distortions are represented as ad valorem price wedges. Border protection estimates combine tariff protection and the tariff equivalents of nontariff barriers. For Brazil and Argentina, these data were estimated by Reincke in Harrison, Rutherford, and Tarr (1997b, appendix B). Otherwise, these data are taken from the GTAP database. They are presented in Harrison, Rutherford, and Tarr (2001, table 9). Other distortions include factor taxes in production, VATS, export subsidies and voluntary export restraints (represented as ad valo-

15. Our aggregation to 24 sectors was undertaken in a manner that ensured that those sectors with significant rates of protection (in the principal trading partners of Chile) are retained as individual sectors. That is, we aggregated sectors that are not important in trade or that have low rates of protection. Aggregation may significantly change the results in applied trade policy analysis, but this type of aggregation results in quite small aggregation bias.

rem export tax equivalents). These are also taken from the GTAP database, except for domestic distortion data in Chile. The latter were estimated for this exercise by Soloaga in Harrison, Rutherford, and Tarr (1997b, appendix A). Lump-sum replacement taxes or subsidies ensure that government revenue in each region stays constant at real benchmark levels. However, for Chile, we capture the marginal efficiency cost of the government having to raise extra revenue through a distortionary domestic tax system. For developing countries, these costs could be quite significant because the revenue losses from trade reform could be sizable.

SOLUTION ALGORITHM. The model is formulated using the GAMS-MPSGE software developed by Rutherford (1999) and solved using the PATH algorithm of Ferris and Munson (2000). Although the model has 11 regions and 24 sectors and is large by historical standards, it is smaller than our Uruguay Round model. Use of demand elasticities as high as those we employ could pose numerical problems in general, but this model solved without difficulty.

II. Policy Results for Chile

We first discuss how Chile will replace the revenue it will lose from lowering its tariffs and the welfare implications of these options. We then discuss the results regarding the preferential trade area policy options and examine how Chile may use unilateral tariff reduction to optimize its trade policy. Finally, we examine the effects of Chile's strategy of additive regionalism.

The Role of the Replacement Tax

Because Chile is reducing tariffs in most of our scenarios, there is a revenue loss to the government. We impose an equi-revenue requirement in all simulations and stipulate explicitly how the additional tax revenue will be generated. We employ the existing VAT, a uniform VAT, or a lump-sum tax.

WELFARE EFFECTS OF THE REPLACEMENT TAX. Collection of the existing VAT is not uniform in Chile. According to the estimates in Harrison, Rutherford, and Tarr (1997b, table 3), it ranges from 0 percent up to 18 percent across sectors. Hence, raising revenue through the VAT generates distortions: when the VAT is increased, resources move into less highly taxed sectors. This reduces any possible gains from the trade policy change. Results for welfare using the existing VAT are presented in column 1 of table 2.

In fact, we estimate the "marginal cost of public funds" of the existing VAT in Chile to equal 7.6 percent. This implies that consumers and producers will have to be taxed 1,076 pesos for the government to receive 1,000 pesos. The 76 pesos are a welfare loss to the Chilean economy.

We also calculate the marginal cost of public funds of the Chilean tariff, which equals 18.5 percent. Despite the fact that the tariff is uniform across sectors, and

TABLE 2. Welfare and Government Revenue Results for Chile's Trade Policy Options

Policy simulation		With replacement taxes as: Existing VAT		Uniform VAT[a]		Lump sum	Combined effect of uniform VAT and trade policy[b]	
		% change in welfare[c]	% change in VAT[d]	% change in welfare[c]	Tariff revenue % of GDP	% change in welfare[c]	% change in welfare[c]	Tariff revenue % of GDP
		(1)	(2)	(3)	(4)	(5)	(6)	(7)
1. FTA with MERCOSUR	(central elasticities)	–0.62	45	–0.40	1.7	–0.43	–0.19	1.8
	(low elasticities)	0.04	17	0.07	2.7	0.08	0.19	2.7
2. Customs union with MERCOSUR	(central elasticities)	–0.95	52	–0.74	1.3	–0.73	–0.62	1.2
	(low elasticities)	–0.20	21	–0.22	2.5	–0.17	–0.14	2.5
3. FTA with NAFTA	(central elasticities)	0.82	48	1.03	0.9	1.04	1.23	0.9
	(low elasticities)	0.30	26	0.31	2.1	0.38	0.43	2.1
4. Zero tariffs on NAFTA imports,	(central elasticities)	–1.11	62	–0.92	0.7	–0.83	–0.64	0.7
no improved access	(low elasticities)	–0.47	30	–0.45	2.0	–0.41	–0.33	2.0
5. FTA with MERCOSUR and	(central elasticities)	0.12	49	0.44	1.7	0.35	0.61	1.7
6% external tariff	(low elasticities	0.06	38	0.11	1.7	0.13	0.21	1.7
6. FTA with NAFTA and	(central elasticities)	1.46	45	1.72	1.1	1.70	1.89	1.1
6% external tariff	(low elasticities)	0.41	41	0.45	1.4	0.49	0.55	1.4
7. Reduce external tariff to 8%	(central elasticities)	0.02	16	0.12	2.9	0.10	0.41	2.9
	(low elasticities)	–0.11	17	–0.08	2.7	–0.06	0.03	2.7
8. Reduce external tariff to 6%	(central elasticities)	0.01	28	0.16	2.3	0.11	0.43	2.3
	(low elasticities)	–0.18	30	–0.14	2.1	–0.14	–0.04	2.1
9. Reduce external tariff to zero	(central elasticities)	–0.26	76	0.02	0	0.09	0.21	0
	(low elasticities)	–0.54	72	–0.45	0	–0.42	–0.37	0

[a]In these scenarios we first create an equilibrium with a uniform VAT, no other domestic taxes, then evaluate the "pure" effects of the trade policy.

[b]These scenarios combine the impacts of the trade policy simulation with going to a uniform VAT and elimination of the domestic output tax, government revenues held constant.

[c]Percentage change in Hicksian equivalent variation as a percentage of GDP.

[d]Required equiproportional increase in the VAT rate across all sectors to keep government revenues unchanged.

therefore imposes no intersectoral distortion costs, the Chilean tariff imposes a higher distortion cost than the VAT because the tariff favors domestic production over imports.

Column 5 of table 2 shows the results of employing a lump-sum tax as the replacement tax. This tax avoids the distortions of a nonuniform VAT because consumer income is taxed in a fixed amount independently of consumer choices. Hence there are no resource allocation effects from the revenue replacement tax instrument. The results show that there is an added welfare cost of using the VAT, as compared with the lump-sum alternative.

Finally, column 3 of table 2 shows the results of using a uniform VAT. In these scenarios, we first counterfactually create an equilibrium in which all other domestic taxes and subsidies are zero and the VAT is uniform. The impact we evaluate is then solely due to the change in trade policy. Because all sectors are taxed and there is no labor-leisure choice, there is no way to take an action that will lower the tax. In other words, there are no resource allocation effects and the uniform VAT is essentially equivalent to a lump-sum or distortionless tax in our model. In addition, any "second-best" interaction effects of distortions between the tariff and the existing VAT will be removed if we start with a uniform VAT and no other distortions (for this reason, the results for the lump-sum tax and the uniform VAT may differ). In these scenarios, we equalize the VAT across sectors and solve for the level of the VAT that is required to compensate for the lost revenue.

REVENUE EFFECTS. Column 2 of table 2 presents the equi-proportional increase in the VAT required to keep government revenue constant. For example, with central elasticities, an FTA with MERCOSUR will require an increase of 45 percent in the VAT rate across sectors. Thus, if the collected VAT rate is 10 percent in a sector, the collected VAT rate will have to increase to 14.5 percent. With central elasticities, there is a strong substitution away from imports that pay tariffs in favor of imports from partner countries that are tariff free. In this case, the revenue requirements for the VAT are quite high to compensate for the lost tariff revenues. With low trade elasticities, the revenue requirement for the VAT is much smaller, ranging from increases between 17 and 26 percent in the three basic preferential trade arrangement scenarios presented in rows 1–3 of table 2.

In columns 4 and 7 of table 2, we show tariff revenues collected in the new equilibrium as a percentage of GDP. In our initial equilibrium, tariff revenues are equal to about 3.6 percent of GDP, but in the preferential trade area scenarios (rows 1–3), they fall to between 0.9 and 2.7 percent of GDP. Thus, in the preferential trade area scenarios, tariff revenues fall to between 25 and 75 percent of the original levels. The loss of tariff revenue is higher with NAFTA (because NAFTA is a larger share of Chilean imports than MERCOSUR) and higher with central elasticities because of the greater trade diversion. The VAT revenues as a percentage of GDP initially constitute about 9 percent of GDP. Depending on the preferential trade area and elasticities, the tariff loss is between 0.9 and 2.7 percent of

GDP. Hence, if the VAT is employed as the replacement tax, it will be necessary for VAT revenues to increase by about 10 to 30 percent.

Some may question whether the implied increase in the VAT is too high. To provide intuition for the model implications for the VAT, consider a particular scenario in which the lost tariff revenue is about 2.5 percent of GDP, as in row 6 of table 2 with central elasticities. Table 2 estimates that the VAT rate will have to increase by 45 percent to a legal rate of about 26 percent. In 1994, the legal VAT rate of 18 percent generated VAT revenues of about 9 percent of GDP, so the legal rate was twice the collected rate. Assuming no change in the rate of VAT evasion, it would appear necessary to raise the VAT by 5 percent to generate 2.5 percent of GDP (that is, from 18 to 23 percent).

The reason that the model predicts a required increase of the legal VAT rate to 26 percent and not 23 percent is that an increase in the tax will induce a shift away from the highly taxed sectors and an erosion of the tax base. Given our model parameters, increases in the VAT continue to generate additions in revenue within the range under consideration. But it is possible that evasion of the VAT could increase. The required legal VAT rate would then increase and the distortion costs of revenue replacement would be still higher than we have estimated. It is possible that the VAT is not a feasible tax for generating considerably more revenue without further reform in collection procedures.[16] Given the uncertainties over rates of evasion of VAT in Chile, these estimates should be taken as indicative of revenue requirements rather than as precise recommendations for the VAT rate. In fact, the next subsection emphasizes the importance of uniformity of collections.

Options for Preferential Trade Areas

RESULTS IN TABLE 2. Table 2 presents the overall welfare results for the trade policy options. Harrison, Rutherford, and Tarr (1997b) give more detailed results on output, imports, and exports for the main scenarios, with central elasticities. Welfare impacts are presented as a percentage of Chile's GDP. They represent changes on a recurring, annual basis, so a 1 percent welfare gain should be interpreted as a 1 percent increase in real income *each year in the future.*

The first row of table 2 presents the results from the scenario where Chile forms an FTA with MERCOSUR. It assumes that each of the MERCOSUR countries represented in the model, Argentina and Brazil, reduces its tariffs, export subsidies, or taxes on its trade with Chile to zero and that Chile does the same for its

16. To quantify these ideas, we simulated Chile's FTA with MERCOSUR and NAFTA, assuming that the services and trade and transportation sectors cannot have their collected VAT rates increased due to evasion. (These are the sectors with low rates of VAT collection and where evasion of the VAT may prevent additional collections; together they produce about 65 percent of Chilean value added.) With central elasticities, the welfare loss in this case from the FTA with MERCOSUR is increased to −0.60 percent of GDP and the gains from the FTA with NAFTA are reduced to 0.12 percent of GDP. As expected, the required rate of VAT increase jumps to about 75 percent.

trade with MERCOSUR. Chile does not adopt the common external tariff of MERCOSUR in this scenario.

The second scenario, shown in row 2 of table 2, represents Chile joining MERCOSUR as part of the customs union. In addition to the requirements of the scenario in row 1, in this case Chile adopts the common external tariff of MERCOSUR. Although Chile has not joined the MERCOSUR customs union, it is a potential policy option, so we evaluate it in this scenario. For simplicity, we assume that the common external tariff that Chile adopts is the import tariff structure that Brazil currently has with the countries that are not in MERCOSUR.[17]

In the third scenario, in row 3 of table 2, Chile forms an FTA with NAFTA. In row 4, primarily to help understand the results, we evaluate the consequences of a free trade agreement between Chile and NAFTA in which Chile does not obtain improved access to the NAFTA market. After discussion of these scenarios, we introduce further simulations to help explain the results and evaluate modified options.

The effects on welfare are dependent on both how Chile chooses to replace the lost tariff revenues and on assumed elasticities. Chile's preferential trade policy options with MERCOSUR lead to a loss of welfare with our preferred central trade elasticities and negligible gains or losses with low trade elasticities. With central trade elasticities, the trade diversion costs of an agreement with MERCOSUR typically dominate the trade creation effects. Moreover, based on the MERCOSUR external tariff, preferential access to the MERCOSUR markets is insufficient to overcome this welfare loss in Chile's markets. Welfare losses are lower with lower assumed elasticities because there is less trade diversion when Chile's consumers are less willing to substitute MERCOSUR's products for those of the rest of the world.[18]

The results indicate that the customs union with MERCOSUR is an inferior outcome for Chile compared with a free trade agreement with MERCOSUR. MERCOSUR's tariff structure is diverse compared with Chile's tariff, which is uniform. Because the welfare costs of trade restrictions tend to increase disproportionately with the height of the tariff, Chile is better off with its own uniform

17. This tariff structure is slightly different than the tariff structure shown for Argentina for two reasons. First, there are exceptions to the common external tariff for Argentina and Brazil, as both countries continue to adapt their tariff schedules over time to the agreed common external tariff. In addition, Argentina and Brazil could well have adopted exactly the same common external tariff at a detailed tariff line level, but have different trade shares across these tariff lines. With the different trade weights, the rates that appear in the GTAP database at the 24-sector level reflect differences in these trade patterns and need not reflect differences in the common external tariff at the detailed tariff line level. For ease of comparison, we also assume in our "Chile customs union with MERCOSUR" scenario that Argentina adopts the tariff of Brazil as its common external tariff. This provides a clean representation of the MERCOSUR customs union for our purposes.

18. These results are consistent with those of Donoso and Hachette (1996) and Muchnik, Errazuriz, and Dominguez (1996). Based on the latter's results, which focus on agriculture, Donoso and Hachette estimate that access to the MERCOSUR market would not offer significant gains to Chile. See also Valdes (1995).

tariff than with the common external tariff of the customs union.[19] That is, part of the costs to Chile of joining a customs union with MERCOSUR derive from the loss of tariff uniformity. Thus, one advantage of a free trade agreement for Chile as opposed to a customs union is that only the customs union requires the adoption of a common external tariff.

In comparing our results in rows 1–3 of table 2 regarding Chile's preferential trade area options, the most important result is that the FTA with NAFTA is beneficial to Chile, whereas the other options are likely to present problems.[20] To ascertain the source of the gain to Chile from a FTA with NAFTA, we performed the simulation in row 4, in which Chile lowers its tariffs against imports from NAFTA countries but does not obtain improved access in NAFTA markets. Although this is not a policy option that Chile would adopt, the results in row 4 show that Chile loses from preferential reduction of its tariffs against NAFTA countries without reciprocal access to NAFTA markets because the trade diversion dominates the trade creation.

To identify even more precisely the source of the access gains from the FTA with NAFTA, we performed a simulation in which access to only one sector was not obtained: nongrain crops. Our estimates of the tariff distortions suggest that the U.S. tariff is likely to be central in this sector: there is a 20 percent tariff on nongrain crops.[21] In other words, Chile applies zero tariffs against NAFTA imports, and NAFTA applies zero tariffs against imports from Chile in all sectors except nongrain crops. Although not shown in table 2, if Chile fails to obtain preferential access in nongrain crops, the welfare gains of 0.82 percent we ob-

19. "Ramsey optimal" tariffs will vary inversely with the elasticity of demand. Typically, however, departures from uniformity do not conform with Ramsey optimal rules, but rather with political economy considerations (see Panagariya and Rodrik 1993).

20. Coeymans and Larrain (1994), Reinert and Roland-Holst (1996), and Hinojosa-Ojeda, Lewis, and Robinson (1995) also find that Chile will gain from a FTA with NAFTA.

21. Although the GTAP database indicates that the U.S. tariff on nongrain crops is 47 percent, we have lowered this to 20 percent in our benchmark equilibrium for two reasons. First, we prefer updated estimates where possible. The most important nongrain crops for Chile are fruits and vegetables, and post–Uruguay Round tariff rates for these products in the U.S. market are the relatively modest figures cited in this note; the higher protection estimates for these products in the GTAP database (averaging 56 percent) were derived from an average of protection estimates in the 1989–94 period. Second, the U.S. protection on these products varies with the season. We have assumed that given production in the opposite hemispheres, when Chilean fruits and vegetables are ready for harvest and export to the United States, they would typically face U.S. tariffs that are in the low range of the seasonal tariffs applied by the United States. Products included in the nongrain crops category of the GTAP database (along with the estimated tariff and tariff equivalent of the nontariff barrier in the United States) are: sugar, 67 percent; oilseeds, including peanuts, 25 percent; coffee, cocoa, and tea, 0 percent; cotton, 31 percent; vegetables (fresh, 0–25 percent; frozen, 17.5–25 percent; dried, 25–35 percent, prepared and preserved, 13.6–14.7 percent); fruits (fresh, 0–20 percent; dehydrated, 0.6–2.2 percent; frozen, 0.7–14 percent; juices, 0–31.3 percent; jams and pastes, 7.0–35 percent; canned, 1.9–20 percent); and other nonfood crops (tobacco and jute), 19 percent. The reduced estimates are closer to the estimates of Butelmann and Meller (1995, p. 376). They report that Chilean fresh, frozen, and canned vegetables face MFN tariff rates in the United States ranging from 9.5 to 17.5 percent (with a few percentage point reductions for the former two categories where GSP treatment applies), and that Chilean fruits face U.S. MFN tariffs from 1 to 10 percent.

tained in the full-access case now drop to a welfare loss of 0.58 percent. Thus, access in nongrain crops is crucial to welfare gains from NAFTA.[22]

These results demonstrate the importance of improved access emphasized by Wonnacott and Wonnacott (1981). Our results show that Chile can gain more from an FTA with NAFTA than it can from global free trade. But Chile can expect to lose from any of the preferential trade agreements we have considered if there is no improvement in access to partner-country markets.

THE IMPORTANCE OF LOW UNIFORM TARIFFS. These results differ from several earlier numerical evaluations of preferential trading areas (Rutherford, Rutström, and Tarr 1997, Harrison, Rutherford, and Tarr 1997a). We speculate that part of the reason that trade diversion dominates trade creation in these estimates is that Chile has a low and uniform tariff. That is, the implementation of a preferential trade agreement in a country that starts with a dispersed tariff structure may result in a reduction in the dispersion of the tariff structure, although this is not true as a general proposition. Potential benefits from a reduction in the dispersion of the tariff, however, are ignored in more aggregated analyses of preferential trade arrangements.[23] To verify this intuition, we have counterfactually created an initial equilibrium in which Chile applies a 22 percent tariff on one-half of its imports and a zero tariff on all others, and then implemented the policy scenarios in rows 1–4 of table 2 (where we have employed existing VAT replacement and central elasticities). The sectors with high tariffs were selected at random, and the experiment was repeated 206 times. The means of the distributions for welfare as a percentage of GDP are as follows: FTA with MERCOSUR, –0.56 percent; customs union with MERCOSUR, –0.44 percent; FTA with NAFTA, 1.47 percent; and FTA with NAFTA but no improved access, –0.52 percent.

The gains of the FTA with NAFTA are significantly larger when based on the hypothetical nonuniform initial tariff structure. Similarly, the losses from the FTA with MERCOSUR are slightly smaller, reflecting a movement toward uniformity. But losses from a preferential reduction in tariffs toward the NAFTA markets remain unless access to the NAFTA market is obtained. These numerical results are consistent with the theoretical results of Hatta (1977), who found that countries will benefit from moving toward uniformity by simultaneously lowering the highest tariff and raising the lowest tariff.

22. Because U.S. protection in milk products is also high, we examined the impact of denial of improved access in NAFTA markets for Chilean products on both nongrain crops and milk products. Chile exports very little milk products, however, so the welfare result was only slightly more adverse for Chile (–0.60 percent of GDP with central elasticities and existing VAT replacement) relative to denial of Chilean access in nongrain crops alone.

23. There is value in further theoretical work into the generality of the impact of preferential arrangements on uniformity. Note also that in our model, elasticities are equal across sectors, so the Ramsey optimal tariff is uniform. A useful exercise would be to evaluate the impact of a preferential trade arrangement where we start from randomly selected elasticities across sectors and see how often Chile gains from preferential trade agreements as we use a large number of distinct sets of elasticities.

In this hypothetical experiment, we find that the ranking of the customs union with MERCOSUR versus the FTA with MERCOSUR is reversed compared with the actual situation represented in table 1. Although Chile still loses from both preferential trade agreements with MERCOSUR, it is intuitive that the customs union produces smaller losses than the FTA because the common external tariff of MERCOSUR is more uniform than the hypothetical Chilean tariff. In the actual situation of table 2, the customs union with MERCOSUR represents *a movement away from uniformity.*

Optimizing Chile's Trade Policy Options

We know from theory that Chile can reduce the trade diversion costs of preferential trade areas if it lowers its external tariff. Thus, a number of economists have recommended that a reduction in Chile's external tariff be combined with its free trade agreements.[24] In rows 5 and 6 of table 2, we evaluate the two FTA options with a simultaneous reduction in the tariff to 6 percent. In rows 7 and 8, we evaluate the impact of lowering the external tariff to 8 and 6 percent, respectively, on a multilateral basis. We evaluate going to global free trade in row 9.

Chile may have a low optimal tariff despite being a small country. If Chilean exports are differentiated from the products of other countries so that Chile in aggregate faces a downward-sloping demand curve for a product, even if individual Chilean producers do not perceive a downward-sloping demand curve, then there will be an optimal export tax to maximize Chilean export profits. The height of the optimal export tax will be inversely related to the elasticity of demand faced by Chile in its export markets, which is in turn determined by how substitutable Chile's products are with those of other countries.[25] In the limit, when Chilean products are perfect substitutes in all its export markets for products from all other countries, Chile has no ability to obtain a higher price by restricting its exports. In this case, the optimal export tax is zero.

Although Chile imposes virtually no export taxes, the Lerner symmetry theorem tells us that in general equilibrium import tariffs are equivalent to export taxes. The import tariff will tax all export sectors roughly uniformly. However, with product differentiation and many sectors, market power on exports differs across sectors and destination markets. Hence the import tariff is not as efficient an instrument as export taxes varying by sector and destination. Nonetheless, if export taxes are ruled out, there is a positive optimal import tariff. However, given the existence of a uniform tariff of 11 percent, we investigate both theoretically and numerically whether the optimal tariff is above or below it.

In our central elasticity scenarios, we have assumed that all countries have an elasticity of substitution between imports from different countries (σ_{MM}) equal

24. Such as Schiff and Sapelli (1996), Corbo (1966), and Leipziger and Winters (1996).

25. Individual competitive firms will price at their marginal costs, but because the country as a whole must accept a lower price to sell more, there is an optimal export tax that equates the marginal revenue received from exports equal to the marginal costs. The more elastic the demand, the lower the optimal export tax.

to 30. We show in Harrison, Rutherford, and Tarr (1997b, appendix C) that the optimal tariff t^* is bounded below by $t^* = \{[\sigma_{MM}/(\sigma_{MM} - 1)] - 1\}$. Thus, even with $\sigma_{MM} = 30$, the optimal tariff is over 3 percent; but in our low elasticity scenarios, with $\sigma_{MM} = 8$, the optimal tariff is over 14 percent.

Considering the preferential trade options in rows 5 and 6 of table 2, there is an expected increase in the estimated welfare gains compared with rows 1 and 3, respectively. With central elasticities, there is a significant improvement in welfare compared with an external tariff of 11 percent. With low elasticities, the adverse terms-of-trade effect of reducing tariffs mitigates the welfare gain from reducing the trade diversion costs. These results show that as long as Chile limits itself to an FTA, it can profit from the increased access it obtains in its partner countries without excessive trade diversion costs, provided it lowers its external tariff sufficiently. In particular, the results in row 5 show that the free trade agreement with MERCOSUR can be expected to yield benefits when the external tariff is lowered to 6 percent. By contrast, comparing rows 5 and 6, we observe that an agreement with NAFTA is worth a lot more than the one with MERCOSUR, largely due to the superior market access of NAFTA.

Rows 7 and 8 of table 2 present estimates of the welfare and replacement tax implications, respectively, to Chile of unilaterally lowering its external tariff. With central elasticities and distortionless domestic taxes (lump sum or uniform VAT), unilateral reduction of the tariff to 6 percent increases welfare, and there are further gains from reducing tariffs from 8 to 6 percent. With the existing VAT as the replacement tax, reducing the tariff to 8 percent increases welfare. However, the distortion costs of the VAT are sufficiently close to the tariff, so that in combination with the small adverse terms-of-trade effects, there are no further gains from tariff reduction below 8 percent. With a distortionless replacement tax, reduction of the external tariff to zero produces positive welfare gains compared with the tariff of 11 percent (row 9). However, because the gains are less than reduction to 6 percent (row 8), the optimal tariff is between 0 and 6 percent.[26]

With existing VAT replacement, there is some limited scope for beneficial reduction of the tariff with central elasticities. Again, with higher elasticities, the optimal tariff is lower and the gains from tariff reduction would increase.

Sectoral Impacts

Tables 6 and 7 in Harrison, Rutherford, and Tarr (1977b) present the impacts under central elasticities on output, exports and imports at the 24-sector level from three of the principal trade policy options: the FTA with MERCOSUR, the FTA with

26. These were the results employed by the vice president of the Central Bank of Chile in his presentation before the lower house committee of the Chilean Parliament when he argued for a reduction of the tariff to 6 percent. In fact, we have separately calculated the optimum tariff with central elasticities at between 3 and 4 percent, and with low elasticities of about 14 percent, assuming lump-sum replacement of tariff revenues in each case.

NAFTA, and unilateral reduction of the tariff to 8 percent.[27] Focusing on the percentage change in output with central elasticities, the sectors that significantly expand under the free trade agreement with MERCOSUR are transportation equipment (dramatically),[28] machinery and equipment, iron and steel, and milk. With the free trade agreement with NAFTA, the sectors that expand more than 10 percent are iron and steel, transportation equipment, milk, nongrain crops, and textiles. With unilateral tariff reduction, the expanding sectors are transportation equipment, iron and steel, and, to a lesser extent, nonferrous metals and mining.

Iron and steel and transportation equipment expand under all three trade policy options, but the other expanding sectors differ. Iron and steel and transportation equipment are both small sectors in Chile; each sector produces less than 1 percent of Chilean value added. However, these are the two sectors that export the most intensively: both export over 90 percent of their output. Preferential or multilateral tariff reduction induces a depreciation in the real exchange rate, which makes exporting more profitable and gives a boost to the sectors that export intensively.

With unilateral tariff reduction, the other sectors that expand (nonferrous metals and mining) are also the ones that export a high percentage of their output. So the real exchange rate impact and export intensity explain the pattern of expanding and contracting sectors with unilateral nondiscriminatory tariff reduction.

With a free trade agreement with NAFTA, textiles, milk, and nongrain crops expand—in addition to the two or three most export-intensive sectors—because these three sectors obtain a substantial improvement in their terms of trade in the U.S. market. We considered earlier how improved access to nongrain crops and milk is crucial to an improvement in Chilean welfare from NAFTA, and these sectoral results are consistent with those welfare results.

With the free trade agreement with MERCOSUR, machinery and equipment and milk expand in addition to transportation and iron and steel. Our data indicate that these sectors are two of the most highly protected in MERCOSUR, so they obtain relatively greater improvement in their terms of trade after implementation of a free trade agreement with MERCOSUR, which induces their expansion.

Additive Regionalism

Butelmann and Meller (1995) articulate the strategy of the government of Chile: to negotiate bilateral free trade agreements with MERCOSUR, NAFTA, and all of its significant and willing trading partners, including the European Union and the rest of South America.[29] They argue that this strategy will progressively lower

27. We also present the sectoral results with low elasticities.

28. Although the expansion is dramatic in percentage terms, it starts from a very small base. Thus, the absolute increase is plausible.

29. The percentage share of Chile's aggregate exports (imports) for its most significant trading partners are: the United States, 14 (25); Brazil, 5 (7); Argentina, 5 (6); the European Union, 32 (23); the rest of South America, 5 (5); and Japan, 17 (10).

the effective average tariff, successively reduce trade diversion costs, and, crucially, will help ensure stability of access to the markets of partner countries. The free trade agreement in late 1996 between Chile and Canada, in which both countries agreed to eschew antidumping actions against each other, is regarded as a notable example of the advantages that the bilateral approach offers. An opposing view within Chile is offered by Donoso and Hachette (1996). They argue that the limited market access of the bilateral agreements with the Southern countries (for example, MERCOSUR) is not worth delaying the benefits of opening up unilaterally, although agreements with the large markets of the United States, the European Union, or Japan would be worthwhile. Moreover, they fear that the MERCOSUR arrangement may restrict broader liberalization.

Table 3 presents estimates of the gains to Chile of progressively adding free trade agreements, using central elasticities and a lump-sum tax as the replacement tax. Columns 1 and 2 reproduce the estimates in table 2. Column 3 shows that although the MERCOSUR agreement independently results in losses to Chile, when combined with an agreement with NAFTA, the impact of an agreement with MERCOSUR is positive rather than negative. The reason is that competition from NAFTA producers greatly reduces the extent and impact of trade diversion.[30] Column 4 of row 1 shows that combining agreements with NAFTA and MERCOSUR with an agreement with the European Union results in a large increase in the gains to more than 5 percent of GDP. Finally, adding a free trade agreement with the rest of South America results in gains of 8.4 percent of GDP. These are enormous estimated gains for a constant-returns-to-scale model. The last column of row 1 excludes the United States from the agreement, but this has only a small negative impact on Chile because it obtains such substantial preferential access in the other markets.

Critics of the government's strategy argue that it is unrealistic to assume that the European Union would grant tariff-free access in its highly protected agricultural products as part of a free trade agreement with Chile. The European Union has steadfastly refused to do so in its Association Agreements with the

30. NAFTA and MERCOSUR combined produce gains of 1.48 percent of GDP, whereas if the results of the NAFTA and MERCOSUR agreements were merely additive (columns 1 and 2) the gains would be only 0.61 percent of GDP. That is, we find that reduced trade diversion from the combined agreements accounts for 0.87 percent of GDP. Because this may appear to be too large a saving due to reduced trade diversion, to verify our explanation we have three additional simulations: (1) Chile unilaterally eliminates tariffs on NAFTA imports without improved access to NAFTA; (2) Chile unilaterally eliminates tariffs on MERCOSUR imports without improved access to MERCOSUR; and (3) Chile unilaterally eliminates tariffs on NAFTA and MERCOSUR without improved access to NAFTA or MERCOSUR markets. If our explanation is correct, simulation 3 should result in reduced trade diversion costs of at least 0.87 percent of GDP, compared with additive losses from the first two simulations. In percentage of GDP, the welfare impacts of these three simulations are: (1) –0.83, (2) –0.82, and (3) –0.77. If the losses of the preferential tariff reduction were additive, the total losses would be –1.65 (= –0.83 – 0.82). Because preferential tariff reduction against the two regions combined results in losses of only –0.77 percent of GDP, trade diversion costs are reduced by 0.88 percent of GDP by combining tariff reductions for the two regions.

TABLE 3. Welfare Results of Additive Free Trade Agreements by Chile
(Chilean gains as a percent of Chilean GDP with central elasticities and lump-sum tax replacement)

	Agreements with					
Product coverage	MERCOSUR	NAFTA	NAFTA & MERCOSUR	NAFTA & MERCOSUR & EU	NAFTA & MERCOSUR & EU & rest of SA[a]	Canada & Mexico MERCOSUR & EU & rest of SA[a]
	(1)	(2)	(3)	(4)	(5)	(6)
1. All products included	–0.43	1.04	1.48	5.24	8.4	8.16
2. Excluded products[b]	–0.43	1.04	1.48	2.02	2.48	0.44
3. Excluded products[b] and 6% external tariff	0.35	1.70	2.01	2.29	2.66	0.87
4. Only EU AG products[c] excluded	–0.43	1.04	1.48	2.02	5.48	3.90
5. Only EU AG products excluded and 6% tariff	0.35	1.70	2.01	2.29	5.71	4.44

[a]Rest of SA is South America except for Chile and the MERCOSUR countries.
[b]Excluded products in the agreement with the EU and their tariffs plus nontariff equivalents in the EU are: wheat (57%), grains (74%), nongrain crops (51%), fishing (14%), meat (63%), and milk (129%). Excluded products in the agreement with the rest of South America (and their tariffs) are: nongrain crops (29%), meat (51%), milk (27%), food (34%), beverages and tobacco (55%), textiles and apparel (46%), chemicals, rubber, and plastics (31%), fabricated metal products (43%), and machinery (52%).
[c]Only the agricultural products from the European Union listed in note b are excluded from any of the FTAs.
Source: Authors' calculations.

Central and Eastern European countries and in its Free Trade and Customs Union Agreements with Mediterranean countries, such as Morocco, Tunisia, and Turkey. Hence, it is unlikely to offer concessions to Chile that it has refused to offer to other countries for which it might be viewed as having more to gain geopolitically. Similarly, although more speculatively, it would be doubtful that tariff-free access in the most highly protected products would be provided by the rest of South America because (following Grossman and Helpman 1995) the political-economy interests that obtained such high protection would resist regional competition as well.

Row 2 of table 3 presents results that more realistically reflect possible outcomes due to excluded products. They exclude agricultural products from the agreement with the European Union, and products with tariffs above 25 percent in the rest of South America from that agreement. The results show, as expected, that without preferential access to these highly protected markets, the gains would be dramatically reduced. The last column shows that the United States is crucial to the whole scenario. If the United States is not included in the additive agreements, the gains drop dramatically to 0.44 percent of GDP. The drop in welfare for Chile exceeds the gains from NAFTA alone, showing that competition from (and in) the United States is important to Chile being able to avoid the trade diversion costs of these agreements. Conversely, if Chile can get a free trade agreement with the United States as part of NAFTA, then free trade agreements with MERCOSUR, the European Union, and the rest of South America each add, impressively, about 0.5 percent to Chilean GDP. These gains accrue even when the European Union and the rest of South America exclude their most highly protected items from the agreements.

Proponents of the government's strategy maintain that the trade diversion costs of the free trade agreements will be diminished because Chile will adopt an external tariff of 6 percent. Moreover, though they concede that access to the European Union in agricultural products is unlikely, they maintain that it is possible that Chile will receive full access to the markets of the rest of South America in view of the sustained trend toward open economies in Latin America. In row 3 of table 3, we evaluate the impact of a 6 percent external tariff with the same products excluded from the agreements with the European Union and the rest of South America as in row 2. There are slightly larger gains to Chile from lowering the external tariff, but the United States remains important for substantial gains. In rows 4 and 5, we evaluate additive regionalism where only European Union agricultural products are excluded, so that full access to the rest of South America is obtained. Columns 5 and 6 show that Chile obtains very substantial gains, with a 6 or 11 percent external tariff, if it can obtain tariff-free access to the highly protected markets of the rest of South America.

Thus, if Chile succeeds in including a wide net of countries in its additive regionalism strategy, the estimates of the welfare gains range from 0.44 to 8.4 percent of Chilean GDP. However, table 2 shows that the gains to Chile from unilateral trade liberalization are only about 0.11 percent of GDP. Hence, the estimated gains to Chile from additive regionalism are 4 to 76 times the gains from unilateral trade liberalization. On balance, it appears that Chile has little to lose by pursuing additive regionalism, especially given that additive regionalism is being combined with lowering the external tariff to about 6 to 8 percent.

III. The Impact of Additive Regionalism

Experience with regional trade arrangements has shown that if the agreement is not mutually beneficial to all parties, then it is unlikely to be effectively implemented or sustained (World Bank 2000). Agreements may exist de facto but are not implemented effectively. Thus, the impact on Chile's partner countries in the trade agreements is relevant to the likely success of the strategy in the long run. Moreover, even if the agreements are beneficial to Chile and its partners, if the benefits are derived from losses to countries that are excluded from the agreements, then clearly the agreements would be unattractive from the perspective of the multilateral trading system. Thus, it is important to estimate the impact on partner and excluded countries from the Chilean strategy of additive regionalism, and to assess the impact on the world in general. As a point of comparison, we also estimate the impact of global free trade on the countries and regions of our model.

Table 4 reports welfare gains as a percentage of own-country GDP, for both our central and low-elasticity cases. For comparisons of gains and losses across countries, table 5 presents the estimated welfare gains in millions of 1995 U.S. dollars. The first five columns in row 1 of table 5 reproduce the results for Chile's additive regionalism strategy, which is presented in the first five columns of table 3. Rows 2–11 present results for the other 10 countries or regions in our model. Column 6 presents results for the global free trade scenario.

Impact on Individual Countries and Regions

From the first five columns of table 4, Chile is too small or its trade pattern is sufficiently different for its regional agreements to have more than a trivial impact on about half of the countries and regions in the model.[31] This group includes Japan and the rest of the world (which are excluded from all the agreements evaluated in table 3) and the United States and the European Union (which are excluded in some of the arrangements in table 3 and included in others). Canada is also essentially unaffected by Chile's trade policy options.

The rest of South America and Central America lose in all the agreements from which they are excluded, although the welfare loss is only about 0.05 percent of their GDP. These regions compete with Chile for the markets in MERCOSUR and NAFTA and compete with producers from MERCOSUR and NAFTA for the Chilean market. In both cases, they lose access to markets because there is a decline

31. When we round welfare to the nearest 0.01 percent of GDP, the impact is either 0 or 0.01 percent.

TABLE 4. The Welfare Impact of Chile's Additive Free Trade Agreements and Global Free Trade[a]
(welfare gains as a percent of each country's GDP)

		Agreements with					
Country	Elasticity	MERCOSUR	NAFTA	NAFTA & MERCOSUR	NAFTA & MERCOSUR & EU	NAFTA & MERCOSUR & EU & Rest of SA[b]	Global free trade
		(1)	(2)	(3)	(4)	(5)	(6)
1. Chile	central	–0.40	1.04	1.48	5.24	8.40	1.26
	(low)	(0.00)	(0.37)	(0.60)	(2.55)	(3.31)	(0.68)
2. United States	central	0.00	0.00	0.00	0.00	0.00	0.34
	(low)	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)	(0.18)
3. Canada	central	0.00	0.00	0.00	0.00	0.01	0.42
	(low)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(–0.36)
4. Mexico	central	0.00	–0.02	–0.01	0.00	0.00	–1.38
	(low)	(0.00)	(–0.01)	(–0.01)	(0.00)	(0.00)	(–1.02)
5. Argentina	central	0.06	0.00	0.10	0.12	0.07	0.82
	(low)	(0.00)	(–0.01)	(0.02)	(0.02)	(0.01)	(0.60)
6. Brazil	central	0.02	–0.01	–0.04	–0.04	–0.02	0.94
	(low)	(0.00)	(–0.01)	(0.00)	(0.00)	(–0.01)	(0.24)
7. Central America	central	0.00	–0.06	–0.05	–0.04	–0.06	9.70
	(low)	(0.00)	(–0.03)	(–0.03)	(–0.05)	(–0.06)	(4.42)
8. Rest of SA	central	0.00	–0.03	–0.06	–0.04	–1.19	4.40
	(low)	(0.00)	(–0.02)	(–0.04)	(–0.05)	(–0.22)	(1.25)
9. European Union	central	0.00	0.00	0.00	0.00	0.00	2.74
	(low)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(1.17)
10. Japan	central	0.00	0.00	0.00	0.00	0.00	3.43
	(low)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(1.98)
11. Rest of the world	central	0.00	0.00	0.00	0.00	0.00	1.97
	(low)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.54)

[a]All products included in agreements and lump-sum tax replacement.
[b]Rest of SA is South America except for Chile, Argentina, and Brazil.
Source: Authors' calculations.

TABLE 5. The Welfare Impact of Chile's Additive Free Trade Agreements and Global Free Trade[a]
(welfare gains in millions of 1995 US dollars)

Country	Elasticity	Agreements with					
		MERCOSUR	NAFTA	NAFTA & MERCOSUR	NAFTA & MERCOSUR & EU	NAFTA & MERCOSUR & EU & Rest of SA[b]	Global free trade
		(1)	(2)	(3)	(4)	(5)	(6)
1. Chile	central	−291	414	590	2090	3350	504
	(low)	(−67)	(149)	(239)	(1013)	(1318)	(270)
2. United States	central	−7	51	−29	138	60	19,972
	(low)	(−24)	(306)	(231)	(59)	(−11)	(10,833)
3. Canada	central	5	−20	−22	23	49	243
	(low)	(4)	(−15)	(−13)	(14)	(19)	(−2058)
4. Mexico	central	13	−58	−44	−11	15	−4539
	(low)	(1)	(−35)	(−35)	(−3)	(0)	(−3315)
5. Argentina	central	63	−1	222	264	147	1832
	(low)	(44)	(−18)	(54)	(54)	(28)	(1327)
6. Brazil	central	214	−42	−171	−161	−70	3912
	(low)	(108)	(−36)	(15)	(−11)	(−21)	(1004)
7. Central America	central	4	−37	−32	−23	−38	6112
	(low)	(3)	(−21)	(−21)	(−29)	(−36)	(2680)
8. Rest of So. America	central	−34	−56	−95	−73	−2024	7456
	(low)	(−28)	(−39)	(−75)	(−90)	(−376)	(2110)
9. European Union	central	−184	−156	−336	−88	−200	207,413
	(low)	(−28)	(−241)	(−317)	(156)	(86)	(88,720)

10. Japan	central	−58	−19	−30	81	−2	127,664
	(low)	(−30)	(−48)	(−69)	(−76)	(−91)	(73,711)
11. Rest of the world	central	92	−73	−50	−115	6	85,111
	(low)	(29)	(−89)	(−100)	(−229)	(−232)	(23,348)
12. Sum for included[c]	central	−14	387	546	2255	1327	
	(low)	(85)	(405)	(491)	(1282)	(1043)	
13. Sum for excluded[d]	central	−169	−384	−543	−130	−34	
	(low)	(−73)	(−492)	(−582)	(−424)	(−359)	
14. Sum over all countries	central	−183	3	3	2125	1293	455,680
	(low)	(12)	(−87)	(−91)	(858)	(684)	(198,626)

[a]All products included in agreements and lump-sum tax replacement.
[b]Rest of SA is South America except for Chile, Argentina, and Brazil.
[c]Sum of the welfare impact for countries included in the agreement.
[d]Sum of the welfare impact for countries excluded from the agreement.
Source: Authors' calculations.

in the demand for their exports due to preferential access arrangements between Chile and its partners, which adversely affects their terms of trade and welfare.[32]

Perhaps most interesting is that although the rest of South America loses from being excluded by Chile, the biggest loss for this region by far is when the rest of South America *is included* with Chile in a free trade agreement (along with the European Union, NAFTA, and MERCOSUR, as shown in column 5 of table 4). The rest of South America has high protection on the products mentioned in the footnotes for table 3. To the extent that Chilean imports displace imports from other countries in the rest of South America, it loses tariff revenue on imports. Although there is some trade creation from tariff-free access to Chilean imports in the rest of South America, the tariff loss dominates the trade creation due to the high level of the tariffs.[33] Moreover, comparing columns 4 and 5 in table 5, the addition of the rest of South America to the coalition of Chile, MERCOSUR, the European Union, and NAFTA results in an aggregate reduction in welfare for the partner countries (see row 12). The gains to the other partners in this agreement are less than the losses to the rest of South America. So there are insufficient benefits to allow the gainers to compensate the rest of South America for its losses.

For Mexico, competition from Chile for preferred access in the U.S. market results in a very small negative impact of including Chile in NAFTA. However, Chile is too small to make a significant difference to Mexico in the U.S. market. When Chile combines an agreement with NAFTA with an agreement with MERCOSUR, the diversification of Chilean exports results in still less displacement of Mexican exports in the United States, so the negative impact on Mexico of Chile in NAFTA is reduced. When Chile adds the European Union to its group of free trade agreements, the diversification of Chilean exports reduces the small negative impact on Mexico of Chile's preferential access to the United States to virtually zero. By contrast, in the global free trade scenario discussed below, Mexican losses are substantial due to the erosion of preferential access in U.S. markets from the whole world.

Brazil and Argentina both lose from Chile joining NAFTA due to erosion of preference margins in both Chile and NAFTA markets. But Argentina and Brazil both gain small amounts from a MERCOSUR free trade agreement with Chile. The latter fact is partly explained by improved access to the Chilean market for MERCOSUR producers. It is also likely that part of the explanation for this result is that Brazil and Argentina reduce the trade diversion costs of MERCOSUR when they add new partners. That is, Chile will compete with Brazilian producers in Argentine markets. This will reduce the trade diversion costs of Argentina from importing Brazilian products under the MERCOSUR agreement. Of course, Chile could well displace imports from the rest of the world in Argentine markets, which

32. This is consistent with the evidence of Winters and Chang (2000). They find that the price of imports from the United States and Korea in Brazil fell after the formation of MERCOSUR.

33. If the high-tariff products are excluded from the free trade agreement with Chile, the losses are reduced to about one-third of their level (to –0.36 percent).

could increase Argentine trade diversion costs. But as more countries are added to a network of preferential trading arrangements, the trade diversion costs associated with earlier partners are reduced, especially if these are large countries that interject significant competition.[34] Comparing columns 4 and 5 in table 5, Brazil and Argentina both lose from Chile negotiating a free trade agreement with the rest of South America. This is likely due to a terms-of-trade loss in the markets of the rest of South America.

Aggregate Impact of Chile's Additive Regionalism Strategy

Even if Chile gains from an agreement or set of agreements, there is the question of whether Chile gains only because other countries lose. In table 5, we convert the percentage gains and losses of table 4 into gains and losses in millions of 1995 U.S. dollars. This allows us to compare gains and losses across countries and arrive at a total for the world. Row 12 sums the welfare effects for countries that are included in the agreement. For example, Chile-MERCOSUR (column 1) includes Chile, Argentina, and Brazil in our model. Row 13 sums the welfare effect for all countries that are not part of the agreement (for example, all countries other than Chile, Argentina, and Brazil in the case of Chile-MERCOSUR). Row 14 sums over all countries.

From row 12 in table 5, trade diversion dominates the Chile-MERCOSUR agreement to the extent that even the members of the agreement lose in the aggregate. But this agreement is the only one we consider that results in losses for the member countries. Other agreements considered in table 5 are "North–South" agreements (in particular, all include the United States), and we estimate that all of these result in aggregate net benefits for the member countries, although at least one member loses in all of them. The inclusion of the United States means that significant competition is injected into the markets of participating members, and this reduces the likelihood of trade diversion dominating.

From row 13 in table 5, all of the preferential arrangements we consider result in losses for the excluded countries or regions. These results are consistent with the results of Winters and Chang (2000). Employing ex post data, they show that there can be a very significant negative welfare effect (through negative terms-of-trade effects) on countries excluded from regional arrangements. In particular, they estimate that MERCOSUR induced losses for the United States, Germany, Japan, the Republic of Korea, and Chile of about $800 million per year, which was about 9 percent of the value of their exports to MERCOSUR.[35]

For the world as a whole, with central elasticities, the agreement with MERCOSUR results in losses of $183 million, primarily due to the trade diversion costs for Chile and the terms-of-trade loss for the European Union. Independent of elas-

34. It is possible, however, that a new partner could divert imports from an excluded country and add to the trade diversion costs on balance.

35. We estimate a very small negative effect for Central America as a result of Chile forming an FTA with NAFTA.

ticities, the agreements in the first three columns result in essentially a zero impact for the world or for the three excluded regions outside of the Western Hemisphere (rounded to the 0.01 percent of their own GDP). With NAFTA involved, Chile has significant gains, but the terms-of-trade loss for the excluded countries is almost as much as the gains to Chile, so the impact on the world is small.

In columns 4 and 5 of table 5, the gains for the world become significant when the European Union is added or when the European Union and the rest of South America are added to Chile's network of agreements. The main reason for the much larger gains to the world is that the gains to Chile become very large when it obtains preferential access to the markets of the European Union and the rest of South America. As explained, given the high protection on selected products in the rest of South America, the trade diversion costs in this region significantly reduce the gains to the world from including this region in Chile's network of free trade agreements.

Impact of Global Free Trade

The results for global free trade are presented in column 6 of tables 4 and 5. As expected, the gains to the world vastly exceed the gains from any regional arrangement. Even the included countries to any agreement gain more from multilateral global free trade than any individual regional arrangement (although the impact on Chile of an agreement with NAFTA is close). These results emphasize the importance of moving toward lower trade barriers in the multilateral context. Mexico is an exception (as is Canada in the low-elasticity case). Mexico sees losses from global free trade due to the erosion of favored access to the U.S. market.

IV. Conclusions

Our results for Chile point to some general themes regarding regional trading arrangements. One clear theme is that improved market access in preferential trading areas is important. In the case of Chile, trade diversion costs dominate the welfare effects of bilateral agreements unless sufficient market access is obtained in partner countries (or third-country tariffs are lowered). The North–South agreements generally provide sufficient access to make them beneficial, but the South–South agreement we examined did not (although Chile could lower its external tariff to make the South–South arrangement beneficial). We show that efficient replacement taxes are important with changes in either regional or unilateral trade policy and provide greater scope for trade policy action. Finally, our range of estimates for the gains from additive regionalism indicate that Chile has little to lose by pursuing this strategy and may potentially gain many multiples of the gains from unilateral trade liberalization.

We find that the excluded countries lose from all of the regional arrangements that we examine. In addition, partners to these preferential arrangements sometimes lose.

Chile's additive regional arrangements have an almost imperceptible impact on world welfare. However, we estimate that global free trade generates gains to the world that are enormous in comparison, emphasizing the importance of moving toward lower trade barriers in the multilateral context.

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Trade in International Maritime Services: How Much Does Policy Matter?

Carsten Fink, Aaditya Mattoo, and Ileana Cristina Neagu

Maritime transport costs significantly impede international trade. This article examines why these costs are so high in some countries and quantifies the importance of two explanations: restrictive trade policies and private anticompetitive practices. It finds that both matter, but the latter have a greater impact. Trade liberalization and the breakup of private carrier agreements would lead to an average of one-third lower liner transport prices and to cost savings of up to US$3 billion on goods carried to the United States alone. The policy implications are clear: there is a need not only for further liberalization of government policy but also for strengthened international disciplines on restrictive business practices. The authors propose an approach to developing such disciplines in the current round of services negotiations at the World Trade Organization.

Maritime transport costs have a profound influence on international trade. In many cases, their trade-inhibiting effect dwarfs that of customs duties.[1] For instance, the average incidence of transport cost exceeds that of tariffs on imports from the majority of U.S. trading partners (figure 1). More generally, economic research highlights the role of transport costs in determining geographical patterns of trade, production, industrial structure, and income (Venables and Limao 1999). Interesting new work even suggests that transport costs (as an element of trade costs) help explain a variety of puzzles in the field of international macroeconomics, such as the well-known home biases in consumption and investment and the excessive volatility of exchange rates (Obstfeld and Rogoff 2000). These observations are interesting from a policy point of view, however, only if something can be done about these costs. Are transport costs exogenously determined by technological developments or can they be influenced by policy?

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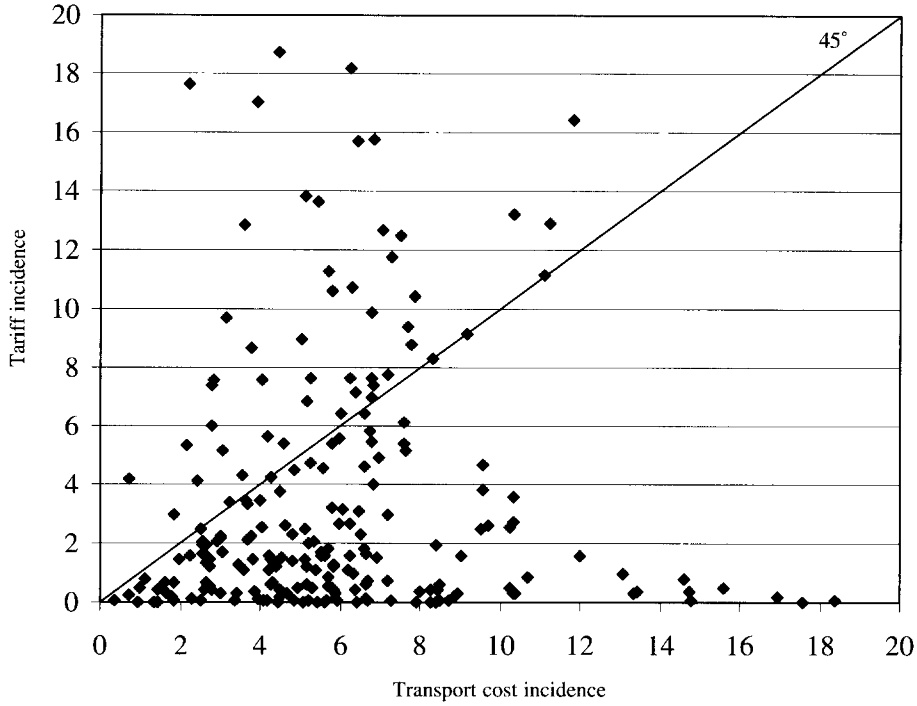
1. This has been demonstrated in several studies. See Waters (1970), Finger and Yeats (1976), Sampson and Yeats (1977), Conlon (1982), and Amjadi and Yeats (1995).

FIGURE 1. The Relative Importance of Transport Costs and Tariffs for U.S. Imports, 1998



Note: Each dot represents a country. The tariff incidence is calculated as the ratio of actual duties paid over import values. Similarly, the transport cost incidence represents the share of transport charges in import values. Five countries (Benin, Guinea, the Solomon Islands, Togo, and Samoa) exhibit a transport cost incidence greater than 20 percent and are not shown.
Source: U.S. Bureau of the Census.

Some researchers argue that restrictive trade policies keep maritime transport costs high, notably the cargo reservation schemes and monopoly rights granted to providers of port and auxiliary services (Bennathan 1989, Amjadi and Yeats 1995, Francois and others 1996; Hummels 1999). Some also argue that private anticompetitive practices—primarily but not exclusively of the maritime conferences—are responsible for keeping costs high (Francois and Wooton 1999, Hummels 1999). However, most observers also argue that both public and private trade-restrictive policies are becoming less important (White 1988, Franck and Bunel 1991, World Trade Organization [WTO] 1998). Yet the available evidence suggests that transport costs, especially for liner trade, are not falling—despite dramatic improvements in technology, especially in the form of containerization (Hummels 1999).

This article seeks to assess the relative importance of public and private trade-restrictive actions in explaining the price of maritime transport services. To measure these prices, we use newly published data on U.S. waterborne transport from the U.S. Department of Transportation. A major advantage with these data is that they are broken down by type of service—liner, bulk, and tanker. It is more difficult to put together a comprehensive data set on public policies and private practices, a problem that has inhibited meaningful empirical research in this area. The few attempts to measure the restrictive impact of government policies have only limited coverage (McGuire, Schuele, and Smith 2000), and there has not been, as far as we know, an attempt to use existing information on carrier agreements.[2] This article draws on a database, created as part of the World Bank's services project, which contains information on both policy and private rate-fixing arrangements affecting maritime trade with the United States.

These data made it possible to carry out the econometric analysis presented here. Our estimates confirm, first of all, the importance of all the standard determinants of transport prices, ranging from distance to technology. More interesting, we find that both public policy and private practices continue to exercise a significant influence on maritime transport prices. Somewhat surprisingly, private anticompetitive practices seem to have a stronger influence on prices than public restrictions.

What are the implications for policy? The negotiations on maritime transport were the only post–Uruguay Round services negotiations that completely failed. This failure implied an unfortunate loss of political momentum for reform of domestic policies and, less obviously, a lost opportunity to develop procompetitive rules. To some extent, an effort was made to develop rules that would ensure nondiscriminatory access to port services.[3] But these rules, concerned primarily with ensuring market access, did little to protect consumers from the anticompetitive practices of international cartels. An international initiative is needed because these practices cannot be adequately addressed only through national competition policy, given the weak enforcement capacity of small states. A further reason for developing a first-best international response to these practices is to prevent recourse to an inferior national response: recall that the cargo-sharing schemes imposed by many developing countries were primarily a response to the perceived power of conferences. A possible way forward is to strengthen the provision of the General Agreement on Trade in Services (GATS), dealing with anticompetitive business practices to ensure that collusive pricing does not erode the gains from liberalization.

2. Kang (2000) uses the policy indicators developed by McGuire, Schuele, and Smith (2000) to estimate the impact of restrictive maritime policies on bilateral shipping margins, defined as the ratio of cost-insurance-freight import values to free-on-board export values. This approach suffers from well-known data problems (import and export values are not reported by the same statistical entities) and by the undesirable property that shipping margins vary with unit values of shipped goods. The empirical approach adopted here addresses both of these problems.

3. In some respects, the approach to port services, which can be seen as essential facilities often controlled by major or monopoly suppliers, was analogous to the approach to basic telecommunications networks established in the procompetitive regulatory principles.

I. An Overview of International Maritime Transport

Maritime transport services consist of three types of activities: *international maritime transport* (freight and passengers), that is, the actual transportation service performed once the commodity is on board a ship in a country until the moment when the vessel reaches the destination port of a different state;[4] *maritime auxiliary services*, that is, any activities related to cargo manipulation in ports and on ships;[5] and *port services*, that is, activities related solely to ship management in ports.[6] This article uses data pertaining to restrictions affecting each segment of the market.

Due to differences in commodity types as well as to technological improvements in the shipping industry (most importantly, containerization), international maritime freight transport has developed specialized branches. Thus, *liner shipping*—meaning maritime transport of commodities by regular lines that publish in advance their calls in different harbors—is distinct from *tramp shipping*, referring to transport performed irregularly, depending on momentary demand. Typically, liner carriers transport commodities with a higher degree of industrial processing using containers, whereas noncontainerized raw materials (crude and refined oil, iron ore, grain, coal, and bauxite), generically known as bulk, tend to be carried in tramp carriers.[7]

Tramp shipping is generally believed to be a fairly competitive market, mostly free from restrictions (WTO 1998). By contrast, liner shipping has traditionally been subject both to private cartel-like arrangements and government restrictions. This article concentrates on the liner segment of the market.

Cargo Reservation Schemes

Over time, the most important category of barriers applied to international maritime transport has been various cargo reservation schemes. These schemes require that part of the cargo carried in trade with other states must be transported only by ships carrying the national flag or interpreted as national by other criteria. These policies have typically been justified by either security (self-sufficiency in times of war) or economic (infant industry) concerns.

4. International transport as defined by GATS excludes cabotage, which refers to transportation of commodities between ports of the same country.

5. In the GATS classification, maritime auxiliary services include maritime cargo handling, storage and warehousing, customs clearance, container station and depot, maritime agency, and maritime freight forwarding.

6. In the GATS classification, port services include pilotage, towing and tug assistance, provisioning, fueling and watering, garbage collecting and disposal, port captain's services, navigation aids, shore-based operational services, and emergency repair facilities.

7. Bulk traffic is typically divided into two categories: tanker (including crude oil and oil-related products) and dry bulk (including iron ore, grain, coal, bauxite, and phosphates). Note that the distinction between liner and bulk is not watertight. There exists a gray area that includes break-bulk (that is, loose, noncontainerized cargo transported using liners), general cargo (nonbulk commodities transported on liners without using containers), or containerized goods transported by tramp carriers.

Cargo reservation takes various forms. It can be imposed unilaterally, so that ships flying national flags are given the exclusive right to transport a specified share of the cargo passing through the country's ports. An alternative and more common form involves cargo sharing with trade-partner countries on the basis of bilateral or multilateral agreements. A specific form of such a scheme is the U.N. Conference on Trade and Development (UNCTAD) Liner Code of Conduct or the 40-40-20 rule. This legal instrument, which was adopted in 1974 and entered into force in 1983 by its ratification by more than 70 countries, was meant to counteract the anticompetitive actions of liner conferences, which are cartel-like arrangements. In many cases, access of outside shipping companies to a liner conference used to be restricted,[8] so governments applying the Liner Code required these cartels to divide the cargo transported according to the following rule: 40 percent for ships belonging to the exporting country, 40 percent for ships belonging to the importing country, and 20 percent for ships belonging to other countries. These restrictions were intended to encourage the development of the shipping industry of developing countries.

Cargo reservation schemes have declined in significance, as more countries have phased them out formally or chosen not to implement them. For example, in Asia, Indonesia adopted an "open sea" policy in the late 1980s, Thailand abolished all cargo reservations in 1993, and Korea's commitment to phase out its "designated cargo system," made in 1995 on becoming a member of the Organisation for Economic Co-operation and Development (OECD), was fully implemented by 1998. In Africa, Côte d'Ivoire and Senegal are among the countries that have eliminated cargo-sharing schemes. In Latin America, Chile pioneered liberalization in 1979 and Peru phased out most restrictions in the early 1990s. A further indication of the reduced importance of cargo sharing is the spread of "open registries" in many countries and the intensification of the "deflagging" process, that is, the transfer of ships to open registries to enable the ship owners to benefit from more efficient cost conditions (WTO 1998). The UNCTAD Liner Code, which was never applied on a large scale, is even less visible today, being applied mostly on routes between West Africa and Europe.[9]

Nevertheless, the evidence we have obtained on policy suggests that 11 of the countries in our sample, ranging from Benin to India, still have in place reservation policies that at least nominally restrict the scope for trade. Most of these countries subscribed to the UNCTAD Liner Code, whereas others (for example, Brazil, China, and Nigeria) implemented schemes that were similar in spirit.

8. The United States banned closed conferences. Cargo sharing and shipping conferences interacted over time, and in many cases authorities tailored their policies by taking into account the presence of carrier agreements. For example, Chile's cargo reservation mechanism, before the liberalization of the past decades, was designed so as to favor access of Chilean shipping companies into conferences and to restrain conference pressures on nonaffiliated carriers (Bennathan 1989).

9. In many countries, national shipping companies that had access to the reserved share, but did not possess sufficient technical means for its transportation, used to sell their preferential right to cargo, a practice that resulted eventually in a higher transport cost.

Nicaragua offers an example of a country that imposes a reciprocity condition, that is, access for foreign ships depends on whether their home countries grant Nicaraguan ships access.

There are few empirical studies of the impact of cargo reservation. The most frequently cited is Bennathan's (1989) analysis of the determinants of freight charges in Chile's export trade with the United States before and after elimination of cargo reservation. The results show that the indicators of cost-based pricing have greater explanatory power after liberalization than the indicators of demand-based pricing. This is seen as evidence of increased competition in the liner shipping industry. In another study, Pálsson (1997) suggests that in the South American market, the abolition of cargo sharing led to a decline in shipping rates to and from Europe by 20–50 percent, and to and from the United States by 25–35 percent. This study also indicates that cargo rates from Europe to Abidjan and Dakar declined by 10–20 percent a year after liberalization in 1995 (Pálsson 1997).[10] None of these studies provide any details on how these estimates were obtained, and to our knowledge there is not any cross-country analysis of the impact of cargo-reservation schemes.

Price Fixing and Other Cooperative Agreements

Maritime carriers enter various types of agreements, which help them enjoy advantages that arise from cooperation on technical or commercial matters. Far from being a recent phenomenon, carriers' collusive habits are deeply rooted in the history of maritime transportation, and the first shipping conferences, covering the routes between the United Kingdom and Calcutta, date back to 1875. By joining carrier agreements, shipping companies retain their juridical independence but consent to common practices with the other members regarding pricing, traffic distribution, and/or vessel capacity utilization. Examples of carrier agreements that were recognized in U.S. regulation by the end of 1998 were conference agreements, cooperative working agreements, joint services agreements, pooling agreements, space charter agreements, and trans-shipment agreements.

Conference agreements are made between two or more ocean common carriers, and provide for the fixing of and adherence to uniform tariff rates and conditions of service.[11] Conferences are the most widespread type of rate-binding agreement. In the United States, conferences are required by law not to restrict the entry and exit of any shipping company. Therefore, shipping conferences in U.S. foreign trade are "open," whereas those covering other routes may be closed

10. There are also some studies of the impact of the U.S. Jones Act, which prohibits foreign shipping firms from transporting goods or people from one U.S. location to another. Estimates of the price-increasing effect range from 100 percent (USITC 1991) of the average world price to a high of 300 percent (White 1988). Francois and others (1996) estimate that the welfare costs of this protection (assuming a conservative 100 percent price difference) are at least $3 billion a year.

11. Because conferences are a characteristic of liner shipping, they are also referred to as liner conferences.

to outside carriers.[12] The U.S. Shipping Act of 1984 defines cooperative working agreements as agreements that establish exclusive, preferential, or cooperative working relationships, but that do not fall precisely within the arrangements of any specifically defined agreement. Only some of the carrier agreements have a rate-binding clause, that is, they declare that they engage in unique price setting for transport services provided by all members.

The high incidence of conferences and other types of carrier agreements in maritime transport is due to the fact that the United States, the European Union, and many other countries exempt shipping conferences from antitrust regulation on the ground that they provide price stability and limit uncertainty regarding available tonnage.[13] The exemption from antitrust law is compounded by the Federal Maritime Commission's role in helping police price-fixing arrangements. The 1984 U.S. Shipping Act required all ocean carriers to file their rates with the Federal Maritime Commission and publish their rate and schedule information. Secret discounting on filed rates was considered illegal. Through the imposition of fines, the commission was authorized to ensure that the filed rates were actually charged.[14] However, conferences were required to allow for independent action, meaning that members could post a rate different from the conference rate, provided they notified the conference in advance. Although this provision created some flexibility, there was probably limited incentive to make public preannounced price cuts that were likely to be matched by rivals.

In recent years, the power of conferences has eroded for two reasons. The first is the entrance in the market of strong and efficient outside shipping companies. Containerization and other forms of technological progress have made it possible for outsiders to supply the same services as the conferences at lower costs to consumers. A second development is the change in regulations affecting international shipping, notably the U.S. Ocean Shipping Reform Act of 1998, which amended the Shipping Act of 1984 and went into effect in May 1999. While preserving the antitrust immunity of the rate-setting conference system, the Ocean Shipping Reform Act allows for the confidentiality of key terms (prices are included in this category) in contracts between shippers and carriers. This amendment is bound to create greater scope for price competition.

12. Recently, the European Commission claimed that steps taken by the Trans Atlantic Conference Agreement (TACA) to comply with the "open" conference obligations of U.S. law had constituted an abuse of their dominant position. It was alleged that TACA offered inducements to certain shipping lines to enter trans-Atlantic trade as parties to the conference rather than as independents (Levitt 2000).

13. Francois and Wooton (1999). See also Davies (1986), who states that "generally these cartels have been exempted from domestic legislation on competition, primarily because of jurisdictional problems stemming from the international character of the industry, but also because governments have judged them useful for promoting the health of both international trade and national merchant marines" (p. 300).

14. The rationale for these measures was ostensibly to protect small shippers from being disadvantaged by their inability to extract discounts from shipping companies.

In response to these developments, two types of arrangements have begun to emerge. First, shipping lines now sometimes enter discussion agreements. These allow conference and nonconference carriers serving a particular trade lane to discuss and share information about rates, costs, capacity, and service. The members may adopt voluntary rate, capacity, and service guidelines. Second, shipping companies and conferences tend to enter more wide-ranging organizations, such as consortia, alliances, and global alliances. There are two interesting questions, only the first of which we address in this article: How much do the traditional conferences still matter? Although these new arrangements are different from conferences from a juridical point of view, how different are they in actual behavior?

Some recent events provide implicit evidence of the continued influence of collusive practices. Although price fixing by conferences is exempted from the scope of competition law, the abuse of a conference's dominant position and the extension of collusion to other areas have provoked the wrath of European competition authorities. In 1992, the European Commission imposed fines against the members of the French/West African Ship Owners Conference.[15] The commission found that the conference had deterred the entry of other operators by a combination of loyalty agreements with shippers and predatory pricing against nonconference lines. Furthermore, competition between lines belonging to different conferences had been prevented by a partitioning of shipping routes: members of one conference were prohibited from operating in the ports served by another conference unless they first obtained membership, through a long, uncertain procedure, of that second conference.

In 1998, the European Commission fined the Trans Atlantic Conference Agreement (TACA) a sum of $314 million. The European Commission concluded that the conference, which controlled more than 60 percent of the traffic crossing the Atlantic at the time, set prices not only for the ocean leg but also for inland transportation by truck or train as well.[16] In May 2000, the European Commission imposed a penalty on 15 liner shipping companies that were members of the Far East Trade Tariff Charges and Surcharges Agreement (FETTCSA)—an agreement abandoned in 1994 following action by the European Commission. Altogether, the companies controlled 80 percent of the traffic between northern Europe and the Far East. Again, the target of action was not price fixing per se, but the FETTCSA members' collective strategy of not offering discounts from published fares.[17] Finally, reports in the maritime press also suggest that despite an

15. See p. 13 of the Annex of the Communication from the European Community and its Member States to the WTO Working Group on the Interaction between Trade and Competition Policy, WT/WGTCP/W/140, June 8, 2000.

16. As reported by CNN, the event marked a new record for fines imposed by the European Commission on a cartel. This was the first time that any EU authority had assessed the compatibility of liner conference practices with EU competition law.

17. See "FETTCSA: Commission fines shipping lines for an illegal price agreement on the Europe / Far East trade" (DN: IP/00/486), available from the European Commission's Web page.

increase in entry, the limited reductions in transport costs are attributable to the legal privileges granted to shipping company agreements.[18]

Notwithstanding this evidence, the issue of whether liner conferences were in a position to exercise market power has provoked some debate, primarily over the question of whether liner shipping markets are contestable. One view is that liner shipping markets satisfy a list of a priori conditions of contestability (Davies 1986, Zerby 1988). The entrant and the incumbent lines have access to the same technology and, provided that the market is not affected by any other distortions (such as cargo reservation), all shipping lines are equally placed with respect to access to cargo. Ship mobility and an active secondhand market imply that no significant sunk costs arise in the industry. Furthermore, the incumbent shipping lines are likely to provide a slow price response to the new entrants, especially if the former are organized in conferences, where price decisions require consensus among members. The frequent entry and exit on certain maritime routes has been cited as evidence of the contestability of these markets (Davies 1986).

An alternative view questions each of these assertions (Pearson 1987, Jankowski 1989). First, it argues that building up goodwill represents a substantial sunk cost, and lines cannot enter and exit markets with complete disregard of the effect this has on their reputation. Advertising and agency costs expended to establish regular liner services are also examples of intangible sunk costs.[19] Second, there is evidence that conferences have developed quick price-response mechanisms to respond to entry, including the use of action committees vested with the power to match the rates offered by an outsider. These arguments are in line with the findings of the European Commission cited previously. Finally, the frequency of entry and exit is clearly not convincing evidence of contestability: after all, at least in equilibrium, a contestable market would witness no entry at all.

As far as we know, there has been only one attempt to examine econometrically the influence of conferences. Clyde and Reitzes (1995) find no statistically significant relationship between freight rates and the market share of conferences serving a route. However, they find that the level of freight rates is significantly lower on routes where conference members are free to negotiate service contracts directly with shippers. On this basis, they conclude that the evidence on whether liner conferences are effective cartels is at best mixed. They suggest that there is an alternative interpretation of their results: the industry's antitrust immunity and tariff filing and enforcement requirements potentially facilitate the ability

18. See, for example, "Obstacles Lie Ahead," *1999 Year-end Economic Review*, *Bangkok Post*, 1999.

19. Franck and Bunel (1991) suggest that it may be appropriate to distinguish between two market segments in liner shipping by the type of entry criterion. In the first, entry is easy and competition from occasional outsiders is strong because they adopt a hit-and-run strategy and are not concerned with staying in business. In the second, an outsider is interested in competing on a long-term basis with the conferences and providing as high-quality services as them, so it is more difficult to comply with the entry conditions.

of all carriers to collude, not just those carriers that are conference members. Furthermore, in seeking to determine the influence of the market share of conferences, they consider only the routes on which conferences exist. They do not test a more basic hypothesis that the critical effect is of the existence of a conference on a particular route rather than its precise share of the market.

Restrictions on Port and Auxiliary Services

Both port and auxiliary services, particularly cargo handling, have tended to be monopolized. Liberalization of these services has two aspects. One is to ensure that foreign ships serving the domestic market obtain nondiscriminatory access to such services. The second is to allow competition, domestic and foreign, in the supply of the service itself.

Seaports are typically coordinated by public or, in fewer cases, private organizations called port authorities. Depending on the role assumed by these institutions, seaports can be classified into different categories. With *landlord ports*, the port authority owns and manages port infrastructure and private firms provide the rest of port and maritime auxiliary services; private firms are able to own superstructure and operate assets pertaining to infrastructure by concession or licensing. With *tool ports*, the port authority owns both infrastructure and superstructure, but private firms provide services by renting port assets through concessions or licenses (for example, Antwerp, Belgium). Finally, with *service ports*, the port authority owns assets and supplies services by directly hiring employees.

Trujillo and Nombela (1999) argue that the landlord port is the most desirable category from the efficiency point of view, since it allows private enterprise and market forces to play a role in the supply of services, while preventing monopolization of essential assets by private firms. For instance, in the case of Puerto Nuevo in Buenos Aires, six terminals were competitively commissioned to the private sector, with substantial foreign participation in the case of three.[20] The government also established free entry in the sector by allowing any operator to build, manage and operate a port for public or private use. These reforms transformed Argentine ports from the most expensive in Latin America to among the cheapest. Average charges per container declined from $450 to $120 and container time at port declined from 2.5 to 1.3 days (Trujillo and Estache 2001). Chile also witnessed a significant improvement in port performance after the competitive allocation of the right to operate ports, with five major world operators participating in the bidding consortia (Foxley and Mardones 2000).[21]

20. Terminals 1 and 2 were originally awarded to an international consortia headed by P&O Australia in partnership with Fasce SA, a local stevedore company (Trujillo and Estache 2001). Terminal 5 was awarded to an international consortium headed by the Manila-based international operator International Container Terminal Services, Inc.

21. Hutchinson, P&O, Stevedoring Services of America, HHLA and ICTSI. In fact, World Bank (2000) reports that the top nine international terminal operators account for 40 percent of the world's container liftings.

This anecdotal evidence indicates that international participation in the provision of terminal services is now a reality and that the introduction of competition can make a substantial difference in performance. With this broad benchmark in mind, we seek to capture some of the restrictions in place on port and auxiliary services.

II. The Model

In this section, we develop an econometric model of liner transport prices for U.S. imports. The analysis focuses on the ocean leg of the journey because the data available do not directly capture the price of maritime auxiliary services and port services.[22] Nevertheless, the analysis includes policy restrictions affecting the latter type of services. This is because the restrictions are likely to have an adverse effect on the efficiency with which these services are supplied to liners and hence push up the costs of liner services—for example, because of longer waiting or unloading times.[23]

We do not formally derive our estimation equation from a fully specified structural model of competition or collusion among liner companies, but our approach can be best understood in terms of a simple constant-elasticity pricing formula. This pricing rule relates the U.S. dollar price of shipping product k from foreign port i (which is located in country I) to U.S. port j (which is located in U.S. customs district J), P_{ijk}, to the marginal cost for this service, $MC(i, j, k)$, and a markup term, $\Phi(I, J, k)$:

$$P_{ijk} = \Phi(I, J, k)\ MC(i, j, k). \tag{1}$$

The markup term is a function of the elasticity of demand perceived by liner companies serving the routes between country I and customs district J for product k. The pricing formula in equation (1) could, for example, be easily derived from a model of Cournot competition.

Taking natural logs of equation (1) yields

$$p_{ijk} = \phi(I, J, k)\ mc(i, j, k), \tag{2}$$

where lowercase letters refer to natural logs of the respective variables.

Unfortunately, we do not have any direct information on the costs of maritime transport operations. We therefore decompose the *marginal cost term*, $mc(i, j, k)$, as follows:

$$mc_{ijk} = \alpha_J + \lambda_k + \gamma T_{ijk} + \delta d_{ij} + \eta q_{iJ} + \rho CR_I + \varphi^1 PS^1{}_I + \varphi^2 PS^2{}_I. \tag{3}$$

22. More precisely, the data reflect transport charges incurred in bringing the merchandise from alongside the carrier at the port of export and placing it alongside the carrier at the first U.S. port of entry.

23. The possibility of measurement error provides a more mundane reason for considering the impact of restrictions on the port and auxiliary services. Although in principle the liner transport prices do not include the prices of these services, in practice such a clean truncation may not have been possible.

The first term, α_J, reflects an effect specific to each U.S. customs district. It captures differences across customs districts in port services and other auxiliary services, such as cargo handling, and has been included for the reasons noted above. The second variable, λ_k, is a product-specific effect that captures differences in the physical properties of shipped goods, such as weight or size.

The third effect is a technological effect represented by the share of goods shipped in containers, T_{ijk}. Since containerization is likely to reduce the marginal cost of liner services, we expect the coefficient γ to have a negative sign.[24] The fourth cost variable is (the natural log of) the shipping distance between foreign port i and the main port in customs district J, d_{iJ}. There is some evidence that the effect of shipping distance on transport cost becomes less important for longer distances (Hummels 1999), and so we expect $0 < \delta < 1$. Fifth, we include an economies-of-scale effect represented by (the natural logarithm of) the total value of U.S. imports carried by liners (including nontextile goods) between foreign port i and district J, q_{iJ}. If there are economies of scale with regard to traffic originating from the same port, we expect the coefficient η to be negative.

Finally, we add three policy indicators that capture restrictions maintained by I's government affecting the supply of maritime services by foreigners. These restrictions are expected to lead to inefficiencies and the employment of outdated technology. Specifically, CR_I is a dummy that indicates whether exporting countries maintain any form of cargo reservation policy for the domestic shipping fleet affecting trade with the United States. $PS^1{}_I$ is an index that captures the existence of barriers to the foreign supply of cargo-handling services, considered to be one of the most important auxiliary services. $PS^2{}_I$ is an index that measures the number of port services (for example, pilotage, towing, navigation aids, and waste disposal) that are mandatory for incoming ships. In the absence of more direct data on the openness of the port services regime, the extent to which the use of such services is mandatory is used as an indicator of the restrictiveness of the port services regime. As noted above, the costs of auxiliary and port services are not directly captured by the maritime price data, but restrictions in both are relevant because they could push up the costs of liner services.

The *markup term*, $\phi(I, J, k)$, is assumed to depend on the following four variables:

$$\phi(I, J, k) = \mu_k + \tau CR_I + \psi^1 A^1{}_{IJ} + \psi^2 A^2{}_{IJ}. \tag{4}$$

The first term, μ_k, reflects a product-specific effect that captures differences in transport demand elasticities across sectors. Note that the transport demand elasticities are derived from the final demand for product k in the United States. The second variable is again the variable that captures the existence of cargo reservation policies, which directly limit the extent of competition from foreign liners and thus may push up markups. The third and fourth effects, $A^1{}_{IJ}$ and $A^2{}_{IJ}$,

24. Over 80 percent of U.S. imports are containerized. Noncontainerized shipments are because certain foreign ports, mostly in the developing world, are not yet equipped with container terminals.

are due to the existence of collusive agreements among liner companies on routes between country I and customs district J. We distinguish between two kinds of collusive agreements: price-fixing agreements (which include most conferences) and cooperative working agreements that do not have a binding rate-setting authority. A single agreement typically covers routes between the ports of a foreign country and one or more U.S. coastal districts that each consists of several customs districts. Because collusion between liner companies is likely to push up markups, we expect both coefficients ψ^1 and ψ^2 to show a positive sign. But conference and other price-fixing agreements are likely to be more powerful and to have a greater impact on transport prices than cooperative working agreements, that is, we expect $\psi^1 > \psi^2$.

Substituting equations (4) and (3) into equation (2) and inserting an error term, ε_{ijk}, we obtain

$$(5) \qquad p_{ijk} = \alpha_J + \beta_k + \gamma T_{ijk} + \delta d_{IJ} + \eta q_{IJ} + \psi^1 A^1{}_{IJ} + \psi^2 A^2{}_{IJ} + \omega\, CR_I + \varphi^1\, PS_I^1 + \varphi^2\, PS_I^2 + \varepsilon_{ijk},$$

where $\beta_k \equiv (\lambda_k + \mu_k)$, $\omega \equiv (\rho + \tau)$, and we expect the coefficients on the three policy indicators, ω, φ^1, and φ^2, to have a positive sign.[25]

We calculate the transport price, p_{ijk}, as the share of liner transport charges in import values for good k (at the six-digit HS aggregation) multiplied by the unit value of imports. The U.S. Department of Transportation defines transport charges as all freight, insurance, and other charges (excluding import duties) incurred in bringing the merchandise from alongside the carrier at the port of export and placing it alongside the carrier at the first U.S. port of entry.[26] However, actual data reported may include charges for port services and inland transportation.[27] To reduce the potential bias resulting from differences in inland transportation costs, we exclude observations for which the origin of the import is different from the country of the port of shipment (for example, landlocked countries) as well as all in-transit shipments.[28] The appendix provides additional information on the construction and sources of all variables.

Table 1 presents an overview of our estimation data set. It covers all U.S. imports carried by liners from the 59 countries for which we could find information on maritime policies. Data refer to 1998. Liner imports account for around 65 percent of the total value of maritime imports, the remaining 35 per-

25. Because we estimate both product fixed effects and customs district–specific effects, we need to drop one dummy variable (for one customs district) to avoid perfect colinearity among the explanatory variables.

26. If insurance costs are not closely correlated with transport charges, there is the possibility that our transport price variable is distorted. However, this should at least partially be remedied by the inclusion of product fixed effects, because differences in insurance costs are likely to be greatest across products.

27. According to e-mail communication with an official from the U.S. Department of Transportation.

28. Note that we do not exclude the trade originating in third countries and in-transit traffic when calculating total import values q_{IJ}.

TABLE 1. Overview of U.S. Imports Carried by Liners in 1998

Countries	Number of countries	Liner transport charges (million $)	Liner import charges (million $)	Share of liner maritime imports (%)		Share of liner imports in total imports (%)	
				Total	Non-oil[a]	Total	Non-oil[a]
Developing	37	3,940	82,400	64.88	74.82	48.76	53.10
Industrial	22	3,080	104,500	64.20	66.24	50.73	51.23
Total	59	7,020	186,900	64.71	70.04	50.13	52.38

[a]Excluding HS category 27.
Source: U.S. Department of Transportation and U.S. Bureau of Census.

cent being carried by tramp services.[29] About half of all U.S. imports (including all modes of transport—maritime, air and road) from the 59 countries considered are carried by liners.

III. THE ESTIMATES

We begin with ordinary least squares estimation of equation (5) over the entire dataset. The error term ε_{ijk} is assumed to be independently distributed across exporting countries, but we allow for interdependence among observations within each country.[30] The results are presented in table 2.

Although the coefficients mostly accord with our expectations, this empirical approach has a weakness: it ignores competition from alternative modes of transportation, expressly tramp maritime services (bulk and tanker), air transport, and road transport (in the case of Canada and Mexico). For a number of product categories, it is likely that shippers face an explicit tradeoff between the quality and cost of shipping a good by these alternative modes of transport. One approach to remedying this problem is to exclude all products for which competition from tramp maritime and air services is important. Since it is difficult to make a clean separation based on product characteristics alone, we adopt a method relying on the revealed importance of the alternative modes. Specifically, we exclude all observations where either the share of air transport as a percentage of total imports for shipping product k from country I to customs district J is positive, or the share of tramp services for a particular product k on all routes between country I and district J exceeds 15 percent.

29. However, if we exclude U.S. oil imports (HS category 27), this share rises to 70 percent and liner transport becomes relatively more important for developing countries.

30. Instead of using a fixed-effect specification as in equation (5), we also estimated a model with random product effects and maintaining the customs district fixed effects. This model yielded very similar estimation results. Moreover, the Hausman test rejected the null hypothesis that the individual effects are uncorrelated with our regressors in the model, supporting the use of fixed instead of random product effects.

TABLE 2. Full-Sample Fixed-Effects Model

Variable	Estimate	
Distance	0.298**	(4.97)
Containerization	–0.071**	(–2.80)
Total liner imports	–0.017*	(–2.07)
Price-fixing agreements	0.488**	(5.52)
Cooperative agreements	0.050	(1.21)
Cargo reservation	–0.067	(–0.77)
Cargo-handling services	–0.203*	(–2.44)
Mandatory port services	0.357**	(2.52)
Number of products	4,356	
Number of observations	250,237	
F-statistic	65.11**	
Adjusted R^2	0.775	

*Significant at the 5 percent level.
**Significant at the 1 percent level.
Note: The dependent variable (liner transport prices), distance, and total liner imports are expressed in natural logs; all other variables are expressed in actual levels; fixed effects are product-specific and U.S. customs district–specific (see text). The regression assumes an independently distributed error term across exporting countries, but allows for interdependence among observations within each country. *t*-statistics are in parentheses. The *F*-statistic tests the joint significance of all independent variables (except the fixed effects).
Source: Authors' calculations.

This reduces our sample size from 250,237 to 98,997 observations. The estimation results with the reduced sample are presented in the first column of table 3.

Although these results are in line with our expectations, it is possible that the exclusion of observations introduces a sample bias in our estimation. We therefore adopt a sample selection model, where we estimate the likelihood of a shipment having no competition from air and tramp services (as defined above) in two separate probit equations. The explanatory variables in these probit equations are (the natural logs of) the unit value and the unit weight of shipments and, in the case of air transport, a dummy variable that captures the existence of an open-skies agreement between country *I* and the United States.[31] We estimate

31. Because the unit weight is unavailable for selected shipments, the number of observations in the probit regression is somewhat smaller than in the full sample. In the case of tramp services, we also included country fixed effects, except for Benin, for which the share of tramp services was below 15 percent for all observations. As in the liner pricing regression, we assumed that the error term in each probit equation is independently distributed across exporting countries, but allowed for interdependence among observations within each country.

this model using the Heckman two-step estimation procedure, assuming that the error terms in the two probit regressions are uncorrelated.[32]

The results of the sample selection model are presented in the second to fourth columns of table 3. In the "air" probit equation, the estimated coefficient on unit value is significantly negative and the coefficient on unit weight is significantly positive, suggesting that valuable and light products are more likely to be sent by air. By contrast, in the "tramp" probit equation, the coefficient on unit value is significantly positive and the coefficient on unit weight is significantly negative, indicating that tramp services are primarily used for heavy commodities with low unit values.[33]

In the final regression, we exclude Mexican and Canadian imports from our (already reduced) sample. For these two countries, road transport is an alternative mode of transport that may compete with maritime and air services. Table 4 presents the estimation results with both the simple reduced sample and the sample selection approach, which are similar to those presented in table 3.

Estimates of the Model Coefficients

The results from the different estimating methods reveal a reassuring consistency. The estimated coefficient on distance lies between 0.2 and 0.3 and is always significantly different from both zero and one. This confirms that transport cost increases with distance but less than proportionately. As we expected, containerization, as measured by T_{ijk}, works to reduce liner prices, the estimated coefficient being statistically significant. The coefficient on the total value of U.S. imports carried by liners, q_{ij}, takes a small and significant negative value. This suggests that there are economies of scale with regard to traffic originating from the same port, and that small countries or economies with small trading volumes may be relatively disadvantaged.

Consider now the impact of restrictions on trade in maritime services. The most striking finding is the strong positive impact on liner prices of the existence of rate-binding conference and other price-fixing agreements. The existence of cooperative working agreements has a weaker impact that is not always statistically significant. These results confirm our expectation that price-fixing agreements matter and are more important than cooperative working agreements.[34]

32. See Maddala (1983, p. 282) for a description of this model. The assumption that the error terms in the two probit regressions are uncorrelated seems reasonable: decisions on whether to ship goods by air or by vessel are likely to be independent of decisions on the mode of maritime transport.

33. Interestingly, the explanatory power is higher in the air probit regression than the tramp probit regression.

34. It is possible, in principle, that the formation of collusive carrier agreements is an endogenous variable—that collusion is more likely on more profitable routes. To account for this possibility, we estimated a treatment-effects model that corrects for the possible selectivity bias of the dummy variable on price-fixing carrier agreements (see Greene 1997, pp. 981–82, and Maddala 1983, p. 6). Similar to the sample selection model, we used the Heckman two-step estimation procedure that first estimates a probit model of the selection process and then the regression model with an additional selectivity correction variable. Our explanatory variables in the probit model were exporter GDP, unit weight, unit

The evidence on policy restrictions is mixed. The coefficient of the variable capturing the existence of cargo reservation policies is close to zero and not statistically significant in any of the regressions. This result gives credence to the claim that cargo reservation policies no longer exert an important influence on liner trade. The estimated coefficient on the restrictiveness index of cargo-handling services is the only one that has a counterintuitive sign and is statistically significant in the first set of estimates (table 2), but with other, arguably more reliable methods (tables 3 and 4), it ceases to be significant. Recall that our dependent variable captures the cost of complementary services not explicitly but only to the extent that they feed through into the ocean-leg liner prices. In this respect, the index on the restrictiveness of port policy probably has a stronger claim to significance. Our estimates would seem to confirm this—the coefficient is consistently positive and statistically significant. This result also seems in line with current wisdom that the biggest policy hurdles to competitive provision of shipping services are to be found at the ports rather than in the ocean leg.[35] However, it must be kept in mind that we are only using an indirect measure of port policy restrictiveness.

Estimates of the Consequences of Policy Changes

The estimated model can be used to calculate hypothetical reductions in transport prices due to both the breakup of private carrier agreements and allowing greater competition in the provision of port services. For this purpose, we take the estimated coefficients from the sample selection model in table 3, which we consider to be the most reliable estimates both from an economic and econometric standpoint. Table 5 presents the simulated price reductions. The breakup of conference and other price-setting agreements would lead to a more dramatic reduction in transport prices (32 percent) than the breakup of cooperative working agreements (18 percent), whereas the liberalization of port services would cause a 35 percent drop in the price of liner services.[36]

If we compute the trade-weighted percentage reductions in transport prices across all observations included in the sample selection model, the average total

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value, and total liner traffic between the exporting country and the importing coast district. Only exporter gdp made a positive and significant contribution to the likelihood of observing price-fixing agreements. The selectivity correction parameter, however, had a negative sign in the main regression and, accordingly, inflated the coefficient on the price-fixing dummy variable. This result would suggest that rate-binding carrier agreements typically occur on routes with lower prices. This counterintuitive finding may be due to the inadequacies of our explanatory variables in the probit equation. Alternatively, the formation of liner agreements may be less an outcome of market forces and more the result of historical and institutional forces, which are exogenously determined.

35. Of course, savings from the liberalization of port services are likely to be greater when their full impact on aggregate maritime transport costs is taken into account.

36. The policy simulation makes the simplifying assumption that the mandatory use of certain port services, such as pilotage, cannot be justified by safety or related concerns. This assumption does not seem unreasonable, given that safety standards can also be enforced in more liberal policy environments.

TABLE 3. Reduced-Sample and Sample Selection Models

Variable	Reduced-sample model	Sample selection model: Air probit	Sample selection model: Tramp probit	Sample selection model: Liner transport prices
Distance	0.202** (4.58)			0.228** (5.27)
Containerization	−0.132** (−3.89)			−0.116** (−3.00)
Total liner imports	−0.018* (−2.39)			−0.025** (−3.40)
Price-fixing agreements	0.443** (5.78)			0.379** (5.01)
Cooperative agreements	0.132* (2.64)			0.202* (2.53)
Cargo reservation	−1.106 (−1.13)			−0.099 (−1.01)
Cargo-handling services	−0.104 (−1.14)			−0.064 (−0.56)
Mandatory port services	0.307** (2.26)			0.437** (2.83)
Unit value		−0.385** (−18.44)	0.130** (13.12)	
Unit weight		0.448** (17.49)	−0.131** (−11.52)	
Open skies agreement		0.007 (0.07)		
Sample selection correction (air)				0.399** (3.19)
Sample selection correction (tramp)				−0.733* (−2.37)
Number of products	4,214			4,208
Number of observations	98,997	250,159	250,159	98,815
F-statistic	39.43**			42.51**
Adjusted R^2	0.779			0.783
Pseudo R^2		0.130	0.054	

*Significant at the 5 percent level.
**Significant at the 1 percent level.

Note: The dependent variable (liner transport prices), distance, total liner imports, unit value, and unit weight are expressed in natural logs; all other variables are expressed in actual levels; fixed effects are product-specific and U.S. customs district-specific (see text). All regressions assume an independently distributed error term across exporting countries but allow for interdependence among observations within each country. The sample selection correction variables are computed following Heckman's two-step estimation procedure. *t*-statistics (for liner price regressions) and *z*-statistics (for probit regressions) are in parentheses. The *F*-statistic tests the joint significance of all independent variables (except the fixed effects).

Source: Authors' calculations.

TABLE 4. Reduced-Sample and Sample Selection Models without Mexico and Canada

Variable	Reduced-sample model	Sample selection model: Air probit	Sample selection model: Tramp probit	Sample selection model: Liner transport prices
Distance	0.221** (4.12)			0.215** (4.79)
Containerization	–0.147** (–4.49)			–0.142** (–4.13)
Total liner imports	–0.017* (–2.19)			–0.024** (–3.30)
Price-fixing agreements	0.464** (5.65)			0.372** (4.60)
Cooperative agreements	0.124* (2.58)			0.192* (2.46)
Cargo reservation	–0.092 (–0.95)			–0.089 (–0.88)
Cargo-handling services	–0.107 (–1.17)			–0.068 (–0.59)
Mandatory port services	0.298** (2.24)			0.410** (2.69)
Unit value		–0.402** (–27.28)	0.131** (12.84)	
Unit weight		0.469** (27.56)	–0.132** (–11.30)	
Open skies agreement		0.021 (0.22)		
Sample selection correction (air)				0.465** (3.94)
Sample selection correction (tramp)				–0.694* (–2.26)
Number of products	4,190			4,184
Number of observations	97,676	247,673	247,673	97,518
F-statistic	36.74**			50.83**
Adjusted R^2	0.781			0.784
Pseudo R^2		0.136	0.054	

*Significant at the 5 percent level.

**Significant at the 1 percent level.

Note: The dependent variable (liner transport prices), distance, total liner imports, unit value, and unit weight are expressed in natural logs; all other variables are expressed in actual levels; fixed effects are product-specific and U.S. customs district–specific (see text). All regressions assume an independently distributed error term across exporting countries but allow for interdependence among observations within each country. The sample selection correction variables are computed following Heckman's two-step estimation procedure. t-statistics (for liner price regressions) and z-statistics (for probit regressions) are in parentheses. The F-statistic tests the joint significance of all independent variables (except the fixed effects).

Source: Authors' calculations.

TABLE 5. Simulated Reductions in Transport Prices

Simulation	Breakup of cooperative working agreements	Breakup of price-fixing agreements	**Cumulative effect of the breakup of private carrier agreements**	Liberalization of port services	*Cumulative total effect*
1. Percentage reductions on restricted routes	18.30	31.56	**44.09**	35.43	*63.90*
2. Trade-weighted percentage reductions across all observations in our dataset	7.11	18.71	**24.25**	8.66	*30.76*
3. Total savings across all observations in our dataset:					
Absolute value(in million $)	140	371	**484**	201	*637*
As a percent of total transport charges[a]	5.59	14.80	**19.42**	7.99	*25.59*
4. Projected total savings across all exporting countries and all sectors (in million $)[b]	575.1	1,522.6	**1,997.9**	822.0	*2,632.7*

Note: These calculations are based on the estimated coefficients of the sample selection model in table 3. Given the functional form of the regression equation, the individual effects do not sum to the total effect.

[a]The share of total savings in total transport charges is equivalent to the unweighted average percentage reductions in transport prices.

[b]The projected total savings in the last row apply the percentage savings in total transport charges estimated for the reduced sample to total liner transport charges for all U.S. imports.

Source: Authors' calculations.

reduction would be 30.8 percent—made up of the cumulative effects of the breakup of carrier agreements (24 percent) and the liberalization of port services (9 percent). Total savings would sum to $637 million of transport charges. To get a sense of the overall magnitudes involved, we can project these savings to total U.S. imports carried by liners across all sectors and all routes. Our simulations reveal that the removal of public restrictions to liner trade would lead to savings of up to $822 million and the breakup of private cartels would bring about additional savings of up to $2 billion.

There are two important qualifications to these estimates. First, the pattern of restrictions in our limited sample may not be representative of the pattern of restrictions in trade of all products across all routes. Second, competition from other modes of transport for some products may limit the ability of carrier agreements to fix prices. But note that our simulation pertains to the savings arising from goods carried to the United States alone. The imports of the United States are only about a fifth of total world merchandise imports. So global gains from the elimination of all forms of restriction are likely to be substantially larger, particularly if we take into account the indirect benefits from reducing impediments to trade.

IV. Conclusion

Our estimates confirm the general belief that cargo reservation policies, which proliferated in the 1970s and 1980s, are no longer an important barrier to trade. However, it emerged that both public policy, specifically in the form of restrictions on the provision of port services, and private practices continue to exercise a significant influence on maritime transport prices. Interestingly, private anticompetitive practices have a stronger influence on prices than public restrictions do.

These results challenge the notion that collusive carrier arrangements have lost their significance over the past decade. In defense, maritime industry sources frequently point to the fact that liner operators hardly break even and, on this basis, argue that there is little scope for price reductions. But it is well known that protection and cartel-like behavior in the presence of fixed costs can lead to inefficient entry and reduced profitability. The benefits of competition typically arise not only from increased allocative efficiency—that is, pricing close to costs—but also from increased internal efficiency—that is, a reduction in costs. There may be scope for increasing this latter type of efficiency in the maritime industry.

Our results need to be qualified. First, we focused only on routes leading to the United States. Although there is need for further research on other routes, the paucity of transport data in other countries is a major constraint. Second, the analysis herein has focused solely on the maritime leg of the transport journey and has not examined distortions on the inland section. Evidence suggests that the ocean leg accounts for a little more than a third of total door-to-door shipping charges (OECD 1968, Livingston 1986). Unfortunately, there are no comprehensive data on such charges. An ambitious future research program would seek to disaggregate the components of door-to-door shipping charges and subject them to an analysis similar to that carried out in here. A critical component of such a program would be to develop better measures of the restrictiveness of port and auxiliary services than have been used here.

Notwithstanding these qualifications, this article has certain implications for policy. The elimination of policy restrictions to trade in maritime transport services is likely to produce substantial gains. Many of these restrictions can be removed unilaterally, and the GATS can be used to bind the openness to reduce uncertainty and the possibility of policy reversals. But it is not enough to eliminate policy restrictions. There is also a need to deal with the private anticompetitive practices of international maritime cartels. Large states can probably tackle such practices unilaterally through their own competition laws, despite the extraterritoriality problems involved. But small states with limited enforcement capacity are at a disadvantage; the problem is accentuated by the fact that major trading countries have diluted the application of competition disciplines to the maritime sector. One positive development described earlier is the elimination by the United States of some of the provisions in its shipping law that helped police price-fixing arrangements. Whether collusion can be sustained in the absence of such facilitating devices is open to question. But we would argue that there is cause for concern as long as the basic rate-setting conference system continues to enjoy antitrust immunity.

An international initiative would seem desirable. One approach would be to deal with the problem by creating sector-specific competition rules, as in the case of basic telecommunications. Or, if such anticompetitive practices also affect other services sectors, there may be a need to strengthen the general GATS disciplines. Currently, Article IX of the GATS (which deals with private anticompetitive practices) has little substance, providing only for an exchange of information and consultations. The current round of services negotiations offers an opportunity to strengthen this provision.

What form could such a strengthening take? We believe that the harmonization of either sector-specific or general competition rules is probably neither feasible nor necessary. Our proposal is much simpler and would involve the creation of two obligations. The first would end the exemption of collusive agreements in the maritime sector from national competition law. The second would create the right of foreign consumers to challenge anticompetitive practices by shipping lines in the national courts of countries whose citizens own or control these shipping lines. The second obligation is necessary to deal with a possible failure to enforce and already has a precedent in the WTO rules on intellectual property and government procurement.[37]

Would it be feasible to create such rules? History does not provide cause for optimism. The procompetitive rules in basic telecommunications, in line with

37. See Mattoo and Subramanian (1997) for an elaboration of this argument.

most WTO rules, were designed to protect the market access rights of foreign suppliers, and conventional political-economy forces supported their creation. To establish rules that enable small countries to protect their consumers from foreign oligopolies will be far more difficult. In fact, the negotiating history of the GATS reveals successful opposition to the strengthening of Article IX from some of the countries that exempt maritime conferences from the scope of their antitrust laws. However, the reluctance of many developing countries to make liberalization commitments under the GATS did not strengthen their case. One strategy in the current round of services negotiations would be for a coalition of developing countries to put forward an offer of substantial liberalization conditional on the strengthening of Article IX. By targeting the twin maladies of maritime trade, such a strategy, if successful, would provide substantial global benefits.

Appendix: Data

Data on liner transport charges, import values, the percentage of containerized cargo, total imports carried by liners and the market share of tramp services are from the Waterborne Trade Database compiled by the U.S. Department of Transportation. The containerization variable is measured in terms of the weight of goods shipped. Tramp services are defined as bulk and tanker services. Unit values, unit weights, and the market share of air services are computed from the U.S. Merchandise Imports Database published by the U.S. Department of Commerce. This source does not publish data separately by foreign and U.S. ports; we therefore have to use these variables at the more aggregate level, that is, U.S. trading partners and U.S. customs districts.

Shipping distances were kindly provided from a private service called BP Marine. Some missing ports that are included in the Waterborne Transport Database had to be approximated by the closest neighboring port. Information on private carrier agreements between U.S. coastal districts and individual countries comes from the Federal Maritime Commission (1998). We excluded agreements signed before 1970 and also those with an unspecified regional coverage (for example, the Far East), because the *de facto* coverage of such agreements may only relate to a few particular routes. The potential bias introduced by this exercise is likely to be small because most routes covered by such regional agreements are also covered by country-specific agreements. As mentioned in the text, we construct two dummy variables to account for the presence of carrier agreements on maritime routes. The first refers to conferences and other price-fixing agreements and the second captures cooperative working agreements that do not have a binding rate authority. Data on the existence of open-skies agreements were taken from the Web site of the U.S. Department of Transportation.

The three indicators of trade restrictions are constructed based on information compiled from the following sources: WTO (1994), various WTO Trade Policy Reviews, GATS schedules of commitments (available online at http://gats-info.eu.int/index.html), APEC Individual Action Plan submissions (available online at http://www.apecsec.org), unpublished OECD documents, ECLAC (1999), EU Market Access Database (available online at http://mkaccdb.eu.int), and various editions of the National Trade Estimate Report on Foreign Trade Barriers compiled by the U.S. Trade Representative (available online at http://www.ustr.gov). In some cases, Greg McGuire kindly supplied data from the above sources.

The *cargo reservation* dummy variable is assigned a value of 1 if a country has a bilateral agreement involving cargo sharing with the United States, if it is a signatory of the U.N. Code of Conduct for Liner Conferences and applies Article 2 of the code in its trade with the United States, or if it sustains any kind of unilateral cargo reservation scheme; and 0 otherwise. The *cargo-handling services* index measures restrictions or special requirements imposed in a country to potential foreign suppliers of cargo-handling services (foreign suppliers means, in this case, locally registered companies with foreign participation in their capital or branches of firms established in other countries). The index values are 0 if there is no restriction, 0.25 if minor restrictions exist, 0.5 if a joint venture condition is imposed, 0.75 if a very high national participation in the capital of the company is required, and 1 if foreign companies are not allowed to provide cargo-handling services at all. In selected countries, consultations with industry experts suggested a slightly different assessment of policy restrictiveness in cargo handling than the one implied by the sources listed above. But a modification of the rankings did not lead to substantial changes in our empirical findings. The index on *mandatory port services* assigns a score of 0.125 for the existence of each of the following mandatory services: pilotage, towing, tug assistance, navigation aids, berthing, waste disposal, anchorage, and other mandatory services.

Table A-1 lists the countries for which we could find information on the three policy indicators and that are included in our estimation set. The table also shows the assigned values of these policy variables as well as the average value of the dummy capturing the two types of collusive carrier agreements (the latter lying between 0 and 1, if not all U.S. coastal districts are covered by the agreements affecting a particular country).

TABLE A-1. Indicators of Maritime Policy and Carrier Agreements, 59 Countries

Country	Cargo reservation	Cargo-handling services	Mandatory port services	Price-fixing carrier agreements	Cooperative working agreements
Argentina	0	0	0.13	0.00	1.00
Australia	0	0	0.13	1.00	1.00
Belgium	0	0	0.06	1.00	0.00
Benin	1	1	0.00	0.00	0.00
Brazil	1	0.5	0.75	0.00	1.00
Brunei	0	0	0.00	0.00	0.00
Canada	0	0	0.13	0.00	0.00
Chile	0	0	0.25	0.43	1.00
China	1	0.5	0.00	0.00	0.00
Colombia	0	0.5	0.13	0.50	1.00
Costa Rica	0	0	0.00	0.00	1.00
Côte d'Ivoire	0	0	0.25	0.00	1.00
Cyprus	0	1	0.31	0.00	0.00
Denmark	0	0	0.06	1.00	0.00
Dominican Republic	0	0.25	0.25	0.50	1.00
Ecuador	0	0	0.00	0.43	1.00
Egypt	1	0.75	0.75	0.00	0.00
El Salvador	0	0	0.00	0.00	1.00
Finland	0	0	0.25	0.00	0.00
France	0	0	0.38	1.00	0.00
Germany	0	0	0.38	1.00	0.00
Ghana	1	1	0.50	0.00	1.00
Greece	0	1	0.19	0.00	0.00
Hong Kong	0	0	0.25	0.00	0.00
Iceland	0	0	0.13	0.00	0.00
India	1	0	0.00	0.00	1.00
Indonesia	0	1	0.06	0.00	0.38
Ireland	0	0	0.13	1.00	0.00
Italy	0	0.25	0.50	0.38	0.00
Jamaica	0	0.5	0.00	0.00	0.60
Japan	0	0.75	0.13	0.89	1.00
Korea, Rep. of	0	0	0.38	0.00	0.00
Malaysia	0	0	0.25	0.00	0.38
Mauritius	0	1	0.38	0.00	0.00
Mexico	0	0.5	0.38	0.00	1.00
Morocco	1	0.5	0.13	0.00	0.00
Netherlands	0	0	0.50	1.00	0.00
New Zealand	0	0	0.38	1.00	1.00
Nicaragua	1	0	0.00	0.00	1.00
Nigeria	1	0	0.50	0.00	1.00
Papua New Guinea	0	0.5	0.00	0.00	0.00
Peru	0	0.5	0.00	0.50	1.00
Philippines	0	0.5	0.00	0.00	0.38
Poland	0	0.25	0.00	0.00	0.00
Portugal	0	0	0.13	1.00	0.00
Romania	0	0	0.63	0.00	0.00
Senegal	0	0	0.00	0.00	1.00
Singapore	0	1	0.38	0.00	0.33

(continued)

TABLE A-1. (continued)

Country	Cargo reservation	Cargo-handling services	Mandatory port services	Price-fixing carrier agreements	Cooperative working agreements
Spain	0	0	0.06	1.00	0.00
Sweden	0	0	0.06	1.00	0.00
Taiwan	0	0.5	0.00	0.00	0.00
Thailand	0	0.5	0.63	0.00	0.38
Togo	1	0	0.00	0.00	0.00
Tunisia	0	0.5	0.13	0.00	0.00
Turkey	0	0	0.00	0.43	0.00
United Kingdom	0	0	0.31	1.00	0.00
Uruguay	0	0	0.00	0.00	1.00
Venezuela	1	0	0.00	1.00	1.00
Vietnam	0	0	0.00	0.00	0.50

Note: The indicators on "price-fixing carrier agreements" and "cooperative working agreements" show the average value of the 0-1 dummy variable used in the estimation. This value lies between 0 and 1 if not all U.S. coastal districts are covered by the agreements affecting a particular country.

Source: Authors' calculations.

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Bank Risk and Deposit Insurance

Luc Laeven

Arguing that a relatively high cost of deposit insurance indicates that a bank takes excessive risks, this article estimates the cost of deposit insurance for a large sample of banks in 14 economies to assess the relationship between the risk-taking behavior of banks and their corporate governance structure. The results suggest that banks with concentrated ownership tend to take the greatest risks, and those with dispersed ownership engage in a relatively low level of risk taking. Moreover, as a proxy for bank risk, the cost of deposit insurance has some power in predicting bank distress.

Banking crises have shown not only that banks often take excessive risks but that risk taking differs across banks. Some banks engage in more risks than their capital could bear if the downside potential of the risks fully materialized; others are more prudent and would be able to weather a banking crisis. Whether different types of banks take different risks is not well known.

To see whether there is a relationship between risk-taking behavior and bank characteristics such as ownership structure, I analyze a large sample of banks in different economies. I measure the degree of a bank's risk taking by the value of deposit insurance services implicitly extended to the bank by the safety net to guarantee its deposits. This implicit deposit insurance cost is calculated by applying a well-known technique that models deposit insurance as a put option on the bank's assets.

The results provide empirical support for this method of assessing the risks of a bank. Implicit deposit insurance premiums are higher for banks in crisis countries and have some power in predicting bank distress.

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I. LITERATURE

Many countries have implemented deposit insurance schemes to prevent bank runs and to provide liquidity to banks in case bank runs do occur. In most countries that have explicit deposit insurance the schemes insure deposits only up to a certain limit, offering limited-coverage deposit insurance. In some countries, such as Turkey, the schemes insure deposits in full, providing a blanket guarantee. The advantage of a deposit insurance system that provides a blanket guarantee is that it fully eliminates bank runs. The disadvantage is that it destroys all potentially beneficial information production and monitoring by depositors. Bhattacharya, Boot, and Thakor (1998) show that partial deposit insurance encourages market discipline, exercised through bank monitoring by informed depositors, and that regulatory measures, such as limited regulatory forbearance and tough bank closure rules, may control bank risk taking. Underlying the partial insurance conclusion is the presumption that informed depositors—with their own assets at risk—will do a better job of monitoring banks than government regulators will.

Demirgüç-Kunt and Huizinga (1999) find empirical evidence that adopting an explicit deposit insurance scheme involves a tradeoff between increased depositor safety and reduced market discipline by bank creditors. Demirgüç-Kunt and Detragiache (1999) provide empirical evidence based on a large sample of countries that explicit deposit insurance increases banking system vulnerability in countries with weak institutional environments. More generally, Kane (2000) argues that the design of financial safety nets should take country factors into account, particularly the informational environment and the enforceability of private contracts.

Since Merton (1977), deposit insurance has typically been modeled in the literature as a put option on the bank's assets. Marcus and Shaked (1984) were the first to implement Merton's (1977) model and empirically test over- and underpricing of insurance premiums. Ronn and Verma (1986) claim that Marcus and Shaked (1984) incorrectly look at the preinsurance value of bank assets. They designed a model that looks at the postinsurance value of bank assets and incorporates capital forbearance by the bank regulators. Duan (1994, 2000) develops a maximum likelihood framework to estimate the value of deposit insurance. Duan's method is free of some of the statistical problems of Ronn and Verma's (1986) method, an issue discussed in more detail later.

Many empirical studies have applied these methods. Few of them look at developing economies, however. Duan and Yu (1994) calculate deposit insurance premiums for 10 exchange-listed depository institutions in Taiwan (China) in 1985–92. Using Duan's (1994, 2000) maximum likelihood estimation method to assess implicit deposit insurance premiums, they find that the deposit insurance agency subsidized these institutions in all years except 1989. They also find that the methods of Ronn and Verma (1986) and Duan (1994, 2000) produce significantly different estimates of the cost of deposit insurance. Fries, Mason, and Perraudin (1993) apply Ronn and Verma's (1986) method to 16 Japanese banks in 1975–92 and similarly find that the institutions were subsidized by the deposit insurance agency. Kaplan (1998) applies Duan's (1994, 2000) method to calculate risk-adjusted deposit insurance premiums for 15 Thai banks during the precrisis period of 1992–97. She finds that the cost of deposit insurance was highest for the banks that were nationalized, closed, subject to intervention, or sold to foreigners during the crisis period of 1998.

In this article I claim that the implicit cost of deposit insurance for a bank is a proxy for the risk taking of that bank, because the cost of insuring the deposits of a risky bank should be higher. The argument underlying this claim is that providing deposit insurance generates incentives for banks to take on risk, often reflected by excessive loan growth. This moral hazard behavior of banks has been described at length in the deposit insurance literature (for an excellent overview see Bhattacharya, Boot, and Thakor 1998). Merton (1977, 1978) first highlighted the attendant moral hazards of deposit insurance. Using a formal model, Bhattacharya and Thakor (1993) show that deposit insurance invites insured banks to seek excessive portfolio risk and to maintain liquid reserves lower than the social optimum. I estimate the cost of deposit insurance by calculating the risk-adjusted deposit insurance premium that a bank should have been paying under a risk-adjusted deposit insurance scheme, given its risk taking. This is the implicit deposit insurance premium. A high implicit cost of deposit insurance is taken as an indication of a risky bank. I use this approach both for countries with explicit deposit insurance and for countries with implicit insurance.

Because no country has actually implemented a market-based risk-adjusted deposit insurance scheme,[1] the risk-adjusted deposit insurance premium is fictitious. In fact, Chan, Greenbaum, and Thakor (1992) showed that it is impossible to implement a risk-sensitive deposit insurance pricing scheme that is incentive compatible unless banks are permitted access to rents, through explicit regulatory subsidies or restrictions on entry into banking.

In countries with an explicit deposit insurance scheme, the difference between the implicit premium and the premium that the bank actually pays to the deposit insurance fund indicates whether deposit insurance is under- or overpriced. Deposit insurance would be underpriced if the difference between the implicit premium (also known as the fair premium) and the actual premium is positive. I am not interested in the over- or underpricing of deposit insurance, but merely in estimating overall banking risks, so I focus on the implicit deposit insurance premiums.

Many countries do not have an explicit deposit insurance scheme in which every bank pays a certain premium to a deposit insurance fund. Nevertheless, most governments are expected to rescue troubled banks to protect depositors

1. Although most countries with an explicit deposit insurance scheme have designed flat-rate premiums, some countries, such as Argentina and the United States, have risk-sensitive premiums.

from losses, even if deposits are not officially insured. That is, these countries have implicit deposit insurance. In banking systems with implicit deposit insurance, the cost of the insurance is measured by the value of the deposit insurance put option.

II. Methodology

Merton's (1977) model of deposit insurance can be used to calculate the implicit cost of deposit insurance. Merton shows that the payoff of a third-party guarantee of payment to a firm's bondholders where there is no uncertainty about the guarantee obligation being met is identical to that of a put option where the promised payment corresponds to the exercise price and the value of the firm's assets V corresponds to the underlying asset.

Merton (1977) applies this model to a bank for which the debt issue corresponds to deposits. Because most deposits are of the demand type, the model's assumption of term debt issue is not strictly applicable. However, if one interprets the time until maturity as the time until the next audit of the bank's assets, then from the point of view of the guarantor, deposits can be treated as if they were term and interest bearing. Two more assumptions are made. First, it is assumed that deposits equal total bank debt and that both principal and interest are insured. Second, it is assumed that the bank's asset values follow geometric Brownian motion.

$$d\ln V_t = \mu dt + \sigma dW_t \tag{1}$$

where V is the value of assets, t is time, μ is the instantaneous expected return on assets, σ is the instantaneous expected standard deviation of asset returns, and W indicates a standard Wiener process. The Black and Scholes (1973) option pricing model can be used to value the deposit insurance per unit of deposits:

$$g = \Phi\left(\sigma\sqrt{T-t} - h_t\right) - \left[\frac{(1-\delta)V_t}{D}\right]\Phi(-h_t), \tag{2}$$

where $h_t = \dfrac{\ln[(1-\delta)V_t/D] + [\sigma^2/2](T-t)}{\left(\sigma\sqrt{T-t}\right)}$, g is the value of the deposit insurance guarantee per dollar of insured deposits, Φ is the cumulative normal distribution function, T is the time until maturity of the bank debt, D is the face value of the bank debt, and δ is the annualized dividend yield.

To implement the model, the two unobservable variables, the bank's asset value V and the volatility parameter σ, have to be estimated. Ronn and Verma (1986) suggest using two restrictions for the identification of these two unknowns. The first restriction is obtained by viewing the equity value of the bank, which is directly observable, as a call option on the bank's assets with a strike price equal to the value of the bank's debt:

$$E_t = V_t\Phi(d_t) - D\Phi\left(d_t - \sigma\sqrt{T-t}\right), \tag{3}$$

where $d_t = \frac{\ln[V_t / D] + [\sigma^2 / 2][T - t]}{(\sigma\sqrt{T - t})}$. Ronn and Verma (1986) modeled equity as being dividend protected; therefore dividends do not appear in equation (3). The Black-Scholes (1973) formula thus defines a one-to-one mapping between the unknown asset value and the observed equity value.

Ronn and Verma (1986) used the relationship between the equity and asset volatility, which can be obtained by applying Ito's Lemma to equation (3), as the second restriction

$$\sigma = (\sigma_E E_t) / [V_t \Phi(d_t)] \quad (4)$$

where σ_E is the standard deviation of equity returns.

Because the market value of equity is observable and the equity volatility can be estimated, two nonlinear restrictions are now in place for identifying the two unknowns. Using data on bank debt, bank equity, and equity volatility, equations (3) and (4) can be solved simultaneously for V and σ. Given these values, equation (2) is used to solve for the value of deposit insurance per dollar of deposits, which I interpret as the implicit cost of deposit insurance. For this approach to be valid, the time until maturity, T, of the put and call options must be the same. Ronn and Verma (1986) use Merton's (1977) assumption that the time until maturity of the debt is equal to the time until the next audit. They interpret the strike price of the put option as equal to the total debt of the bank, rather than to total deposits only. This assumes that all the debts of the bank are insured and that they are issued at the risk-free interest rate.

Ronn and Verma (1986) estimate instantaneous equity volatility by the sample standard deviation of daily equity returns and therefore impose the condition that equity volatility is constant. Duan (1994, 2000) points out that this premise is inconsistent with the underlying theoretical model of Merton (1977), in which equity volatility is stochastic. Therefore, the Ronn and Verma (1986) estimator does not possess the properties normally expected from a sound statistical procedure, such as consistency and efficiency.

Duan's (1994, 2000) maximum likelihood framework for estimating the value of deposit insurance is consistent with the assumption of Merton's (1977) theoretical model that equity volatility is stochastic. With the process in equation (1), the one-period transition density of the unobserved values of the bank's assets can be characterized by $\ln(V_{t+1} / V_t) \sim N(\mu,\sigma^2)$. Therefore, the log-likelihood function for a sample of unobserved V_t can be expressed as

$$\begin{aligned} L_v(V_t, t = 1,\ldots,n; \mu,\sigma) = &-[(n-1)/2]\ln(2\pi\sigma^2) - \sum_{t=2}^{n}\ln V_t \\ &-[1/2\sigma^2]\sum_{t=2}^{n}[\ln(V_t / V_{t-1}) - \mu]^2 \end{aligned} \quad (5)$$

Because the call option formula (equation [3]) is an element-by-element transformation from an unobserved sample of asset values to an observed time series of equity values, the log-likelihood function for the observed sample of equity values can be written as

$$L(E_t, t=1,\ldots,n;\mu,\sigma) = -[(n-1)/2]\ln(2\pi\sigma^2) - \sum_{t=2}^{n}\ln\left[\hat{V}_t(\sigma)\Phi(\hat{d}_t)\right] - (1/2\sigma^2)\sum_{t=2}^{n}\left[\ln\left(\hat{V}_t(\sigma)/\hat{V}_{t-1}(\sigma)\right)-\mu\right]^2, \tag{6}$$

where $\hat{V}_t(\sigma)$ is the unique solution to equation (3) for any σ, and $\hat{d}_t$ corresponds to d_t with $\hat{V}_t(\sigma)$ in place of V_t. In the expression above, I have used the fact that $\partial E_t / \partial V_t = F(d_t)$.

With the log-likelihood function in equation (6), an iterative optimization routine can be used to compute the maximum likelihood estimates. According to Duan (1994, 2000), these estimators are consistent. Given starting values for μ and σ and data on equity values E_t and debt D_t, equation (3) can be solved to yield a series of bank asset values V_t. Equation (6) is then used to solve for $\hat{\mu}$ and $\hat{\sigma}$. This process is iterated to find the maximum likelihood estimates of $\hat{\mu}$ and $\hat{\sigma}$ and their standard errors. Using the put option formula for deposit insurance (equation [2]), one can solve for the value of the guarantee per dollar of deposits and its standard error. The asymptotic distribution of the estimator for the deposit insurance premium is reported in Duan (1994, 2000). Although Duan (1994, 2000) correctly points out the deficiency of the Ronn and Verma (1986) method, I nevertheless apply both methods for comparative purposes. The main focus, however, is on the Duan (1994, 2000) estimates. (In the section on estimates I compare those obtained from applying these two methods.)

I calculate the deposit insurance premiums under the assumption that all bank debts (both deposits and other debt liabilities) are fully insured. Total bank liabilities are therefore used for the variable D in equation (3). This assumption is made for simplicity. In reality, banks carry both insured and uninsured debts. In particular, some deposit insurance schemes insure only certain types of deposits or provide only partial insurance by insuring up to a certain level. Nevertheless, given the bailout practices of deposit insurance funds around the world, a valid argument can be made that de facto insurance extends to all liabilities of an insured bank. Moreover, some countries have explicitly covered bank debt other than deposits.

I assume that the next audit of the bank will take place in one year and that the maturity of the debt also equals one year. I thus model deposit insurance as a limited-term contract. Because the government is likely to give the bank some forbearance after it finds out that the bank is undercapitalized, modeling deposit insurance as a one-year contract seems restrictive. Moreover, Pennacchi (1987) has shown that the assumption of a limited-term contract can lead to underestimates of the cost of insurance. However, because the level of regulatory control is unknown ex ante, I prefer to model deposit insurance as a limited-term contract, acknowledging that the cost of deposit insurance might be underestimated. As long as a possible underestimation is similar across banks, the method remains valid for comparative purposes. Moreover, it is likely that regulatory control is weaker in countries with weak banks, so that the cost of deposit insurance would be underestimated for the riskiest banks. Any comparative results found using a limited-term contract would thus probably have been even stronger had deposit insurance been modeled in a multiperiod environment. (For models that allow for unlimited-term contracts, see Pennacchi 1987 and Hovakimian and Kane 2000.)

I estimate annual equity volatility by using a sample of daily equity returns and following Fama (1965), who suggested ignoring days on which the exchange is closed. Observations are also excluded for days on which it is announced that the bank will be restructured, merged, or closed down, because such announcements tend to lead to large jumps in share prices, which have a distortionary effect on the estimated volatility of equity returns. These corrections imply that $\sigma_E = \sqrt{n}\sigma_{E,n}$ is used as the estimate of annualized equity volatility to compute the Ronn and Verma (1986) deposit insurance estimates, where n is the actual number of trading days per year minus the trading days on which large jumps occurred, and $\sigma_{E,n}$ is the bank's daily equity volatility based on n daily equity returns. In most countries n is around 252 days. In estimating the Duan (1994, 2000) deposit insurance premiums, I correct for missing data by accommodating the log-likelihood function in equation (6) accordingly. This correction was used by Duan and Yu (1994).

III. Data

The selection of economies and banks for the sample was based on several criteria. I wanted to focus on banks in emerging market economies, because these banks are thought to be riskiest and because they tend to have more diverse ownership structures. As a control group, I also wanted to include a number of highly developed economies, whose banks were expected to provide a benchmark for a low level of risk taking. Within each economy I had to restrict the sample to exchange-listed banks because I needed data on bank market capitalization and dividend yields.

Because the put option approach to valuing deposit insurance assumes that stock markets are efficient, the sample is limited to economies that have relatively large and liquid stock markets. The International Finance Corporation classifies 14 economies as emerging markets in which the total market capitalization of listed companies exceeded US$50 billion and the monthly turnover ratio 2 percent in mid-1999. The sample is also limited by excluding countries with heavily regulated financial sectors. To this end, the financial liberalization dates in Williamson and Mahar (1998) were used to exclude countries that had not started to liberalize their financial sectors before 1990.[2] The remaining sample of emerging market economies numbers 12. As a result of banking data limitations, related mostly to a lack of data on the ownership structure of banks, an-

2. This criterion excludes China and India.

other five countries were excluded.[3] The final sample of emerging market economies numbers seven: Argentina, Chile, Indonesia, the Republic of Korea, Malaysia, Taiwan (China), and Thailand.

This sample includes the four East Asian countries that experienced banking crises in 1997–98: Indonesia, Korea, Malaysia, and Thailand. In addition to Taiwan (China), several other Asian economies were included in the sample to examine whether implicit deposit insurance costs differ between economies that have been heavily affected by the 1997 East Asian financial crisis and those that have not been. These are Hong Kong (China), Japan, and Singapore, the only three economies in Asia considered to be developed. To assess the effects of the crisis, data are needed for the crisis years 1997–98 as well as for some years before the crisis. As a benchmark group for a low level of risk taking, the sample includes the four largest Western economies: France, Germany, the United Kingdom, and the United States.

Thus the final data set includes listed banks from 14 economies: 2 Latin American countries, the 4 East Asian crisis countries, 4 other economies in East Asia, and the 4 major Western economies. Across these 14 economies data were collected on 144 listed banks during the period 1991–98. The banks in the sample include most major listed banks in each economy.[4]

To limit the number of listed banks in Japan, the sample includes only the long-term credit banks (3), the city banks (9), and the trust banks (7) and thus excludes the mostly smaller regional banks (127). To limit the number of listed banks in the United States, the sample includes only the 22 largest U.S. banks: the multinational banks (6) and the super-regional banks (16) as defined by Goldman Sachs (2000).

Data on daily stock market capitalization and annualized dividend yields were collected from Datastream. The data range from 1991 to 1998 and thus include the East Asian crisis years 1997–98. Total deposits at year-end, net loans at year-end, and ownership data were taken from BankScope. For missing observations, Bloomberg was consulted. Ownership data were collected as follows. Four forms of concentrated ownership were distinguished: state-owned (the state, treasury, military, or another government institution owns shares in the bank), family-owned (a family or individual owns shares in the bank), company-owned (a manufacturing company owns shares in the bank), and owned by another financial institution (another financial institution owns shares in the bank). Banks with no concentrated owners (dispersed ownership) are classified as widely held. A number of ownership dummy variables were defined that are related to this classification of ownership and based on different thresholds of shareholdings. The threshold for a majority shareholding is 50 percent of shares and that for a major shareholding

3. These data limitations exclude Brazil, Greece, Israel, Mexico, and South Africa.

4. The distribution of banks across economies is as follows: Argentina (5), Chile (2), France (4), Germany (8), Hong Kong (China) (12), Indonesia (8), Japan (19), Korea (22), Malaysia (10), Singapore (5), Taiwan (China) (8), Thailand (12), the United Kingdom (7), and the United States (22).

is 20 percent. The BankScope data set is also used to construct a dummy variable that indicates whether the bank is affiliated with a business group or not. A bank is classified as group affiliated if it is either a subsidiary of a diversified business group or if more than 50 percent of its shares are held by a nonfinancial company.

For the 144 banks, 950 observations were collected, spanning eight years. Data are missing for 202 observations. These data are missing for several reasons. Some banks did not report accounting data for each year, some were listed on the exchange during only part of the sample period, and some were delisted during the sample period because of government intervention or merger activity.[5] Missing observations for 1998 are due largely to bank restructuring that took place after the East Asian financial crisis of 1997.

Country-specific data were also collected. Gross domestic product (GDP) per capita and inflation rates were taken from the International Monetary Fund's International Financial Statistics database. As a proxy for the quality and enforcement of a country's legal system, figures were taken from the law and order index of the *International Country Risk Guide,* published by the PRS Group. The law and order index ranges from 0 to 6, with higher values indicating higher quality (less risk). Law and order are assessed separately, with the value for each ranging from 0 to 3. The law subcomponent is an assessment of the strength and impartiality of the legal system, and the order subcomponent is an assessment of popular observance of the law. Data on bank concentration and foreign bank penetration were taken from the World Bank's Financial Structure Database. Finally, data were taken from Demirgüç-Kunt and Huizinga (1999) and Demirgüç-Kunt and Sobaci (2000) on the features of economies' deposit insurance schemes, particularly on whether insurance is implicit or explicit and on the size of the officially charged, explicit insurance premiums (table 1).

5. In Indonesia Bank Tiara Asia, a private foreign exchange bank, was taken over in 1998 by the Indonesian Bank Restructuring Agency and is therefore missing for 1998. Although Indonesian Bank Danamon merged with state-owned bank PDCFI in 1998, both banks continued reporting separately for another year, so Bank Danamon's 1998 data could be included. In Japan two long-term credit banks have been delisted—Long Term Credit Bank (October 26, 1998) and Nippon Credit Bank (December 14, 1998)—and nationalized. Because both banks reported 1998 deposit data, 1998 data for these two banks could be included as well. For Korea, Commercial Bank of Korea and Korea Long Term Credit Bank were excluded because they were not listed, and Donghwa Bank was excluded because it began operation only in 1996. The sample of Korean banks changes in 1998 because of merger activity. Commercial Bank of Korea and Hanil Bank merged in 1998, creating a new bank called Hanvit Bank, and on September 8, 1998, Hana Bank announced a merger with Boram Bank (to become effective in 1999). Korea First Bank was sold to New Bridge Capital (United States) as of December 30, 1998, although trading was not suspended until June 25, 1999; Kookmin Bank announced a merger with Korea Long Term Credit Bank on August 25, 1998. Accounting data for Seoul Bank continued to be reported until 1998, although the bank was nationalized in 1998 and subsequently sold to HSBC Bank on February 22, 1999. In Malaysia Kwong Yik Bank was acquired by RHB Capital and officially delisted on August 26, 1997, so it was not included. In Thailand lack of data required the exclusion of Laem Thong Bank, Nakornthon Bank, and Union Bank of Bangkok. In addition, for 1998 data are missing for Bangkok Bank of Commerce, which was closed and delisted that year, and for First Bangkok City Bank, which was acquired by the government in February 1998 and merged with state-owned Krung Thai Bank in 1999.

TABLE 1. Features of the Deposit Insurance Schemes in the Sample Economies

Economy	Type of scheme	Year established	Insurance premium (percentage of insured deposits)
Argentina	Explicit	1979	0.36–0.72 (risk based)
Chile	Explicit	1986	Callable
France	Explicit	1980	Callable, but limited
Germany	Explicit	1966	0.03, but can be doubled
Hong Kong (China)	Implicit	NA	NA
Indonesia	Implicit	NA	NA
Japan	Explicit	1971	0.04
Korea, Rep. of	Explicit	1996	0.05
Malaysia	Implicit	NA	NA
Singapore	Implicit	NA	NA
Taiwan (China)	Explicit	1985	0.015
Thailand	Implicit	NA	NA
United Kingdom	Explicit	1982	On demand (with a maximum of 0.3 percent)
United States	Explicit	1934	0.00–0.27 (risk based)

NA: Not applicable.

Note: If an economy has an explicit deposit insurance scheme, the table reports the year in which it was established and the size of the annual insurance premium. Korea had implicit deposit insurance before 1996.

Source: Demirgüç-Kunt and Sobaci (2000).

IV. DEPOSIT INSURANCE ESTIMATES

I calculate the annual implicit costs of deposit insurance as one-year put options on the value of bank assets for the 144 banks for each year in 1991–98 using both the Ronn and Verma (1986) method and the Duan (1994, 2000) method (table 2, panel 1). (Throughout the rest of the article RV indicates estimates based on the Ronn and Verma method, and Duan indicates estimates based on the Duan method.) At first sight the estimates produced by the two methods seem to differ widely. In particular, the RV estimates seem to be higher on average than the Duan estimates. Nevertheless, the correlation between the estimates from the two methods is 57 percent,[6] and Spearman's rank correlation is 85 percent.[7] These results indicate that although the two methods produce estimates that differ in size, they produce similar rankings. In other words, the methods tend to identify similar groups of banks as the riskiest.

Because the distribution of both estimates is highly skewed to the right because of some large positive outliers, I also compare the estimates once they have been

6. The high correlation is confirmed in a simple OLS regression with the Ronn and Verma (1986) estimates as the dependent variable and the Duan (1994, 2000) estimates as the explanatory variable. In fact, a Wald test does not reject (at the 5 percent significance level) the hypothesis that the regression coefficient of this regression differs from one. These regression results should be interpreted with caution, however, because measurement error in the explanatory variable causes the OLS estimates to be statistically inconsistent.

7. Note that these figures suffer from measurement error in the deposit insurance estimates.

TABLE 2. Implicit Deposit Insurance Costs for the Sample Banks Estimated Using Two Methods, 1991–98
(basis points of total bank debt)

	Panel 1		Panel 2	
	RV	*Duan*	*RV**	*Duan**
Summary statistics				
Mean	35.13	19.36	1.14	1.03
Median	0.42	0.08	0.35	0.09
Maximum	4,721.06	1,431.95	8.46	7.27
Minimum	0.00	0.00	0.00	0.00
SD	206.13	85.66	1.64	1.58
Skewness	15.00	10.40	1.76	1.55
Correlation				
Correlation coefficient	0.57		0.80	
Rank correlation coefficient (Spearman's rho)	0.85		0.85	

Note: *RV* indicates the deposit insurance cost estimated by applying the Ronn and Verma (1986) method, and *Duan* the deposit insurance cost estimated by applying the Duan (1994, 2000) method. In panel 2 the estimates are transformed as follows: $RV^* = \ln(1 + RV)$ and $Duan^* = \ln(1 + Duan)$.

Source: Author's calculations.

transformed by the log operator. Because the estimated cost of deposit insurance is zero for some banks, I first add one to each estimate of the cost of deposit insurance before applying the log operator. After this rescaling of the estimates, the results of the two methods are more similar (table 2, panel 2). The correlation is around 80 percent, and the rank correlation around 85 percent.[8]

Despite strong rank correlation, the results of the comparison indicate some remaining variation between the two estimates. I therefore conclude that the RV method and the Duan method produce different estimates of the cost of deposit insurance. This result was found earlier by Duan and Yu (1994), although their assessment is restricted by the small number of banks in their study (10, compared with 144 in my analysis).

In the subsequent analysis I focus on the Duan estimates of the cost of deposit insurance, because they are theoretically and statistically superior. This means that all findings are based on these estimates. The Duan method has the added advantage of allowing estimation of the standard error of the deposit insurance cost estimates. For comparative purposes, I also report the RV estimates.

Estimates of the implicit cost of deposit insurance averaged by year show that for most economies in the sample the cost of deposit insurance increases over the period, from an average of 7 basis points in 1991 to 62 in 1998. More specifically, the average cost of deposit insurance is higher during the crisis period 1997–98 than during the precrisis years (table 3).

8. Again, it should be noted that these figures suffer from measurement error.

TABLE 3. Estimated Implicit Deposit Insurance Costs across Years, Economies, and Ownership Forms, 1991–98
(basis points of total bank debt)

Across years				Across economies				Across ownership forms			
Year	*RV*	*Duan*	No.	Economy	*RV*	*Duan*	No.	*Owner20*	*RV*	*Duan*	No.
1991	2.12	6.66	71	Argentina	31.36	17.81	25	*Company*	82.14	43.93	111
	(4.80)	(22.93)			(66.09)	(58.43)			(289.06)	(107.22)	
1992	4.68	3.40	88	Chile	0.02	0.00	8	*Family*	106.42	56.58	78
	(9.25)	(7.63)			(0.04)	(0.01)			(551.70)	(212.82)	
1993	1.03	2.70	116	France	2.37	7.72	29	*OtherFI*	54.67	23.02	79
	(3.15)	(18.24)			(4.98)	(12.93)			(149.55)	(55.57)	
1994	1.22	3.19	129	Germany	0.18	6.17	54	*State*	35.67	16.53	63
	(2.84)	(17.38)			(0.51)	(17.20)			(161.12)	(36.95)	
1995	5.75	5.47	136	Hong Kong (China)	37.85	13.74	79	*Widely*	15.06	10.04	627
	(29.93)	(28.32)			(98.67)	(31.61)			(73.96)	(51.44)	
1996	0.79	4.04	138	Indonesia	154.37	83.99	55				
	(2.33)	(20.84)			(412.59)	(147.90)					
1997	35.30	53.20	143	Japan	12.43	13.91	149				
	(72.96)	(138.37)			(69.95)	(55.33)					
1998	206.20	61.88	129	Korea, Rep. of	36.58	20.13	125				
	(522.44)	(163.04)			(89.12)	(88.60)					
				Malaysia	25.85	20.86	60				
					(81.91)	(45.67)					

Singapore	5.98	0.35	37
	(28.79)	(0.90)	
Taiwan	1.34	3.81	57
(China)	(2.22)	(10.41)	
Thailand	135.95	58.26	93
	(530.62)	(196.76)	
United	1.34	2.29	48
Kingdom	(3.56)	(7.15)	
United	0.40	0.63	131
States	(1.44)	(2.71)	
Average	35.13	19.36	950
	(206.13)	(85.66)	

Note: The cost of deposit insurance across years is averaged over all banks and across all economies. The cost for each economy is averaged over all banks in the economy and across all years. The cost across ownership forms is averaged over all banks in the ownership category and across all years. The variable *Owner20* is identical to "company" if a company owns more than 20 percent of the shares, "family" if a family owns more than 20 percent, "otherFI" if another financial institution owns more than 20 percent, "state" if a government institution owns more than 20 percent, and "widely" if no concentrated group owns more than 20 percent. Standard deviations of the costs of deposit insurance are in parentheses. No. refers to the number of observations in each category.

Source: Author's calculations.

Over the sample period the cost of deposit insurance (averaged across all banks in the economy and over all years) is highest for the four East Asian crisis countries: Indonesia (84 basis points), Thailand (58), Malaysia (21), and Korea (20). The cost is lowest for the four highly developed Western economies—the United States (0.6), the United Kingdom (2.3), Germany (6.2), and France (7.7)—as well as for highly developed Singapore (0.4) and for Chile (0.0) and Taiwan (China) (3.8). In Taiwan (China) the financial system is predominantly state-owned and banking is heavily regulated, which might explain the result suggesting that Taiwanese banks take low risks. The low estimate for Chile may not accurately reflect the riskiness of the average Chilean bank, because the sample includes only two Chilean banks. The implicit deposit insurance premiums calculated for banks in Hong Kong (China) (14 basis points), Japan (14), and Argentina (18) are somewhere in the middle.

The estimates of deposit insurance cost indicate that risk taking also differs across forms of ownership. The cost estimates (averaged over all banks in the economy and over all years) are as follows: family (57 basis points), company (44), other financial institution (23), state (17), and widely held (10). These figures indicate that concentrated ownership links between banks and other parties, such as in the Japanese *keiretsu* or the Korean *chaebol*, increase risk taking by banks and that dispersed ownership of banks is to be preferred. State ownership has an intermediate impact on a bank's risk taking. Note that not all economies have banks in all five ownership categories. In Western countries, for example, most banks are widely held. On the other extreme, in Indonesia most banks have concentrated ownership, with 32 percent of Indonesian banks in the sample having an owner that holds at least 20 percent of shares.

Group affiliation also increases the cost of deposit insurance. For the 35 banks in the sample that are affiliated with a business group, the cost averages 45 basis points, whereas the cost for the nonaffiliated banks averages 18 basis points.

V. Empirical Analysis

In the previous section I quickly interpreted the summary statistics of the calculated implicit costs of deposit insurance. Although these summary statistics show some clear patterns, in this section I conduct a more accurate analysis of the differences in the cost of deposit insurance across economies, periods, and ownership forms using econometric techniques to control for bank-specific effects. I transform the variables with the log operator and estimate a log-linear model. Because the cost of deposit insurance is estimated to be zero for some banks, I use $\ln(1 + Cost)$ as the dependent variable, rather than $\ln(Cost)$, where *Cost* is the implicit cost of deposit insurance in basis points of total debt, calculated using either the RV method or the Duan method. With the transformed estimate of the implicit cost of deposit insurance as the dependent variable, I estimate a series of ordinary least squares (OLS) regression models. Although the costs of deposit insurance are estimates, measurement error in the dependent variable can be absorbed in the disturbance of the regression and ignored. The results are presented with White's (1980) heteroskedasticity-consistent standard errors.

Ownership, Size, and Credit Growth

First I regress the cost of deposit insurance on dummy variables for dispersed ownership, country, and year. The dispersed ownership dummy variable takes the value one if no shareholder owns more than 5 percent of the shares in the bank, and zero otherwise. The country dummy variables control for differences in institutional environments across economies. The United States and the year 1991 are used as benchmark variables to prevent multicolinearity.

The results show that the cost of deposit insurance in 1991–98 is higher on average for banks in Argentina, France, Germany, Hong Kong (China), Indonesia, Japan, the Republic of Korea, Malaysia, Taiwan (China), Thailand, and the United Kingdom than for banks in Chile, Singapore, and the United States (table 4, model [a]). Notably, the cost of deposit insurance is relatively high for banks in the financial crisis countries. In the sample period the cost is highest for Indonesian banks—around 7.7 basis points higher than for U.S. banks.[9] For Thai, Korean, and Malaysian banks the cost is 3.4, 1.9, and 1.5 basis points higher than for U.S. banks.

The cost of deposit insurance is relatively high in 1997 and 1998—5.0 and 5.5 basis points, respectively, higher than in 1991. This result is expected because 1997 and 1998 are the East Asian crisis years. Controlling for country and time effects, I find that the cost of deposit insurance for widely held banks is 0.2 basis points lower than for banks with concentrated ownership. I find similar results if I use a dispersed ownership dummy variable that takes the value one if no shareholder owns more than 20 percent of the shares in the bank rather than a dummy variable using 5 percent as the cutoff.

To control for bank-specific size effects, I add the amount of net loans outstanding at the end of the year as a variable to the previous model. The results are similar (table 4, model [b]). Again, the cost of deposit insurance for widely held banks is 0.2 basis points lower than for banks with concentrated ownership. In addition, the cost is higher for small banks, with bank size measured by total net loans outstanding. The size effect is only marginal, however. For example, all other things equal, a 10 percent increase in loans would lead to a 1.2 percent decrease in the cost of deposit insurance. A possible explanation for this

9. The effect on the cost of deposit insurance is calculated as $\exp(\beta) - 1$, where β is the coefficient of the respective dummy variable. The effect is an average effect for the sampled banks in the economy over the sample period.

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TABLE 4. Deposit Insurance Cost and Dispersed Ownership

	(a)		(b)		(c)		(d)	
Variable	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*
Constant	−0.342**	−0.134	2.378***	1.875***	1.820***	1.535***	0.105	−0.071
	(0.164)	(0.196)	(0.553)	(0.657)	(0.417)	(0.584)	(0.103)	(0.163)
Argentina	1.505***	0.864**	0.981**	0.475	1.850***	1.044***	1.811***	0.020
	(0.434)	(0.340)	(0.442)	(0.384)	(0.467)	(0.423)	(0.649)	(0.285)
Chile	−0.269	−0.408	−0.652*	−0.692**	−0.286	−0.465***	−0.136	−0.563***
	(0.388)	(0.355)	(0.388)	(0.355)	(0.107)	(0.133)	(0.086)	(0.131)
France	0.455***	0.866***	0.584***	0.959***	0.127	0.771**	−0.003	0.885**
	(0.134)	(0.310)	(0.141)	(0.318)	(0.091)	(0.319)	(0.065)	(0.352)
Germany	0.082	0.451**	0.192	0.530**	−0.102	0.148	−0.133**	0.219
	(0.140)	(0.216)	(0.150)	(0.222)	(0.063)	(0.193)	(0.056)	(0.204)
Hong Kong	1.052***	0.725***	0.743***	0.492***	−0.080	−0.217	0.083	−0.220
(China)	(0.171)	(0.175)	(0.177)	(0.181)	(0.083)	(0.145)	(0.068)	(0.147)
Indonesia	2.336***	2.160***	1.838***	1.792***	1.001***	0.762***	1.222***	0.591**
	(0.217)	(0.253)	(0.232)	(0.269)	(0.198)	(0.249)	(0.199)	(0.259)
Japan	0.984***	1.012***	1.223***	1.186***	0.461***	0.237**	0.439***	0.313***
	(0.108)	(0.122)	(0.119)	(0.141)	(0.072)	(0.118)	(0.068)	(0.098)
Korea,	1.677***	1.063***	1.390***	0.855***	0.505***	0.018	0.669***	0.103
Rep. of	(0.120)	(0.130)	(0.130)	(0.147)	(0.076)	(0.106)	(0.073)	(0.122)
Malaysia	1.060***	0.917***	0.628***	0.576***	0.042	−0.336**	0.151	−0.278
	(0.168)	(0.193)	(0.188)	(0.216)	(0.139)	(0.147)	(0.138)	(0.171)
Singapore	0.203	−0.084	−0.047	−0.272*	−0.365***	−0.512***	−0.231***	−0.478***
	(0.153)	(0.159)	(0.161)	(0.165)	(0.076)	(0.102)	(0.071)	(0.120)
Taiwan	0.363*	0.482**	0.192	0.353*	0.358***	0.245	0.283***	−0.074
(China)	(0.201)	(0.197)	(0.204)	(0.198)	(0.122)	(0.172)	(0.106)	(0.119)
Thailand	1.878***	1.477***	1.648***	1.303***	0.576***	0.169	0.617***	0.188
	(0.183)	(0.197)	(0.184)	(0.198)	(0.106)	(0.167)	(0.110)	(0.192)
United	0.309**	0.240	0.465***	0.354**	0.155**	−0.003	0.074	0.047
Kingdom	(0.131)	(0.173)	(0.132)	(0.177)	(0.067)	(0.149)	(0.064)	(0.157)

1992	0.440***	0.094	0.444***	0.101	0.450***	0.101	0.584***	0.759***
	(0.141)	(0.182)	(0.144)	(0.186)	(0.127)	(0.170)	(0.126)	(0.173)
1993	-0.065	-0.311*	-0.049	-0.300*	-0.158	-0.423***	-0.019	0.190
	(0.133)	(0.169)	(0.133)	(0.170)	(0.104)	(0.155)	(0.103)	(0.155)
1994	-0.015	-0.039	0.023	-0.011	-0.105	0.159	0.018	0.447***
	(0.129)	(0.165)	(0.129)	(0.167)	(0.101)	(0.152)	(0.102)	(0.154)
1995	0.057	-0.086	0.106	-0.050	-0.035	-0.206	0.108	0.418***
	(0.131)	(0.168)	(0.131)	(0.171)	(0.101)	(0.154)	(0.097)	(0.153)
1996	-0.197	-0.117	-0.121	-0.061	-0.283***	-0.237	-0.136	0.455***
	(0.121)	(0.167)	(0.123)	(0.170)	(0.098)	(0.160)	(0.926)	(0.169)
1997	1.538***	1.786***	1.587***	1.814***	—	—	—	—
	(0.157)	(0.197)	(0.158)	(0.199)				
1998	2.734***	1.869***	2.796***	1.915***	—	—	—	—
	(0.188)	(0.196)	(0.190)	(0.198)				
Widely5	-0.182**	-0.203**	-0.159**	-0.189**	-0.128***	-0.284***	-0.128***	-0.297***
	(0.082)	(0.096)	(0.081)	(0.097)	(0.051)	(0.087)	(0.053)	(0.092)
Loan	—	—	-0.161***	-0.118***	-0.089***	-0.051	—	—
			(0.031)	(0.038)	(0.024)	(0.035)		
Loan growth	—	—	—	—	—	—	0.275	0.715**
							(0.207)	(0.341)
Adjusted R^2	0.583	0.431	0.593	0.593	0.410	0.151	0.375	0.122
Observations	950	950	944	944	673	673	569	569

— Not available.

*Significant at the 10 percent level.

**Significant at the 5 percent level.

***Significant at the 1 percent level.

Note: The dependent variable is ln(1 + *Insurance*), where *Insurance* is the cost of deposit insurance in basis points of total debt calculated using either the RV method or the Duan method. *Widely5* is a dummy variable that takes the value one if no shareholder owns more than 5 percent of the shares in the bank, and zero otherwise. *Loan* is ln(*Loan*), where *Loan* is the amount of net loans outstanding at year-end in thousands of U.S. dollars. *Loan growth* is ln(1 + *Dloan*), where *Dloan* is the growth of net loans during the year. The United States provides the benchmark for the country effects, and 1991 the benchmark for the year effects. In addition to country and year effects, model (a) controls only for dispersed ownership. Model (b) adds net loans to control for the size effect. Models (c) and (d) are estimated using 1991–96 data only. Model (c) controls for loan size, and model (d) for credit growth. Heteroskedasticity-consistent standard errors are in parentheses.

Source: Author's calculations.

finding is that the fierce competition many small banks face from large banks may lead the smaller banks to take on more risk.

To check robustness, I repeat the previous analysis while excluding the crisis years 1997–98 (table 4, model [c]). Again, the cost of deposit insurance for widely held banks is significantly lower, this time by about 0.3 basis points. However, the small size effect that I found for the entire sample period is not significant for the period 1991–96.

Because banks often take risks in the form of excessive credit growth, I also estimate a model that controls for loan growth. If the cost of deposit insurance correctly indicates a bank's risk taking, there should be a strong correlation between credit growth and the cost of deposit insurance. The results show that the cost of deposit insurance is larger for banks with high loan growth (table 4, model [d]). For example, all other things equal, a 10 percent increase in credit growth leads on average to a 7.2 percent increase in the cost of deposit insurance. Excessive credit growth has been cited by many as a major factor in the banking crisis that unfolded in East Asia during 1997. This finding is therefore unsurprising, because the sample includes a large number of East Asian banks, some of which we now know took excessive risks.

We have already seen that concentrated ownership leads to a higher cost of deposit insurance. To find out whether the cost of deposit insurance differs across categories of concentrated ownership, I regress the cost on the different categories. I include the ownership dummy variables with absolute majority shareholdings (that is, larger than 50 percent) and the ownership variables with major shareholdings (20–50 percent). For the entire sample period (1991–98) the cost of deposit insurance is higher for banks with majority shareholdings by companies (around 0.9 basis points higher than for banks without concentrated ownership) and other financial institutions (around 0.7 basis points higher; table 5, model [a]). In addition, there is weak evidence that the cost of deposit insurance is higher for banks that are majority owned by the state. Although the effect is statistically significant only at the 11 percent level, the estimated difference (0.8 basis points) is substantial. These banks might have greater access to the safety net not only because they are riskier but also because they have better connections.

When the amount of net loans outstanding is added to the model specification to control for bank-specific size effects, the higher cost of deposit insurance for majority state-owned banks becomes statistically significant. The cost is now 0.9 basis points higher than for widely held banks, even higher than for banks with majority shareholdings by companies and other financial institutions (table 5, model [b]). For these banks the cost of deposit insurance is 0.8 and 0.6 basis points higher than for widely held banks. In addition, the cost of deposit insurance for small banks is slightly higher than for large banks.

For a robustness check, I also exclude the crisis years. The results for 1991–96 again show that majority shareholdings by companies increase risk (table 5, model [c]). But majority shareholdings by other financial institutions or the state no longer increase risk, nor are small banks riskier. Instead, the results show that

TABLE 5. Deposit Insurance Cost and Majority Ownership

	(a)		(b)		(c)	
Variable	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*
State20	0.249	0.164	0.239	0.155	0.234	0.041
	(0.241)	(0.237)	(0.244)	(0.242)	(0.161)	(0.203)
State50	−0.109	0.834	0.001	0.920*	−0.011	1.041
	(0.320)	(0.520)	(0.322)	(0.522)	(0.164)	(0.673)
OtherFI20	0.275	0.059	0.249	0.044	0.182	−0.029
	(0.180)	(0.191)	(0.182)	(0.190)	(0.183)	(0.176)
OtherFI50	0.855***	0.687***	0.752***	0.579**	0.306*	0.241
	(0.265)	(0.254)	(0.267)	(0.254)	(0.175)	(0.195)
Family20	−0.234	0.222	−0.272	0.190	0.070	0.512**
	(0.241)	(0.270)	(0.239)	(0.270)	(0.161)	(0.245)
Family50	0.292	0.254	0.148	0.131	0.319	0.309
	(0.442)	(0.410)	(0.445)	(0.402)	(0.378)	(0.391)
Company20	0.068	0.003	0.058	0.004	−0.111	0.089
	(0.175)	(0.199)	(0.174)	(0.198)	(0.162)	(0.163)
Company50	0.923***	0.898***	0.775***	0.778**	0.568***	0.789**
	(0.263)	(0.317)	(0.264)	(0.319)	(0.220)	(0.395)
Loan	—	—	−0.148***	−0.117***	−0.076***	−0.048
			(0.032)	(0.038)	(0.025)	(0.034)
Adjusted R^2	0.593	0.645	0.600	0.440	0.418	0.159
Observations	950	950	944	944	673	673

— Not available.
*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: The dependent variable is ln(1 + *Insurance*), where *Insurance* is the cost of deposit insurance in basis points of total debt calculated using either the RV method or the Duan method. *Loan* is the log of net loans outstanding at year-end. *State20* is a dummy variable that takes the value one if the state owns 20–50 percent of the shares in the bank; *state50* indicates 50–100 percent state ownership. *Family20* is a dummy variable that takes the value one if a family owns 20–50 percent; *family50* indicates 50–100 percent family ownership. *OtherFI20* is a dummy variable that takes the value one if another financial institution owns 20–50 percent; *otherFI50* indicates 50–100 percent ownership by another financial institution. *Company20* is a dummy variable that takes the value one if a company owns 20–50 percent; *company50* indicates 50–100 percent company ownership. For models (a–d) a constant term and country and year dummy variables were added but are not reported. The United States provides the benchmark for the country effects, and 1991 the benchmark for the year effects. In addition to country and year effects, model (a) controls only for majority ownership effects. Model (b) includes net loans to control for size effects. Model (c) is identical to model b but is estimated for 1991–96 only. Heteroskedasticity-consistent standard errors are in parentheses.

Source: Author's calculations.

major shareholdings by families[10] or individuals increase risk, although to a smaller extent than for companies.

I also compare the cost of deposit insurance for banks affiliated with a business group with the cost for other banks. Because banks affiliated with a busi-

10. Family ownership of firms is common in emerging market economies, particularly in East Asia. Claessens, Djankov, and Lang (2000) show that there is extensive family control in more than half of East Asian corporations and that managers of closely held firms tend to be relatives of the controlling shareholder's family.

ness group might be prepared to support a group member facing financial distress, the cost of deposit insurance for such banks could be expected to be higher than the cost for nonaffiliated banks. I classify a bank as group affiliated if the bank is a subsidiary of a diversified business group or if a nonfinancial company holds more than 50 percent of its shares. I regress the cost of deposit insurance on loan size and on a dummy variable that takes the value one if the bank is group affiliated and zero otherwise. I also include country and year dummy variables. I estimate this regression model both for the entire sample period and for the precrisis period 1991–96. For both periods the cost of deposit insurance is significantly higher for group-affiliated banks, suggesting that those banks might have supported some group members (table 6).

Country-Specific Factors

Thus far I have used country dummy variables to control for differences across economies. In this section I expand the model with country-specific variables. In a first specification I control for differences in two macroeconomic fundamentals: GDP per capita and the inflation rate. Banking systems can be expected to be less risky in economies with high GDP per capita and low inflation. I find that this is indeed the case for the sample when I regress the cost of deposit insurance on a constant, year dummy variables, a dispersed ownership dummy variable, net loans outstanding, GDP per capita, and inflation (table 7, model [a]). The findings are of economic importance. A 10 percent increase in GDP per capita

TABLE 6. Deposit Insurance Cost and Group Affiliation

	(a)		(b)		(c)	
Variable	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*
Group	0.622***	0.651**	0.512**	0.573*	0.476**	0.634*
	(0.251)	(0.301)	(0.251)	(0.301)	(0.214)	(0.387)
Loan	—	—	–0.153***	–0.110***	–0.082***	–0.043
			(0.032)	(0.038)	(0.024)	(0.034)
Adjusted R^2	0.586	0.434	0.594	0.437	0.415	0.151
Observations	950	950	944	944	673	673

— Not available.
*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: The dependent variable is ln(1 + *Insurance*), where *Insurance* is the cost of deposit insurance in basis points of total debt calculated using either the RV method or the Duan method. *Group* is a dummy variable that takes the value one if the bank is affiliated with a group, and zero otherwise. A bank is classified as group affiliated if it is a subsidiary of a diversified business group or if a nonfinancial company holds more than 50 percent of its shares. *Loan* is the logarithm of net loans outstanding at year-end. For each model a constant term and country and year dummy variables were added but are not reported. The United States provides the benchmark for the country effects, and 1991 the benchmark for the year effects. Models (b) and (c) include net loans to control for size effects. Models (a) and (b) are estimated for the full sample period, and model (c) for the precrisis years 1991–96 only. Heteroskedasticity-consistent standard errors are in parentheses.

Source: Author's calculations.

TABLE 7. Deposit Insurance Cost and Macroeconomic and Institutional Variables

	(a)		(b)		(c)	
Variable	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*
State20	0.249	0.164	0.239	0.155	0.234	0.041
Widely5	–0.143*	–0.238***	–0.246***	–0.334***	–0.300***	–0.456***
	(0.078)	(0.091)	(0.078)	(0.089)	(0.084)	(0.117)
Loan	–0.028	0.060*	–0.100***	–0.008	–0.121***	–0.048
	(0.028)	(0.032)	(0.025)	(0.030)	(0.033)	(0.045)
GDP *per capita*	–0.352***	–0.319***	—	—	—	—
	(0.051)	(0.058)				
Inflation	0.410***	0.368***	—	—	0.223***	0.389***
	(0.082)	(0.081)			(0.085)	(0.096)
Law and order	—	—	–2.315***	–2.034***	–1.371***	–1.423***
			(0.366)	(0.368)	(0.445)	(0.556)
Explicit	—	—	—	—	–0.230	–0.002
					(0.157)	(0.177)
Concentration	—	—	—	—	–0.247**	–0.076
					(0.102)	(0.145)
Foreign	—	—	—	—	–0.174***	–0.214***
					(0.041)	(0.052)
Adjusted R^2	0.525	0.389	0.492	0.359	0.349	0.285
Observations	944	944	944	944	688	688

— Not available.
*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: The dependent variable is ln(1 + *Insurance*), where *Insurance* is the cost of deposit insurance in basis points of total debt calculated using either the RV method or the Duan method. *Widely5* is a dummy variable that takes the value one if no shareholder owns more than 5 percent of the shares in the bank, and zero otherwise. *Loan* is the log of the amount of net loans outstanding at year-end. GDP *per capita* is the logarithm of GDP per capita in U.S. dollars. *Inflation* is ln(1 + *Infl*), where *Infl* is the inflation rate in percentage points based on the economy's consumer price index. *Law and order* is the logarithm of the law and order index of the PRS Group, which ranges from 0 to 6 (with 6 indicating an excellent system of law and order). *Explicit* is a dummy variable that takes the value one if the economy has explicit deposit insurance, and zero otherwise. *Concentration* is the log of the ratio of the three largest banks' assets to total banking assets. *Foreign* is the logarithm of the share of foreign bank assets in total banking sector assets. A constant term and year dummy variables were added but not reported. The year 1991 provides the benchmark for the year effects. In addition to year effects, model (a) controls for dispersed ownership, loan size, GDP per capita, and inflation effects. Model (b) controls for dispersed ownership, loan size, and quality of legal system effects. Model (c) controls for dispersed ownership, loan size, the quality of the legal system, the existence of explicit deposit insurance, the concentration ratio, and foreign entry. Model c uses data for 1991–97, because no competition data are available for 1998. Heteroskedasticity-consistent standard errors are in parentheses.

Source: Author's calculations.

leads to a reduction in the cost of deposit insurance of around 3 percent, and a 10 percent decrease in the inflation rate causes a reduction of around 4 percent. In addition, I again find that widely held banks are less risky. However, contrary to the earlier finding, large banks now take on slightly more risk than smaller banks. But this result is not significant at the 5 percent level.

In a second specification I control for differences across economies in the quality and enforcement of laws. Banking systems in economies with poor legal systems are expected to be riskier. This is indeed the case for the sample when the cost of deposit insurance is regressed on a constant, year dummy variables, a dispersed ownership dummy variable, net loans outstanding, and a law and order index (table 7, model [b]). The relationship between the law and order index and the cost of deposit insurance has economic importance. For example, other things equal, an improvement in the law and order index from 3 to 4, or an increase of 25 percent in the rescaled index, is estimated to cause a reduction in the cost of deposit insurance of around 50 percent. Widely held banks are again found to be less risky. This time there is no loan size effect.

In a third specification I add a variable indicating whether the economy has an explicit deposit insurance scheme or not, a variable measuring the concentration of the banking market, and a variable measuring penetration by foreign banks. This specification looks at the combined effect of the macroeconomic environment, the quality of the legal system, the type of deposit insurance scheme, and the market structure of the banking sector on the cost of deposit insurance. To indicate the type of deposit insurance, I use a dummy variable that takes the value one if the economy has an explicit deposit insurance scheme and the value zero if it has an implicit scheme. Nine of the 14 sample economies have an explicit deposit insurance scheme; the other 5 have implicit schemes. I measure the concentration of the banking market by the share of the three largest banks' assets in total banking sector assets, and the penetration of foreign banks by the share of foreign bank assets in total banking sector assets. Because the measure of the quality of the legal system—the law and order index—is highly correlated with GDP per capita (with a correlation of 0.80), I include only the inflation rate to control for the macroeconomic environment. I also add a constant, year dummy variables, a dispersed ownership dummy variable, and net loans outstanding to the model.

The results show that the cost of deposit insurance is lower in economies with low inflation rates and sound quality and enforcement of laws (table 7, model [c]). In addition, foreign bank penetration reduces the cost of deposit insurance. The estimated regression coefficients show that a 10 percent increase in the presence of foreign banks would reduce the cost of deposit insurance by 2 percent. But neither the degree of bank concentration nor the type of deposit insurance scheme has an impact on the cost of deposit insurance. Widely held banks are again found to be less risky.

The economies with explicit insurance are the most highly developed ones in the sample, and the correlation between the explicit dummy variable and the law and order index is 0.51. Any difference between the impact of explicit insurance and that of a good institutional environment on the cost of deposit insurance should therefore be interpreted with caution. The findings do suggest, however, that it is not the type of deposit insurance scheme that matters for the cost of deposit insurance—and for the riskiness of a banking system—but the overall quality and enforcement of rules. This finding does not necessarily contradict Demirgüç-Kunt and Detragiache (1999), who provide empirical evidence showing that explicit deposit insurance increases banking system vulnerability in countries with weak institutional environments.

In addition to moral hazard, explicit deposit insurance schemes can lead to fiscal problems if the premiums charged to the banks are underpriced. I therefore investigate whether the economies in the sample with explicit schemes underprice deposit insurance by setting premiums too low. The official deposit insurance premiums in the sample economies range from 0.0 percent to 0.72 percent of insured deposits (table 8).

In two countries with explicit deposit insurance schemes, Japan and Korea, the premiums actually charged are substantially lower than the average implicit cost of deposit insurance over the period 1991–98. The differences between actual and fair premiums do not differ statistically from zero at any reasonable level of significance, however. So for the economies in the sample with explicit deposit insurance schemes, it cannot be concluded that the official premiums were inadequate, although some banks in the sample probably should have been charged higher premiums to reflect their risks.

TABLE 8. Risk-Adjusted and Official Deposit Insurance Premiums in Sample Economies with Explicit Schemes, 1991–98
(basis points of deposits)

Economy	Risk-adjusted premium (*RV*)	Risk-adjusted premium (*Duan*)	Official premium
Argentina	31.36	17.81	36.0–72.0 (risk based)
	(66.09)	(58.43)	
Germany	0.18	6.17	3.0
	(0.51)	(17.20)	
Japan	12.43	13.91	4.0
	(70.00)	(55.33)	
Korea, Rep. of	36.60	20.13	5.0
	(89.10)	(45.67)	
Taiwan (China)	1.34	3.81	1.5
	(2.22)	(10.41)	
United States	0.40	0.63	0.0–27.0 (risk based)
	(1.44)	(2.71)	

Note: Standard deviations are in parentheses. Note that Korea had implicit deposit insurance before 1996.

Source: For official premiums, Demirgüç-Kunt and Sobaci (2000).

Forecasting Bank Distress

If the cost of deposit insurance is indeed closely related to the risk taking of a bank, as I argue, this proxy for bank risk should have power in predicting bank distress. As a first assessment of the information embedded in the cost of deposit insurance as a proxy for bank risk, I compare the cost of deposit insurance for the sample banks in 1996 across economies to check whether this measure indicates which economies were at risk of a banking crisis. Only for Indonesia, Korea, and Thailand is the average cost of deposit insurance in 1996 significantly higher than zero in both economic and statistical terms. Along with Malaysia, these are the East Asian countries that experienced a banking crisis one year later. Thus the implicit cost of deposit insurance in 1996 correctly indicates banking problems in three of the four East Asian countries that experienced a banking crisis in 1997.

In this section I explore the ex ante power of deposit insurance cost estimates to forecast bank distress in more detail and at a bank level using information on actual bank distress. Governments intervened in banks across East Asia in 1998 (table 9). I analyze the link between intervention in banks and the cost of deposit insurance before 1998 for the sample of banks to assess the power of the method to forecast bank problems.

First I analyze whether the cost of deposit insurance is indeed higher for banks in which the government intervened. I regress the cost of deposit insurance in 1998 on the cost in 1997 and add a dummy variable indicating whether the bank was subject to intervention. As expected, the cost of deposit insurance is higher for banks subject to intervention (table 10). With the difference equal to 91 basis points, the effect is economically significant. Interestingly, the implicit deposit insurance premium for banks not subject to intervention, as measured by the Duan method,

TABLE 9. Banks Subject to Intervention in Selected East Asian Economies, 1998

Economy	Banks
Indonesia	Bank Bali, Bank Danamon, Bank International Indonesia, Bank Lippo, Bank Niaga, Bank Tiara Asia
Japan	Long Term Credit Bank, Nippon Credit Bank
Korea, Rep. of	Cho Hung Bank, Chung Chong Bank, Dae Dong Bank, Dong Nam Bank, Hana Bank, Hanil Bank, Housing and Commercial Bank, Kookmin Bank, Koram Bank, Korea First Bank, Kyungki Bank, Seoul Bank, Shinhan Bank
Malaysia	AMMB Holdings, RHB Capital
Thailand	Bangkok Bank, Bangkok Bank of Commerce, Bangkok Metropolitan Bank, Bank of Asia, Bank of Ayudhya, DBS Thai Danu Bank, First Bangkok City Bank, Krung Thai Bank, Siam City Bank, Siam Commercial Bank, Union Bank of Bangkok, Thai Farmers Bank, Thai Military Bank

Note: Interventions include closure, recapitalization, nationalization, sale to foreigners, and domestic takeovers. The table includes only banks in the sample.

Source: Bongini, Claessens, and Ferri (2001) for Indonesia, Korea, Malaysia, and Thailand. Peek and Rosengren (2001), table 1, for Japan.

TABLE 10. Deposit Insurance Cost and Banks Subject to Intervention

Variable	*RV*	*Duan*
Constant	35.56**	–4.12
	(16.47)	(4.67)
Insurance 1997	1.44*	1.04***
	(0.87)	(0.17)
Intervention	572.82***	91.25***
	(165.27)	(42.08)
Adjusted R^2	0.282	0.393
Observations	128	122

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.
Note: The dependent variable is the cost of deposit insurance in basis points of total debt calculated for 1998 using either the RV method or the Duan method. *Insurance 1997* is the cost of deposit insurance in basis points of total debt for 1997. *Intervention* is a dummy variable that takes the value one if the bank was subject to intervention in 1998, and zero otherwise. Interventions include closure, recapitalization, nationalization, sale to foreigners, and domestic takeovers. Both models are estimated using OLS. Country dummy variables were added but not reported. Heteroskedasticity-consistent standard errors are in parentheses.
Source: Author's calculations.

did not differ between 1998 and 1997. Although not reported in the table, the country effects for the banks subject to intervention are insignificant.

Banks with a high estimated cost of deposit insurance are expected to have a higher chance of failing, because they are thought to take higher risks. To assess the power of deposit insurance cost estimates to forecast bank failure, I estimate a model with a dummy variable taking the value one if the bank is subject to intervention in 1998 as the dependent variable and the cost of deposit insurance in 1996 or 1997 as the independent variable. In 1998 some East Asian governments intervened more heavily in their banking sectors than did others. In Thailand, for example, the government intervened in all banks, whereas the Malaysian government allowed banks to continue operating even though many were undercapitalized. To control for differences in the level of intervention, I add country dummy variables to the model. None is included for Thailand (because all Thai banks in the sample were subject to intervention) or for economies in which no banks were subject to intervention in 1998. I estimate the model using OLS and, because of the discrete nature of the dependent variable, also estimate both a probit and a logit model.

The results show that banks with a high cost of deposit insurance in 1996 or 1997 had a higher chance of failing or being subject to intervention in 1998 than did banks with a low cost of deposit insurance (table 11). In addition to Thai-

TABLE 11. Predicting Bank Distress

	(a)		(b)		(c)	
Variable	*RV*	*Duan*	*RV*	*Duan*	*RV*	*Duan*
Panel A						
Constant	0.089***	0.098***	-1.367***	-1.869***	-2.255	-3.229***
	(0.028)	(0.032)	(0.170)	(0.120)	(0.313)	(0.436)
Insurance	0.002**	0.0011***	0.005**	0.018***	0.009*	0.032***
	(0.001)	(0.0004)	(0.002)	(0.003)	(0.005)	(0.007)
Indonesia	0.551***	0.404	1.686***	-0.362	2.759***	-0.078
	(0.167)	(0.179)	(0.523)	(0.948)	(0.912)	(1.953)
Japan	-0.022	-0.037	-0.030	-0.341	-0.129	-0.625
	(0.078)	(0.077)	(0.421)	(0.458)	(0.811)	(0.893)
Korea, Rep. of	0.479***	0.488***	1.485***	1.552***	2.438***	2.692***
	(0.116)	(0.110)	(0.329)	(0.369)	(0.572)	(0.691)
Malaysia	0.113	0.041	0.523	-0.614	0.887	-1.102
	(0.139)	(0.131)	(0.484)	(0.489)	(0.832)	(0.879)
(Pseudo)-R^2	0.288	0.333	0.266	0.463	0.259	0.457
Observations	143	143	143	143	143	143
Panel B						
Constant	0.131***	0.117***	-1.153***	-1.226***	-1.900***	-2.052***
	(0.036)	(0.036	(0.161)	(0.176)	(0.298)	(0.326)
Insurance	0.039*	0.005***	0.174	0.024	0.298	0.042
	(0.021)	(0.001)	(0.111)	(0.016)	(0.249)	(0.040)
Indonesia	0.347	0.622***	0.867	1.853***	1.530	3.060***
	(0.252)	(0.158)	(0.736)	(0.505)	(1.306)	(0.862)
Japan	-0.027	-0.013	-0.103	-0.032	-0.247	-0.099
	(0.080)	(0.080)	(0.419)	(0.424)	(0.800)	(0.815)
Korea, Rep. of	0.368***	0.414***	1.065***	1.284***	1.753***	2.142***
	(0.134)	(0.122)	(0.371)	(0.345)	(0.642)	(0.576)
Malaysia	0.087	0.105	0.370	0.461	0.612	0.799
	(0.147)	(0.146)	(0.494)	(0.497)	(0.855)	(0.866)
(Pseudo)-R^2	0.199	0.233	0.194	0.222	0.190	0.219
Observations	138	138	138	138	138	138

*Significant at the 10 percent level.
**Significant at the 5 percent level.
***Significant at the 1 percent level.

Note: The dependent variable is a dummy variable that takes the value one if the bank was subject to intervention in 1998, and zero otherwise. Interventions include closure, recapitalization, nationalization, sale to foreigners, and domestic takeovers. *Insurance* is the cost of deposit insurance in basis points of total debt calculated using either the RV method or the Duan method. Country dummy variables have been added for Indonesia, Japan, Korea, and Malaysia. A probit model is estimated. The models in panel A use deposit insurance cost data for 1997 and intervention data for 1998; the models in panel B use deposit insurance cost data for 1996 and intervention data for 1998. Model (a) uses the OLS estimation method. Model (b) estimates a probit model. Model (c) estimates a logit model.

Source: Author's calculations.

land, Indonesia, and Korea also intervened more heavily in their banks than did other countries. These results support my claim that the cost of deposit insurance has some power in predicting the riskiness of banks and forecasting bank distress. The results based on the Duan method show that a one-basis-point increase in the cost of deposit insurance in 1997 raises the likelihood of intervention in 1998 by roughly 1.8 percent according to the probit model and by 3.2 percent according to the logit model.

VI. Conclusions

Arguing that a relatively high cost of deposit insurance indicates that a bank takes excessive risks, I use the cost of deposit insurance to assess the relationship between the risk taking behavior of banks and their governance structure. To do so, I estimate the cost of deposit insurance for a large number of banks in different economies, using RV and Duan techniques. The results show that the cost is highest for banks with concentrated private ownership, especially those predominantly owned by a single company or another financial institution, and, to a lesser extent, for state- or family-owned banks—indicating that these banks tend to take the greatest risks. In contrast, banks with dispersed ownership engage in a relatively low level of risk taking. The cost of deposit insurance also tends to be higher for banks that are affiliated with a business group, are small, have high credit growth, and are located in countries with low GDP per capita, high inflation, poor quality and enforcement of laws, or low penetration by foreign banks. Finally, I find that as a proxy for bank risk, the cost of deposit insurance has some power in predicting bank failures.

The findings support the view that existing government deposit insurance schemes create moral hazard for banks. They also suggest that these incentive problems differ in magnitude between different types of banks—in particular, between banks with different governance structures—and between different types of institutional environments. Banks characterized by concentrated private ownership and operating in an environment with weak institutions tend to take high risks. The ultimate goal should be a financial system in which banks have dispersed private ownership and both shareholders and depositors are protected by proper enforcement of prudent regulation.

Both the findings of the article and the method it proposes for measuring bank risk have importance for policymakers. First, the findings support the view of many policymakers that one of the keys to a sound financial system is dispersed private ownership of banks. Second, the findings indicate that dispersed private ownership of banks is even more important for the stability of financial systems where corporate governance systems, and institutional environments in general, are weak, as in many developing countries. Finally, the article shows that as a proxy for bank risk, the cost of deposit insurance could be a useful additional tool for identifying troubled financial institutions and, at an aggregate level, for providing early warning of banking crises.

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How Different Is the Efficiency of Public and Private Water Companies in Asia?

Antonio Estache and Martín A. Rossi

Several studies have compared the efficiency of publicly and privately owned water utilities and reached conflicting conclusions on the impact of ownership on efficiency. This article provides further evidence by estimating a stochastic cost frontier for a sample of Asian and Pacific regional water companies. The results show that efficiency is not significantly different in private companies than in public ones.

Policymakers in developing countries, eager to resolve the decade-long debate on the gains from privatization of water utilities, are increasingly interested in assessments of the efficiency of public and private water utilities. Most early studies focused on the performance of public and private providers in the United States. Crain and Zardkoohi (1978), estimating a cost function derived from a generalised Cobb-Douglas production function with a dummy variable for ownership, found that publicly owned water utilities in the United States had higher costs than their privately owned counterparts. Feigenbaum and Teeples (1984) used a translog approximation and concluded that they could not reject the hypothesis (at the 5 percent significance level) that the parameters were identical for government and private operation. Byrnes, Grosskopf, and Hayes (1986) measured efficiency directly in terms of a production function and found no evidence that publicly owned utilities are more wasteful or operated with more slack than privately owned utilities. Fox and Hofler (1986) estimated the extent and cost of technical and allocative inefficiency and found no statistical difference in inefficiency for public and private firms, although they did find allocative differences. Overall, these studies leave the impression that there is no convincing evidence of a systematic superiority of one form of ownership over another.

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This article provides further evidence for the debate by estimating a stochastic cost frontier using 1995 data from a sample of 50 water companies in 29 Asian and Pacific region countries. Because of conflicting empirical evidence, justification of models and robustness of results are key issues. For that reason, this study applies two approaches to measuring efficiency—error components (EC) and technical efficiency effects (TEE) models—and runs tests for different specifications with those two approaches. The analysis confirms the U.S. results and suggests that efficiency is not significantly different in private and public companies.

I. The Theoretical Cost Function

Frontiers are typically classified as production or cost functions, with the nature of the sector determining which function to estimate. Most water utility firms are required to provide services at a preset tariff. In simpler terms, they are required to meet demand and may not chose the level of output to supply. Because output is exogenous, the firm maximizes benefits by minimizing the cost of producing a given level of output. Specification of a cost frontier is thus often the natural choice.[1]

The theoretical specification of the cost function is

$$C = f(\mathbf{Y},\mathbf{Z},\mathbf{P})\exp(\varepsilon), \tag{1}$$

where C is total cost, **Y** is the output vector,[2] **Z** is a vector that includes all the relevant exogenous variables needed to allow comparisons across firms, **P** is a vector of input prices, and ε is the error term.

The systematic part of the model is the cost frontier, which determines the minimum cost achievable for a given set of outputs, input prices, and control variables. The error term can be decomposed in two parts:

$$\varepsilon_i = u_i + v_i, \tag{2}$$

where $u_i \geq 0$ and v_i is not constrained. The v_i component captures the effects (for firm i) of the stochastic noise and is assumed to be independent and identically distributed following a normal distribution $N(0,\sigma^2_v)$. This component accounts for measurement error and other random factors, such as effects of weather and strikes, as well as misspecifications in the estimated cost function. The u_i component represents the cost inefficiency and is assumed to be distributed independently from v_i and the regressors. Various distributions have been suggested for this term: half-normal (Aigner, Lovell, and Schmidt 1977), gamma (Greene 1990), and exponential (Meeusen and van de Broeck 1977). The half-normal

1. Nevertheless, utilities frequently do limit the number of customers through the use of two-part tariffs or rationing.

2. Another advantage of the cost frontier over the production frontier is that it deals better with multiple outputs.

distribution is the most commonly used in empirical studies and implies that the majority of the firms are almost efficient. To avoid imposing such an a priori distribution of the inefficiency term, the more flexible truncated normal was adopted (Stevenson 1980), a generalization of the half-normal obtained by truncating to zero a normal distribution with median μ and variance σ^2_u. Setting μ to zero reduces to the traditional half-normal model. Therefore, the null hypothesis Ho: $\mu = 0$ will be tested.

When the error term enters multiplicatively in the cost function (additively after logs where taken), the level of the cost efficiency or overall economic efficiency of the ith firm is

$$EF_i = \exp(-u_i). \tag{3}$$

The problem is that the u_i term is unobservable. Battese and Coelli (1988) show that the best predictor of $\exp(-u_i)$ is obtained by using the conditional expectation

$$E[\exp(-u_i)|\varepsilon_i] = \{[1-\Phi(\sigma_A-\gamma\varepsilon_i/\sigma_A)]/[1-\Phi(-\gamma\varepsilon_i/\sigma_A)]\}^*\exp(-\gamma\varepsilon_i + \sigma^2_A/2), \tag{4}$$

where $\Phi(.)$ is the distribution function of the standard normal random variable. Following the parameterization proposed by Battese and Corra (1977), σ_v^2 and σ_u^2 are replaced with $\sigma^2 = \sigma_v^2 + \sigma_u^2$, $\gamma = \sigma_u^2/(\sigma_v^2 + \sigma_u^2)$, and $\sigma_A = [\gamma(1-\gamma)\sigma^2]^{1/2}$. The parameter γ must lie between 0 and 1, with 0 indicating that the deviations from the frontier are due entirely to noise, and 1 indicating that all deviations are due to inefficiency. This specification allows testing the null hypothesis that there are no inefficiency effects in the model, H_0: $\gamma = 0$, against the alternative hypothesis, H_1: $\gamma > 0$.

FRONTIER version 4.1 (Coelli 1996) is used to obtain the maximum likelihood (ML) estimates of the parameters of this model and the efficiency measures.

II. Data and Estimation

The cost frontier for the Asian water utilities was estimated from a database published by the Asian Development Bank (McIntosh and Yñiguez 1997). The sample covers 50 firms surveyed in 1995 in 19 countries: Bangladesh (2 firms), Bhutan (1), Cambodia (1), China (5, including Hong Kong and Taiwan), Cook Islands (1), Fiji (1), India (4), Indonesia (3), Kazakhstan (1), Republic of Korea (2), Kyrgyz Republic (1), Lao People's Democratic Republic (1), Malaysia (3), Maldives (1), Mongolia (1), Myanmar (2), Nepal (1), Pakistan (3), Papua New Guinea (1), the Philippines (3), Singapore (1), the Solomon Islands (1), Sri Lanka (1), Thailand (3), Tonga (1), Uzbekistan (1), Vanuatu (1), Vietnam (2), and Samoa (1).[3]

The data have the advantage of providing comparable information for all the sample companies, but they have limitations. They cover just one year, and they

3. The 50 Asian water companies were selected jointly by representatives of utilities and the Asian Development Bank (ADB). ADB recruited domestic consultants to assist firms in responding to the questionnaire.

are too sparse to support complete analysis of each company. Because they contain no information on the asset base, it is difficult to assess whether operational costs are consistent with the maintenance requirements of the assets. This is an important issue for regulated firms subject to price or revenue caps because their chief cost-cutting options for meeting caps or other restrictions are to reduce the quality of service or to cut back on maintenance. The relevance of the unavailability of the asset data for the comparison of public and private provision is uncertain. Even so, using the cost data to compare the performance of firms in the sample allows for testing of the robustness of earlier results of studies comparing public and private firms in developing countries.

The data cover operational costs (*COST*, in thousands of U.S. dollars),[4] annual salary (*SALAR*, estimated as the ratio of total salary cost to the number of workers), number of clients (*CLIEN*, in thousands), daily production (*PROD*, in cubic meters per day), number of connections (*CONE*), population density in the area served (*DENS*, in people per square kilometer), percentage of water from surface sources (*ASUP*), number of hours of water availability per day (*QUALI*, in h), percentage of metered connections (*METER*), and a set of qualitative variables on the type of treatment used: chlorination (*DUMCLO*, with a value of 1 when the treatment is chlorination and 0 otherwise) and desalination (*DUMDES*; in fact, only one company uses desalinization).

The estimated function is in line with practice in previous studies (see Stewart 1993, Crampes, Diette, and Estache 1997, Price 1993, Byrnes, Grosskopf, and Hayes, 1986, Fox and Hofler 1986, and Bhattacharyya, Harris, and Rangesan 1995). However, because the only input price available was for labor, an ad hoc cost function was estimated.[5] The dependent variable is operational costs, which include expenditures for personnel, power, parts, materials, and bulk purchase of water in some cases.[6] Included as the main cost drivers are average salary (proxy of the main input price), number of clients, daily production and number of connections (proxies of outputs), population density, percentage of water from surface sources, percentage of metered connections, quality, and two dummy variables that account for differences in the type of treatment used (environmental variables).

One advantage of this methodology is that it allows for the inclusion of environmental variables in the model specification—variables that may affect the performance of the firm but are not entirely under its control. Their inclusion ensures that the various operators of an activity are effectively comparable. Population

4. To make data comparable, gross *COST* data in local currency were converted into U.S. dollars at the rate of exchange as of July 1, 1997, using market rates from the New York Foreign Exchange, rates from the country's central bank, or book rates provided by the International Monetary Fund.

5. Estimation of a cost function requires data on input prices, including capital. However, this information is difficult to obtain (see, for example, Huettner and Landon 1977, for the electricity distribution sector). The usual solution is to formulate an arbitrary cost function, without including the price of the capital input.

6. The average share of labor in operational cost is 35 percent, with a standard error of 20 percent.

density, for instance, plays an important role in defining the network infrastructure, especially in regulated firms that are obliged to serve a specific geographical area. The percentage of water from surface sources is included as a control variable because the costs associated with drawing water very much depend on the water input source. The percentage of metered connections is included as a regressor because the administrative cost is higher than it is for the flat-rate system (Bhattacharyya, Harris, and Rangesan 1995). Hours of water availability are included because that can affect costs even after controlling for outputs.

Twenty-two of the 50 utilities have some form of private sector participation. Major private sector management (concession) is under way in the Philippines, Vanuatu, Maldives, and the Solomon Islands. Other types of private sector participation include billing and collection, leak repair, meter reading, source development, production, and pumping (McIntosh and Yñiguez 1997). Three dummy variables are included to account for this heterogeneity: a dummy concession (*DUMCON*, with a value of 1 if the firm is a concession and 0 otherwise), a dummy administration (*DUMBC*, with a value of 1 if the private sector is involved in billing, collection, leak repair, or meter reading and 0 otherwise), and a dummy for other private sector participation (*DUMOP*). The basis for comparison will be public sector performance. The basic statistics are summarized in table 1.

Because the quality of the estimates of the frontier and efficiency measures depend on the accuracy of the specification of the functional form, the Cobb-Douglas specification was tested. A translog cost function, a more flexible form, was not estimated because the inclusion of the second-order and cross terms would leave the model with very few degrees of freedom. To account for variable returns to scale, the models were run with quadratic terms in output alone and in labor price and output variables. In neither case were the results for the included variables statistically significant. A likelihood ratio test was performed, and the null of the Cobb-Douglas specification could not be rejected.[7] Therefore, a Cobb-Douglas cost function was estimated. The initial model is as follows.

$$\ln COST = \alpha + \beta \ln SALAR + \omega_1 \ln CLIEN + \omega_2 \ln CONE + \omega_3 \ln PROD + \pi_1 \ln DENS + \pi_2\, ASUP + \pi_3\, QUALI + \pi_4\, METER + \pi_5\, DUMDES + \pi_6\, DUMCLO + \pi_7\, DUMCON + \pi_8\, DUMBC + \pi_9\, DUMOP \tag{5}$$

The estimated value of μ in the EC model was 0.09, with a standard error of 1.14. A likelihood ratio test was performed, and since the null hypothesis ($\mu = 0$) could not be rejected, the estimation assumed a half-normal distribution.

For the ordinary least squares (OLS), corrected ordinary least squares,[8] and ML estimates of the EC model, the signs of the coefficients are as expected (table 2). The labor input has a positive and significant sign, as do connections and clients. The other product (daily production) has the expected positive sign but is not

7. A RESET test (second power) showed no evidence of omitted variables in the model.

8. OLS plus a change in the intercept.

TABLE 1. Values of Key Variables for 50 Sample Firms

Variable	Average	SD
COST (thousands of US$)	29,372	67,721
SALAR (dollars)	5,042	8,619
CLIEN (thousands)	2,453	2,945
PROD (m^3/day)	935	1,254
CONE (number)	416	548
DENS (people per km^2)	16,587	33,479
ASUP (%)	0.67	0.41
METER (%)	0.74	0.39
QUALI (hours of water availability per day)	18.98	6.85

Source: McIntosh and Yñiguez (1997).

significant.[9] An improvement in quality increases costs, as does an increase in the proportion of metered clients. Population density has a negative and significant sign, suggesting that it is cheaper to serve more densely populated areas. The proportion of water from the surface is not significant. The dummy variable for desalinization is positive but not significant. The signs on the concession dummy variable and the other private sector participation dummy variable are positive but not significant. Finally, the sign of the dummy administration is negative but not significant.

In the extreme case, where $\sigma_u^2 = 0$ (the ratio of the variance of noise to the total residual variance is equal to 1), the ML and OLS estimates are the same, because the composed errors follow a normal distribution. The ML and OLS estimates in table 2 are quite close, which can be explained by the low value of γ (which is not statistically significantly different from zero) or, analogously, by the high (0.60) ratio of the variance of noise to the total residual variance.[10] These results seem to suggest that OLS is the appropriate model (i.e., there is no need to estimate a frontier as all departures from the cost function are due to noise); with no frontier necessary, all observations can be considered equally efficient, which supports the hypothesis of no differences in cost efficiencies between public and private operators.

To determine the robustness of the results, a second model was estimated in which the inefficiency effects are expressed as a function of the ownership dummy variables. This TEE model, as proposed by Battese and Coelli (1995), is similar to the EC model except that the efficiency error has a mean of m_i instead of 0, where $m_i = \delta x_i$ is a contemporaneous auxiliary regression such that $\mathbf{x}_i$ is a $p \times 1$ vector of variables that may influence the efficiency of the firm and δ is a $1 \times p$ vector of parameters to be estimated simultaneously with the parameters α, β, ω, and π.[11]

9. When the model was run without CONE, the main conclusions were unaffected, but the t-value of PROD increased.

10. Estimated as $\sigma_v^2/(\sigma_v^2 + \sigma_u^2\,\pi/[\pi - 2])$ or $1-(\gamma/[\gamma + (1-\gamma)\pi/(\pi - 2)])$.

11. If x_i contains the value 1 and no other variable, then the model reduces to the truncated normal proposed by Stevenson (1980) and shown here.

TABLE 2. Results for the Error Components Model

Variable	OLS	Corrected OLS	ML
Constant	0.495 (0.53)	0.280	0.139 (0.16)
ln(*SALAR*)	0.293 (6.06)	0.293	0.297 (6.97)
ln(*CLIEN*)	0.671 (3.63)	0.671	0.700 (3.82)
ln(*CONE*)	0.269 (3.95)	0.269	0.285 (4.13)
ln(*PROD*)	0.080 (0.45)	0.080	0.044 (0.25)
ln(*DENS*)	−0.139 (−1.65)	−0.139	−0.148 (−1.88)
ASUP	0.116 (0.46)	0.116	0.106 (0.49)
QUALI	0.029 (1.99)	0.029	0.029 (2.32)
METER	0.320 (1.51)	0.320	0.293 (1.51)
DUMDES	0.577 (0.81)	0.577	0.539 (0.88)
DUMCLO	0.213 (1.01)	0.213	0.195 (1.08)
DUMCON	0.002 (0.008)	0.002	0.044 (0.17)
DUMBC	−0.092 (−0.49)	−0.092	−0.067 (−0.40)
DUMOP	0.195 (1.01)	0.195	0.196 (1.18)
γ	0.420	0.65	(1.15)
Log-likelihood	−19.42		−19.34

Note: The dependent variable is the log of operational cost (ln*COST*). The numbers in parentheses are *t*-statistics.

Source: Authors' calculations based on data from McIntosh and Yñiguez (1997).

The results are similar to those for the EC model.[12] Salary, percentage of metered clients, and hours of water availability all have a positive and significant effect on costs (table 3). As in the EC model, population density has a negative and significant sign and percentage of water from surface sources has a positive but not significant effect. The only difference between the two specifications is on the private-public question, because the concession dummy variable has a negative sign, although, as in the EC model, it is not significant.

Average efficiency is 1.39 in the EC model and 1.44 in the TEE model. The ML estimates (both the EC and TEE models) suggest that the differences between private and public operators are not significant, and similar results arise from the OLS estimates.[13]

III. WHERE DO WE GO FROM HERE?

The results discussed here confirm the very cloudy impression emerging from the U.S. experience and do not provide strong evidence that private providers are globally more efficient than public operators. However, the results highlight

12. With a TEE model including a constant term in the inefficiency term, the main result relating to the public-private issues was unaffected.

13. The TEE and EC models differ in that the EC model allows for different intercepts for the different ownership categories whereas the TEE model assumes the same intercept. Hence the cost efficiency scores from the TEE model are gross because they include the ownership effect while the scores from the EC model are net of this effect (see Coelli, Perelman, and Trujillo 1999 for more on net and gross efficiency).

TABLE 3. Results for the Technical Efficiency Effects Model

Variable	OLS	Corrected OLS	ML
Constant	0.609 (0.71)	0.330	−0.113 (−0.14)
ln(*SALAR*)	0.294 (6.72)	0.294	0.303 (6.94)
ln(*CLIEN*)	0.708 (3.97)	0.708	0.668 (3.30)
Ln(*CONE*)	0.269 (4.48)	0.269	0.305 (4.59)
ln(*PROD*)	0.050 (0.29)	0.050	0.054 (0.30)
ln(*DENS*)	−0.161 (−2.05)	−0.161	−0.127 (−1.53)
ASUP	0.105 (0.43)	0.105	0.150 (0.74)
QUALI	0.031 (2.26)	0.031	0.029 (2.24)
METER	0.255 (1.32)	0.255	0.372 (1.66)
DUMDES	0.537 (0.88)	0.537	0.632 (0.11)
DUMCLO	0.171 (0.87)	0.171	0.238 (1.19)
δ_2 (*DUMCON*)			−0.290 (−0.24)
δ_3 (*DUMBC*)			−0.955 (−0.74)
δ_3 (*DUMOP*)			0.309 (0.98)
γ		0.580	0.752 (2.28)
Log-likelihood	−20.63		−18.80

Note: The dependent variable is the log of operational cost (ln*COST*). The numbers in parentheses are *t*-statistics.

Source: Authors' calculations based on data from McIntosh and Yñiguez (1997).

the difficulty of measuring efficiency, reflecting a long tradition of lack of concern for efficiency among regulators in developing countries. This is changing, however. One of the main regulatory adjustments over the last decade has been the recognition that efficiency does matter, a feeling that is spreading as privatization takes hold around the world. Many regulators have switched from rate-of-return regulation to price or revenue-cap regulation to increase the incentive for firms to minimize costs and to ensure that consumers eventually benefit from these cost reductions.

This means that costs need to be measured much more precisely than they were for the ADB database (McIntosh and Yñiguez 1997). Indeed, if any cost reductions are expected to result from private operation of the sector, they should be associated with efficiency gains rather than quality reductions. Both have to be measured if cost differences—or the lack thereof—across firms are to be explained correctly.[14] This alone explains why efficiency measures are no longer a side show as they were under rate-of-return regulation. The data here do not allow for testing of tradeoffs between efficiency gains and quality reductions.

A related regulatory challenge is how to document the fact that a firm's efficiency gains can come from two different sources. Gains can come from shifts in the frontiers reflecting efficiency gains at the sectoral level. But efficiency gains at the firm level can also reflect a *catching-up effect*. These are the gains

14. See Coelli and others (2001) for a longer discussion.

to be made by a firm not yet on the frontier. Public firms that have to compete with new private entrants who enjoy the latest technology will often be expected to play catch-up or die. These firms should be able to achieve not only the industry gain but also specific gains to offset firm-specific inefficiencies. This catch-up effect is one of the expected benefits to consumers of yardstick competition if regulators can ensure that quality is not the adjustment variable for the least cost efficient firms. Yardstick competition—even implicit, as a consequence of studies of this kind that generate results forcing comparisons—should minimize the scope for major differences between public and private providers. In the end, the inconclusiveness of the comparison of efficiency in public and private water utilities may simply reflect the fact that competition matters more than ownership.

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