The Developing Economies, XXIII-1 (March 1985)

DETERMINANTS OF IRON AND STEEL EXPORTS

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I. INTRODUCTION

THIS paper, covering the time period from 1966 to 1978, analyzes the export behavior of thirteen countries[1] in the area of iron and steel products. These countries together account for over 90 per cent of the world's iron and steel exports. This study uses annual data on exports at three-digit level of disaggregation.[2] We consider eight product classes of iron and steel [SITC (Standard International Trade Classification)[3] product code 671 through 679, except 675[4]]. By pooling cross-sectional data relating to the product classes mentioned above for each country for thirteen years, we estimate an export function which (besides including the relative prices and income as explanatory variables) uses time and product dummies, allowing the intercept to vary across product classes and over time.[5] This enables us to separate the influence of characteristics associated with commodity and time on exports of the iron and steel industry. Thus, our study in its comparative analysis of thirteen major exporting countries attempts to relate the movements in export flows to factors like variation in import preference and changing supply conditions, in addition to the estimation of the influence of price and income factors. An inter-country analysis of the kind attempted here should prove useful. The usefulness of the present study is further enhanced because of two distinctive features: (a) this study is more disaggregative than existing

[1] The countries included in this study are: Belgium and Luxembourg, Brazil, Canada, Federal Republic of Germany, France, India, Italy, Japan, the Netherlands, Republic of Korea, Singapore, the United Kingdom, and the United States.

[2] Data for this study have been drawn from various UN sources, the main sources being *Commodity Trade Statistics* and the *Statistical Yearbook*.

[3] Revised code. In terms of this classification iron and steel products are included in SITC's section 67.

[4] Product code 675 has been excluded since complete data for this three-digit product category could not be obtained for most of the thirteen countries covered here. Again, for some of the countries, we could not get complete data even for the other eight product classes. Thus, we consider five product classes for Japan (671, 672, 676, 678, 679) and four product classes for Brazil (671, 672, 673, 674), Korea (671, 674, 678, 679), and Singapore (672, 673, 674, 679).

[5] The use of dummies relating to time and product variables, by allowing the intercept to vary across product classes and over time, in fact captures the effects of omitted variables, which do not lend themselves easily to direct observation. This technique, known as covariance analysis, was applied for the first time for the analysis of export performance by Ginsburg and Stern [7] and later by Ginsburg [6], Richardson [16] [17], Brada and Wipf [4], and others. For some theoretical details on this technique, see Suits [21].

studies on export demand function which have dealt chiefly with aggregate exports, and (b) this study includes in its purview some developing nations and thus adds to the scant existing empirical evidence available at present for developing nations.

This paper is organized as follows. Section II discusses some methodological issues in the estimation of export function and the limitations of the present study. In Section III, we present the basic model. Empirical results are reported in Section IV. Finally, the main findings of this study are summarized in the concluding section.

II. METHODOLOGICAL ISSUES

Data problems confronted in the estimation of export demand functions have been thwarting attempts at more disaggregative levels and have restricted empirical researchers in adopting an approach much different from that which theoretical considerations require. The non-availability of and gaps in data for the years before 1966 pertaining to product groups at three-digit level of disaggregation for developing countries have constrained us to choose a short time span of thirteen years. Further, this problem coupled with the problem of nonconformity between SITC and ISIC (International Standard Industrial Classification) at the three-digit level has been decisive in our choice of the export function employed and the estimation method adopted in the present study.

In our present formulation, the basic model used is a single equation export function, and as such, it may be subject to the usual criticism that it ignores the simultaneity between the level of exports and the relative export price,[6] hence the OLS (ordinary least squares) estimates of the relative price coefficient may be biased. In our defense, two points may be emphasized. First, the experiments using simultaneous equations technique in export function studies either did not meet with encouraging success, or yielded results similar to those obtained by using OLS.[7] In the estimates of price elasticities from the use of simultaneous equations technique over the use of OLS technique in the studies undertaken at disaggregated level, there has been little or no improvement. Also considering the data problems mentioned earlier, we felt that the difficulties encountered in constructing and estimating a simultaneous equations model, in place of the present single equation, for each of the countries considered here may far outweigh the expected gains in terms of the improvement in the parameter estimates

6 The severity of this attack is lessened if one uses a model yielding estimates of price elasticity of substitution between the exports of two countries, because such a formulation employs the ratio of two export functions which tends to cancel out the simultaneity bias. See Leamer and Stern [13] for details on this point. We have, however, not used the elasticity of substitution model here.

7 See, for example, studies by Morgan and Corlett [15] and Artus and Sosa [1]. In a disaggregated study at three-digit level for the United States, EEC, and Japan, Stone [20] presents estimates of price elasticity of exports using both OLS and 2SLS (two-stage least squares) methods and finds OLS to be the "better" estimator for 55 per cent of the equations and 2SLS for the remaining 45 per cent.

from simultaneous equations techniques. Second, for developed countries there is growing evidence in recent years[8] that, at least for manufactures, the export market is oligopolistic in which price is fixed prior to quantity. The evidence, indeed, casts doubts on the usefulness of determining prices and quantities simultaneously, and recommends instead postulating a recursive process in which price is fixed prior to quantity and the resulting demand is either supplied from production and stock or discouraged by altering non-price competitiveness [22, p. 8]. One may argue here that for developing countries it may not be reasonable to assume that their export market of manufactures is characterized by oligopolistic conditions. For the present study which covers both developed and developing countries, an ideal procedure would then be to test different specifications of export function for each individual country and employ, among other variables, the country specific factors, too. However, in a study that is concerned with the export behavior of thirteen countries, such a procedure would entail an enormous task. A common analytical framework is, therefore, adopted with a view to undertake broad comparisons of the factors influencing the export performance in individual countries.

In a world in which information is less than perfect and adjustments neither costless nor immediate, it may be argued that a legitimate export function should, in its specification, take into cognizance the existence of lags. Junz and Rhomberg [10] in their widely acknowledged study found that the response of trade flows to relative price changes stretched out over a period of four to five years. In their words, "Almost 50 percent of the full effect appears to work through during the first three years, and almost 90 percent during the first five years, following a price change" [10, p. 418]. Given a strong case for the possibility of lagged reactions of export flows to changes in relative prices, we experimented with two lag patterns. In the first case, the coefficients of the four lagged price variables used in our study were left unconstrained; in the second exercise, we imposed an inverted V-type lag pattern, using numerical weights selected on a pattern of elasticity estimates obtained by Junz and Rhomberg for prices lagged respectively be zero, one, two, and three years.[9] In the former case, only the coefficient of the current relative price term was statistically significant; in the latter case the coefficients of the price terms were often insignificant, and in some cases wrongly signed. It seems to us that our failure to find lagged response of export flows to price changes is attributable, at least in part, to the following two reasons. First, the relative prices have been measured in terms of unit values and as a result they represent delivery rather than contract prices, and in delivery prices the response to a relative price change should be more immediate.[10] Second,

[8] See Winters [22] for details.

[9] Junz and Rhomberg found that from a concurrent value elasticity of -0.5, the response declines at first, rises again to a peak of -1.0 in the third year before falling off for the longer lags. See [10, Table 1-B].

[10] Since unit values refer to deliveries, they already allow for gestation and delivery lags. See Winters [22, p. 20] on this point. Also see Stone [20] and Junz and Rhomberg [10] for similar comments.

goods at three-digit level of disaggregation and belonging to the same commodity group, as in our present study, are fairly homogeneous across suppliers and in consequence, the adjustment of buyers and sellers can be expected to be quite quick and without much lag.

As most earlier studies have done, this study uses unit value index as a proxy for price index, and as such it suffers from the usual defects of this measure. However, since the unit values used in our study are at three-digit level, which are actually weighted averages of the four-digit, the weights being four-digit quantities;[11] they are subject to less aggregation bias. In constructing a world unit value index facing each *i*th country for each *j*th product included in the SITC-67 group, we computed the weighted average of unit value indices of all non-*i*th countries in our sample for the *j*th category. The weights used are their relative export shares of that category in 1975.

Though the approach adopted here is akin to that adopted by Ginsburg [6], Ginsburg and Stern [7], and Brada and Wipf [4], we have made the following two modifications. First, we have used dummy variables for commodity and time characteristics, whereas the studies cited above include dummy variables to account for regional characteristics as well. Also, while in our study we use only intercept dummies (as in Richardson [17]), in these studies both slope and intercept dummies have been used. Second, the studies cited above use the export-market-share model in which the ratio of country A's exports to the sum of exports of countries A and B is taken as the dependent variable and regressed on the relative export price. This formulation extended to the present case would imply that we should take as the dependent variable the ratio of *i*th country's exports to total exports of all thirteen countries and the relative price variable should be the ratio of *i*th country's unit value index to the weighted average of unit value indices of its competitors. However, this formulation, when estimated, showed a poor fit for many countries presumably because in this specification the implicit restriction imposed on the coefficient of "total world exports" (i.e., exports of all thirteen countries taken together) to be equal to unity was not warranted. Evidently, this formulation implies that with other factors remaining the same an expansion in the world demand of iron and steel products of a particular three-digit category would be shared by the thirteen countries considered here in proportion to their exports in that category. In the subsequent formulation used later in this paper, the variable "total world exports" was allowed to be unconstrained and entered as a separate explanatory variable (to serve as a proxy for the activity variable in place of world income, sometimes used) in addition to the relative export price variable.

Of late, much attention has been paid in econometric literature on export function to the influence of factors like domestic demand pressure and product quality (including design and delivery time).[12] The influence of such factors enters our model through the dummy variables. Using the coefficients of time dummies

[11] See Shinkai [19].

[12] See, for example, Ball, Eaton, and Steuer [3], Sato [18], Balassa [2], and Dunlevy [5].

we have examined whether the observed pattern in the estimated coefficients of time dummies matches with the variation in domestic demand pressure and the growth of iron and steel industry.

III. THE MODEL

For each country, pooling cross-section data for three-digit commodity classes in SITC-67 for thirteen years (1966–78) we have estimated the following regression equation using restricted least-squares method:

$$\log X_{it} = a + (m_i + n_t) + b \log W_{it} + c \log (P_{it}/Q_{it}) + u_{it},$$

$$\sum m_i = \sum n_t = 0,$$

where

i, t = subscripts for product class and year, respectively,
m_i, n_t = coefficients of product dummies and year dummies, respectively,
X_{it} = quantity of exports of the ith product in year t,
W_{it} = quantity of world exports of the ith product in year t,
P_{it} = export price of the ith product in year t,
Q_{it} = weighted average of competitors' export prices of the ith product in year t,
u_{it} = randomly distributed error term.

It is seen from the above that the export function has been specified in the traditional multiplicative form and estimated in a transformed log-linear form. The two price variables enter in a ratio form as a relative price variable. This implies that a 10 per cent fall in the export price of a country has the same effect on its export performance as a 10 per cent increase in the export prices of its competitors. This is based on the assumption of homogeneity, an assumption which most earlier studies have made.[13] Instead of using a weighted average of national incomes of the importing countries as the income variable, we have used total world exports. This practice has found favor in several earlier studies. It has been pointed out that the use of world trade as the income or activity variable tends to correct for some of the variations in institutional barriers to trade and obviates the problem of creating and interpreting an aggregation of national income [5, p. 132].

The dummy variables allow the intercept term to vary from year to year and from product class to product class, picking up thereby the influence of factors other than income and relative prices. It is assumed on the other hand that the income elasticity b and the price elasticity c are stable not only across years (as in conventional time-series analysis), but also across product classes (as in conventional cross-section analysis and implicitly in any analysis—time-series or cross-section—using aggregate data). This assumption does not seem unreasonable

[13] This assumption is subject to criticism on both theoretical and empirical grounds. See Magee [14] and Stone [20].

since this study deals with iron and steel products only and there is in consequence a built-in similarity in the observations.[14]

Data for this study have been taken from various UN sources, the main source being *Commodity Trade Statistics*. The reported figures on quantity exported (measured in thousand metric tons) have been used for X_{it}. A measure of total world exports W_{it} is formed by summing X_{it} for the thirteen countries covered in this study. The price variables are based on unit values (f.o.b. in U.S. dollars). For each country j and for each product i a price index for its competitors is formed by taking a weighted average of their export prices; the weights being their export shares (in value) in that product category in 1975.

The limitations of the use of unit values in place of export prices are well known. These have been thoroughly investigated by Kravis and Lipsey [12] and need not be discussed here. Suffice it to say that, apart from other defects, unit value index at an aggregate level may give undue weight to an item with relatively low elasticities and also conceal some of the substitution that may take place within a class. Considering, however, the nature of the product being studied and the level of disaggregation it seems to us that the bias in parameter estimates resulting from the use of unit values in place of export prices would not be serious.

IV. EMPIRICAL RESULTS

Estimates of income and price elasticities for exports of iron and steel products for the thirteen countries covered in this study are presented in Table I. It is seen from the table that the estimates of income and price elasticities are almost invariably of correct sign (i.e., consistent with the underlying theoretical considerations) and of plausible magnitude, and for the most part statistically significant. Our finding of a correctly signed and statistically significant coefficient of the relative price variable in most cases for such a large number of countries is remarkable in view of the frequent occurrence of incorrectly signed or statistically insignificant price coefficients in econometric studies on export demand function. Evidently, our results suggest that export prices and the world trade volume (proxy for the income variable) were important determinants of export performance in iron and steel products.

Our estimates show marked inter-country differences in income elasticity for iron and steel exports. The estimated income elasticity is high for France, India, Japan, and Singapore (suggesting thereby that from an increase in total world trade these countries gain relatively more than others); while it is low for Belgium and Luxembourg, Brazil, Italy, and the United States. The coefficient of the world trade variable being negative for Canada, it would appear that income elasticity of Canadian exports of iron and steel products was also low. Comparing income elasticity of demand for aggregate exports across countries, Houthakker and Magee [9] and Goldstein and Khan [8] observed that Japan ranked high

14 Richardson [17] makes a similar argument while pooling data on different product categories of manufacturing exports.

TABLE I

ESTIMATES OF INCOME AND PRICE ELASTICITIES FOR IRON AND STEEL EXPORTS FOR THIRTEEN COUNTRIES, 1966–78

Country	Income Elasticity	Price Elasticity	R^2
Belgium and Luxembourg	0.688*	−0.439*	0.992
	(0.124)	(0.106)	
Brazil	0.508	−1.610*	0.722
	(0.938)	(0.317)	
Canada	−0.071	−0.342	0.896
	(0.283)	(0.225)	
France	1.555*	−0.281	0.976
	(0.188)	(0.201)	
Germany	0.926*	−0.277**	0.991
	(0.119)	(0.118)	
India	1.889*	−0.906*	0.902
	(0.581)	(0.168)	
Italy	0.735*	0.195	0.962
	(0.275)	(0.226)	
Japan	2.151*	−1.063*	0.958
	(0.537)	(0.216)	
Korea	1.240	−0.996*	0.946
	(0.634)	(0.179)	
Netherlands	1.116*	−1.192*	0.985
	(0.183)	(0.081)	
U.K.	0.988*	−0.913*	0.978
	(0.143)	(0.095)	
U.S.A.	0.650**	−0.798*	0.910
	(0.276)	(0.226)	
Singapore	1.918*	−0.928*	0.934
	(0.609)	(0.376)	

Notes: 1. Restricted least-squares method applied to pooled cross-section time-series data with intercept dummies.
2. Figures in parentheses give standard errors.
* Statistically significant at 0.01 level.
** Statistically significant at 0.05 level.

while the United Kingdom and the United States ranked low. It is interesting to note that a similar pattern emerges from our results for the exports of iron and steel products.

As in the case of income elasticity there are also marked inter-country differences in price elasticity. The estimated price elasticity is high for Japan, the Netherlands, the United Kingdom, and all the four developing countries covered in this study (Brazil, India, Korea, and Singapore); while it is low for Belgium and Luxembourg, Canada, France, and Germany. Further, the wrongly signed and statistically insignificant coefficient of the price variable for Italy is suggestive of low price elasticity. Our finding of high and statistically significant price elasticity for all the four developing countries is at variance with the conventional view that demand for exports from developing countries is price inelastic. Our results for the developing countries are in line with the findings of Khan [11] who, in an empirical study covering fifteen developing countries, arrived at the conclusion that prices do play an important role in the determination of exports of developing countries. Our estimates of price elasticity for Japan and the United States are in broad agreement with the estimates presented by Stone [20], but in comparison with Stone's estimate of price elasticity for EEC as an entire unit, our estimates for the majority of EEC countries are substantially low.[15]

Our estimates of price elasticity for most EEC countries being rather low compared to what one would expect, a second exercise has been carried out. The export function for the six EEC countries (Belgium and Luxembourg, France, Germany, Italy, the Netherlands, and the United Kingdom) are estimated as a system using Zellner's [23] estimation procedure for seemingly related regressions so as to incorporate into the export function estimation the interdependence in export performance of the EEC countries since the dominant part of their exports is directed to one another, with the assumption that price elasticity is the same for all six EEC countries. The assumption of a common price elasticity for the EEC countries does not seem unreasonable. Also, the estimate of price elasticity for EEC countries presented by Stone [20] involves this assumption. The results of this exercise are presented in Table II. It is seen from the table that the estimated price elasticity for EEC countries is fairly high, statistically significant, and comparable to the price elasticity estimate for EEC as an entire unit presented by Stone [20]. Estimates of income elasticity for EEC countries presented in Table II are not much different from those presented in Table I. Some difference is noted for Italy, but this is not unexpected since the price elasticity estimate for Italy in Table I is very different from the common price elasticity estimate for EEC countries in Table II.

The coefficients of year dummies reflect the influence of factors other than income and relative prices on export performance over time. Analysis of these coefficients should, therefore, be of interest. We examine first whether the intertemporal variation in the coefficients of year dummies matches with the variation in domestic demand pressure. Evidently, an inverse relationship is expected between the two. To examine this question we have taken the ratio of "apparent consumption" of crude steel to its trend value as a measure of domestic demand pressure and computed the correlation coefficient between this measure of demand pressure and the coefficients of year dummies for each country. Time series on "apparent consumption" of crude steel have been taken from the *Statistical Yearbook*. Though not entirely satisfactory, our measure of demand pressure has the advantage of readily available data and similar measures have been used in earlier studies. Time series on apparent consumption of crude steel not being available

15 Stone's estimates of price elasticity of demand for iron and steel exports are as follows: U.S.A. (−1.20 for product class 672–79); Japan (−4.90 for product class 671 and −1.72 for product class 672–79); EEC (−1.22 for product class 671 and −0.63 for product class 672–79). While comparing these estimates with the present estimates it should be noted, however, that the two studies differ with regard to the time period covered and the estimation procedure.

TABLE II

ESTIMATES OF INCOME AND PRICE ELASTICITIES FOR IRON AND STEEL EXPORTS FOR EEC COUNTRIES, 1966–78

Common price elasticity	−0.755*
	(0.040)
Income elasticity:	
Belgium and Luxembourg	0.716*
	(0.125)
France	1.655*
	(0.165)
Germany	1.116*
	(0.111)
Italy	1.003*
	(0.272)
Netherland	0.999*
	(0.168)
U.K.	0.957*
	(0.137)
Weighted R^2 for the system	0.985

Notes: 1. Zellner's [23] estimation procedure applied to pooled cross-section time-series data, with intercept dummies.
2. Figures in parentheses give standard errors.
* Significant at 0.01 level.

for Singapore, it was excluded from this analysis. Correlation coefficients for the other twelve countries are shown below:

Belgium	0.129	Korea	−0.105
Brazil	−0.258	Netherlands	0.140
Canada	0.077	India	−0.266
France	−0.269	U.K.	−0.312
Germany	−0.336	U.S.A.	0.007
Japan	−0.389	Italy	−0.208

The correlation coefficients are negative (as expected) in most cases, but in no case statistically significant. Thus, it would appear that, in general, variation in domestic demand pressure did not have much impact on export performance in iron and steel products. This conclusion is subject to two important qualifications. First, the ratio of "apparent consumption" of crude steel to its trend value may not adequately represent changes in domestic demand pressure. Second, since the coefficients of year dummies are also affected by factors like product quality, a simple correlation coefficient is not sufficient for assessing the influence of demand pressure. What we require is a multiple regression analysis in which other important factors are brought in. This, however, we could not do for lack of data.

Next, we consider inter-country differences in regard to the broad direction of movement in the coefficients of year dummies. To abstract from short-term fluctuations, five-yearly moving averages have been taken and these are shown in Table III. From this table we can discern three distinct patterns. Accordingly, the countries included in this study have been classified into three groups. In Group I we include those countries for which a distinct rising trend is seen in the coefficients of year dummies, implying thereby that factors other than income and relative prices had a significant favorable effect on export performance. The countries falling in this group are Korea, Japan, Canada, and Italy. In Group II we include those countries for which a distinct falling trend is seen in the coefficients of year dummies, implying thereby that the effect of factors other than income and relative prices has been unfavorable. The United Kingdom, the United States, and France fall in this group. Other countries, for which the pattern of movement in the coefficients of year dummies is not clear, are included in Group III. It is, however, possible to put some of these countries in the first two groups with some ambiguity.

Annual growth rates in iron and steel industry over the period 1966–78 are shown in the last column of the table. It is seen from the table that the countries included in Group I had high rates of growth, while the countries included in Group II had low rates of growth in the iron and steel industry. This is suggestive of a positive association between growth performance and export performance. It may be pointed in this connection that in a cross-country analysis of demand for industrial exports, Sato [18] puts forward the hypothesis that the non-price competitiveness of a country's exports is positively associated with its relative share in the world industrial capacity. This implies that a country growing relatively faster improves the non-price competitiveness of its products compared to others. Using industrial growth rate as a proxy variable for the rate of improvement in non-price competitiveness Sato shows strong empirical support for this hypothesis. The inter-country pattern observed in Table III in regard to movement in the coefficients of time dummies and the growth rate in iron and steel industry is in line with the hypothesis and results of Sato.

V. CONCLUSION

The present paper, pooling cross-section and time-series data at three-digit level of SITC disaggregation for thirteen years, 1966 to 1978, and using dummy variables which allow the intercept of the estimated export function to vary across product classes and over time, has provided a comparative analysis for thirteen countries, developed and developing, in regard to the influence of relative prices and income on the one hand and the characteristics associated with time and product categories on the other, in regard to exports of iron and steel products. Our results on price and income elasticities accord well with existing works in this area. In our study the estimated income and price elasticities are almost invariably of correct sign and of plausible magnitude and are for the most part statistically significant. This shows that relative export prices and world income

TABLE III

MOVING AVERAGES OF THE COEFFICIENTS OF YEAR DUMMIES AND ANNUAL GROWTH RATES IN IRON AND STEEL INDUSTRY, 1966–78

	Moving Averages of the Coefficients of Year Dummies									Annual Growth Rate in Iron and Steel Industry (%)
Group I:										
Korea	−1.604	−1.544	−1.076	−0.445	0.239	0.631	1.079	1.329	1.497	27.36
Japan	−0.234	−0.202	−0.063	−0.037	0.144	0.275	0.184	0.101	0.123	7.26
Canada	−0.284	−0.190	−0.106	−0.056	0.060	0.044	0.104	0.202	0.284	4.51
Italy	−0.326	−0.297	−0.183	−0.139	−0.065	0.077	0.169	0.218	0.340	5.15
Group II:										
U.K.	0.158	0.151	0.096	0.022	−0.069	−0.124	−0.206	−0.195	−0.154	−1.21
U.S.A.	0.152	0.146	0.132	0.118	0.086	−0.007	−0.022	−0.034	−0.124	0.57
France	0.028	0.015	0.018	−0.003	−0.029	−0.042	−0.064	−0.056	−0.020	1.63
Group III:										
India	0.302	0.323	0.094	−0.163	−0.502	−0.708	−0.596	−0.283	−0.006	3.70
Germany	−0.007	−0.003	−0.034	−0.030	−0.006	0.031	0.012	0.006	−0.002	1.90
Singapore	0.119	−0.031	−0.191	−0.258	−0.226	−0.152	−0.114	0.004	0.022	8.31
Netherlands	−0.067	−0.053	−0.023	0.043	0.054	0.054	0.022	0.014	−0.007	5.30
Brazil	0.025	0.133	0.130	0.226	0.152	−0.108	−0.129	−0.104	0.012	11.17
Belgium and Luxembourg	0.030	0.019	0.035	0.051	0.035	0.020	−0.001	−0.022	−0.067	3.00

Note: Annual growth rates in iron and steel industry are based on crude steel production data taken from the *Statistical Yearbook*. Growth rate for Singapore (such data being not available) has been computed from production index for iron and steel for 1970–78 taken from the UN *Yearbook of Industrial Statistics*.

are important variables influencing the export market of iron and steel products. Our results indicate marked inter-country differences in income elasticity and price elasticity. In terms of income elasticity, Japan ranks high; while the United Kingdom and the United States rank low. Estimates of price elasticity for the developing countries are statistically significant and near or above unity. On the other hand, price elasticity estimates for developed countries are in many cases less than one.

We did not find any significant relationship between the movements of the coefficients of time dummies and the demand pressure variable, but the cross-country differences in regard to the movement in the coefficients of time dummies shows in some cases a relationship with the growth rate in the production of iron and steel industry. Considering our results with the findings of Sato [18] it seems to us that relatively faster growth in the production of iron and steel industry in Korea, Japan, Canada, and Italy resulted in a significant improvement in the non-price competitiveness of their products, which in turn affected their export performance favorably.

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