North–South linkages and international macroeconomic policy

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3 Export growth and the terms of trade: the case of the curious elasticities

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A country whose exports grow faster than world income will experience falling terms of trade, unless the income elasticity of demand for its exports is proportionately as high as its export growth rate. This is evident from the conventional export demand equation

$$q_x = \epsilon(p_x - p_w) + \eta y_w \tag{1}$$

where all variables are expressed as rates of change, and q_x is the quantity of exports, p_x is the price of exports, p_w is the price of competing goods in world markets, y_w is real world income, and where $\epsilon\,(<0)$ is the price elasticity and $\eta\,(>0)$ is the income elasticity of export demand. It follows that a condition for stable terms of trade $(p_x - p_w = 0)$ is

$$\eta = q_x / y_w \tag{2}$$

Econometric estimates of income elasticities of export demand indicate that, by and large, this condition holds empirically (Houthakker and Magee, 1969). Countries whose export growth rates are relatively high are shown to have correspondingly high income elasticities of demand for their exports. The one-to-one relationship between estimates of the income elasticity of export demand (η) and the rates of growth of exports relative to world income (q_x/g_w), dubbed by Krugman (1989) the '45-degree rule', is illustrated for samples of developed and developing countries in Figures 3.1(a) and (b) respectively.[1]

The correspondence across countries between export growth rates and income elasticity estimates is too close to be purely coincidental. This curious empirical regularity demands an explanation. There are two possibilities: either (i) the income elasticity of export demand influences the rate of growth of exports, or (ii) the rate of export growth influences, or is systematically related to a bias in, estimates of the income elasticity of export demand.

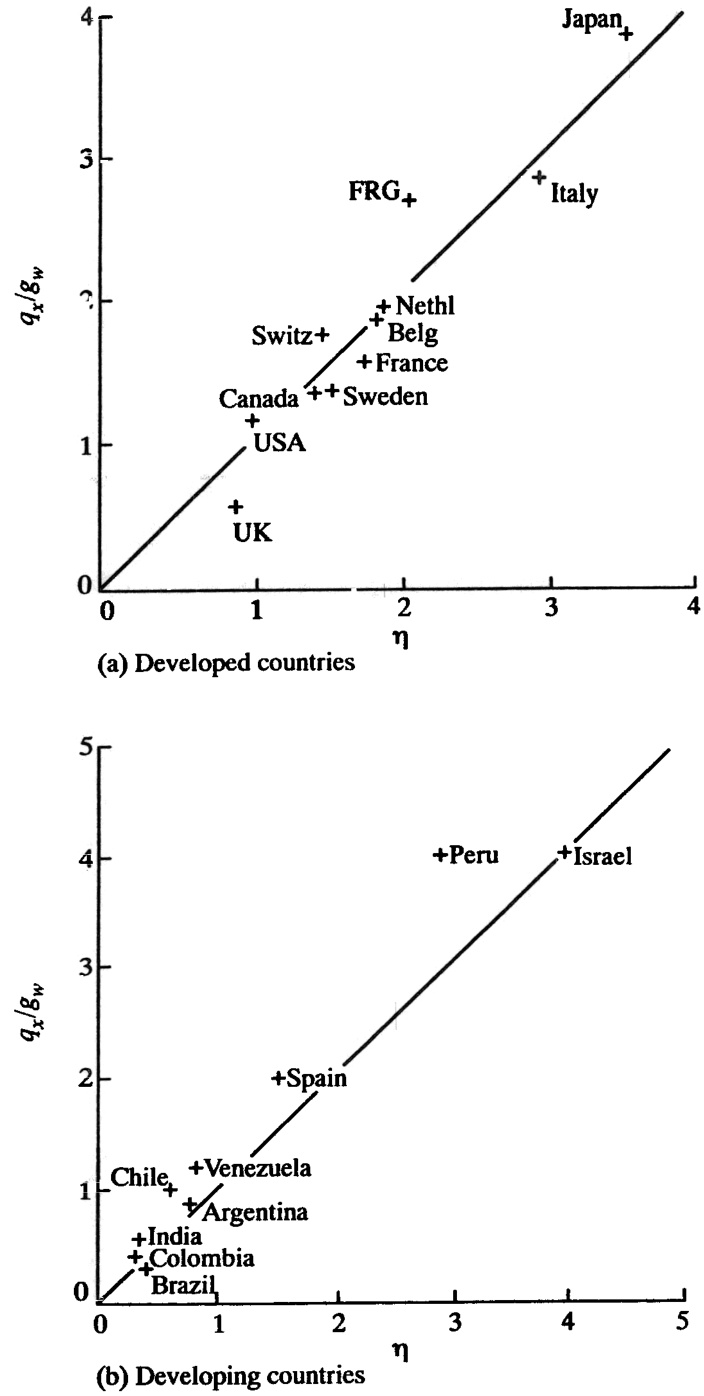

Figure 3.1 The '45-degree rule' for samples of developed and developing countries

1 Hypothesis A: income elasticity of demand influences the rate of export growth

The income elasticity of export demand could determine, or at least influence, export growth in several ways, all of which require that export demand be price inelastic. A low price elasticity of export demand is, in fact, an empirical regularity as robust as the so-called '45-degree rule'. Goldstein and Khan's (1985) exhaustive survey of the literature found a 'consensus view' that the price elasticity of export demand is generally between -0.5 and -1.0, whether the estimate was for geographically and economically large or small countries, for developed or for developing countries, for primary or for manufactured exports.

The rate of growth of export volume (q_x) may become dependent on the income elasticity of export demand (η) if:

(i) policymakers restrict exports to avoid terms-of-trade losses;
(ii) real wages are fixed, eliminating the possibility of real devaluations;
(iii) the supply of exports depends on balance-of-payments constrained imports of investment goods.

However, there is no evidence that the first two mechanisms obtain in practice, while the third mechanism implies a stable income elasticity of export demand for which there is little empirical support (Riedel, 1984). Indeed, it is the non-uniformity of the income elasticity of export demand that we seek to explain. On the face of it, therefore, it would seem that it is export growth that influences estimates of the income elasticity, rather than the other way around.

2 Hypothesis B: export growth influences estimates of the income elasticity of demand

What is suggested is not that export growth influences the elasticity *per se*, but instead that it is systematically related to a bias in conventional estimates of the income elasticity of export demand. Clearly, something is amiss when estimates of the income elasticity of export demand take values which vary so widely for countries at similar levels of development, exporting similar bundles of goods.

Three possible sources of bias in conventional estimates have been suggested: (1) that ordinary-least-squares (OLS) estimates, such as those of Houthakker and Magee, may be biased because they ignore the simultaneous interaction of export supply and demand; (2) that the income elasticity estimates may be biased because they ignore changes in product quality and other forms of non-price competition; and (3) that for a small, price-taking country the conventional estimation procedure is inappropriate and, not surprisingly, yields misleading results. We consider each of these in turn.

2.1 Simultaneous equation bias

An indication of the relevance of simultaneity bias in OLS estimates of export demand equations is revealed by comparing Houthakker and Magee's estimates with those of Goldstein and Khan (1978), who simultaneously estimate export supply and demand equations using full-information-maximum-likelihood (FIML) techniques for a similar sample of countries and similar time period.[2] As shown in Figures 3.2(a) and (b), comparing the two sets of results, the income elasticity estimates are broadly consistent with one another. The '45-degree rule' seems to hold whether the income elasticities are estimated by OLS or by simultaneous equation techniques.

The two sets of price elasticity estimates, on the other hand, differ significantly, with Goldstein and Khan obtaining generally better results in the sense that there are fewer instances of a perverse (i.e., positive) sign. The fact that both price and income elasticity estimates very widely among the countries in the sample, all of which have broadly similar economic structures and face the same world economy, suggests that something other than simultaneity bias is amiss.

2.2 Product quality bias

The argument that estimates of the income elasticity of export demand are biased by the failure to account for changing product quality was revived recently by Krugman (1989), whose reasoning runs as follows:

> Fast growing countries expand their share of world markets, not by reducing the relative prices of their goods, but by expanding the range of goods that they produce as their economies grow. What we measure as exports and imports are not really fixed sets of goods, but instead aggregates whose definitions change over time as more goods are added to the list. What we call 'Japanese exports' is a meaningful aggregate facing a downward-sloping demand curve at any point in time; but as the Japanese economy grows over time, the definition of that aggregate changes in such a way as to make the apparent demand curve shift outward. The result is to produce apparently favorable income elasticities that allow the country to expand its economy without the need for secular depreciation. (p. 1039)

Krugman illustrates this argument in the context of his well-known increasing returns model of intra-industry trade. In that model there are

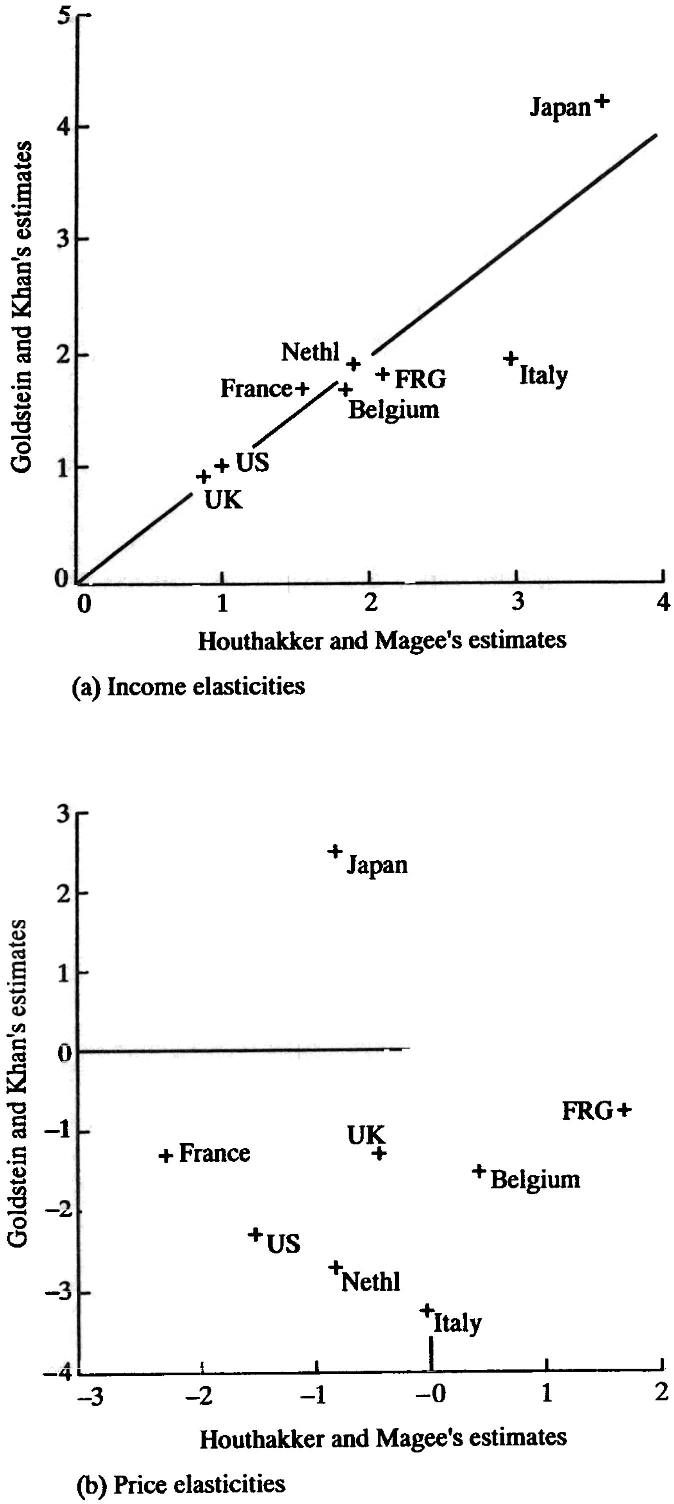

Figure 3.2 Houthakker and Magee's *versus* Goldstein and Khan's estimates of export demand elasticities

no relative price effects; input growth results in a country producing a greater number of goods, not a larger quantity of any good it already produces. Obviously, in such a model of terms-of-trade effect from export growth is ruled out by assumption. Unfortunately, Krugman offers no empirical support for this explanation of the '45-degree rule'. Instead, he suggests (p. 1031) that 'this empirical regularity lends support to a particular view of international trade', which rather confuses whether it is his model of international trade that explains the empirical regularity, or the other way around.

There have, however, been attempts to test the hypothesis that inordinately high income elasticity estimates are due to a bias resulting from the omission of product quality in export demand models. Perhaps the first to examine this hypothesis was Sato (1977), who anticipated Krugman (1989) by a dozen years. Sato contended that the conventional export demand function (e.g., equation (1)) is misspecified by its exclusion of a role for non-price competition, the main form of which, he argued, is product differentiation. Sato suggested that the proper specification of the export demand function is

$$q_x = \epsilon(p_x - p_w) + \eta y_w + \gamma z_x \qquad (3)$$

where z_x is an index of the rate of change in the quality of exports relative to competing goods in world markets, and γ is the elasticity of demand with respect to product quality change. It follows that estimation of (1) yields a strongly biased estimate of η if y_w and z_x are closely correlated.

Sato hypothesized that 'non-price competitiveness is significantly associated with an exporting country's growth performance' (p. 456). As a proxy for growth performance, and hence export quality change, Sato used the growth of industrial capacity, which he showed to be closely correlated across countries with conventional estimates of the income elasticity of export demand. This would be fine if we had evidence that the rate of growth of industrial capacity was indeed correlated with product quality change or with the proliferation of product varieties. Unfortunately, we do not. Hence, the association between the two is purely a matter of conjecture. These findings, therefore, do no more than to illustrate, again, the empirical regularity of the '45-degree rule'.

A similar approach was taken more recently by Helkie and Hooper (1988) who introduce a measure of foreign countries' capital stock into the import demand equation that they estimate for the US, their premise being that such a measure can 'capture the price effects of the introduction of significant new product lines' (p. 20). As they anticipated, inclusion of this proxy variable for changing product variety had the effect of lowering the estimated income elasticity of import demand in the US.

A similar finding is reported by Feenstra (1994), who incorporates a measure of product variety directly into the import price index used in estimating import demand parameters for six disaggregated manufactured goods imported into the US from developing countries over the period 1964–87. Feenstra finds that 'allowing for quality change in the imports from developing countries generally resulted in a price index that rose more slowly than conventional measures'. On the other hand, techniques used by the US Department of Labor to correct for quality change, he notes, 'result in an import price index that rises faster than conventional measures'. Feenstra finds that for two of the six products – men's leather athletic shoes and colour television receivers – the estimated income elasticity of import demand declined when adjustment was made for product quality, as he defines it. In the other cases the adjustment had no significant effect on the income elasticity estimates.

2.3 *The small country case*

It is curious that those who argue that conventional income elasticity estimates are biased are not equally suspicious of the price elasticity estimates that come from the same regression equations. Krugman (1989), for example, goes to great lengths to explain why the conventional estimates of income elasticities of export demand are biased upward, but accepts without question the empirical evidence that price elasticities are low, perhaps because it accords with his model which emphasizes product differentiation and monopolistic competition. However, a price elasticity of -0.5 to -1.0, which is the 'consensus view' in the econometrics literature, suggests far more than product differentiation – it suggests that goods produced in different countries are not even close substitutes. Taken at face value, the conventional estimates of the price elasticity of export demand indicate that most countries have significant market power, implying a potentially important role for optimal export taxes, though this implication is rarely drawn.

Suppose, however, that the export demand equation being estimated is for a small country, one that is a price taker in world markets, exporting goods which are very close, if not perfect, substitutes for those which are produced by its competitors in world markets. In this case, using least-squares techniques to estimate the conventional export demand equation, with quantity as the dependent variable and relative prices and world income as explanatory variables (i.e., equation (1)) is inappropriate, either as a single equation or as part of a simultaneous system of export supply and demand equations. The reason is very simple: for a small country, export quantity is determined exclusively by supply, and export price by the world price.

Estimation of equation (1) for a small, price-taking country yields meaningless results. Indeed, if the export price and world price are identical, estimation of (1) would be impossible since the variable matrix would be singular in that case. If, however, there is variance in relative prices due to measurement errors and other 'noise' factors, an estimate may be obtainable, but little explanatory power would be likely to be claimed by relative price. Instead, world income would generally turn out to be the more statistically significant explanatory variable in estimates of equation (1), even for a small country, since it has been shown by Sato that world income is highly correlated with key export supply variables, such as industrial output. In the case of a small, price-taking country, therefore, the 45-degree rule is a likely, and highly misleading, outcome.

There is, however, a very easy way around this problem. It is simply to estimate the inverse, rather than the standard form, of the export demand equation. If there is any doubt about whether the country being investigated is indeed small, the estimation should be done as part of a system of simultaneous equations. The inverse of (1) is

$$p_x = p_w + (1/\epsilon)q_x - (\eta/\epsilon)y_w \tag{4}$$

which allows for the possibility that ϵ is infinite, something the regular form of the export demand equation does not do. In estimating (4), if it is found that the coefficient on p_w is not significantly different from one, and the coefficients on q_x and y_w are not significantly different from zero, the small country case is confirmed. For such a country, exports can, of course, grow at any rate without affecting the terms of trade.

Two case studies have been published which estimate both the standard form and the inverse export demand equation as part of a simultaneous supply and demand system (Riedel, 1988 and Athukorala and Riedel, 1991). The first was a study of Hong Kong's manufactured exports, using quarterly data for the period 1972 to 1984 (Riedel, 1988); the second was a study of Korean exports of machinery and transport equipment, using quarterly data for the period 1977 to 1988 (Athukorala and Riedel, 1991). Both studies took special care with data collection and with the specification of an export supply equation which was estimated simultaneously by the two-stage least-squares technique. Following previous work, both adopted the partial adjustment mechanism to capture the short-run dynamics, though the focus of the two studies was the long run.

Table 3.1 Published case studies of two small countries: Hong Kong and Korea

	Standard form Dep. var.: Q_x			Inverse form Dep. var.: P_x	
Explanatory var.	P_x/P_w	Y_w	P_w^a	Q_x	Y_w
Parameters	ϵ	η		$1/\epsilon$	$-\eta/\epsilon$
Estimates for					
Hong Kong	−0.70	4.03	1.00	−0.05	0.14
	(−3.78)	(27.00)		(−0.83)	(0.63)
Korea	−0.84	7.22	1.00	−0.002	−0.96
	(−2.15)	(7.93)		(−0.005)	(−1.00)

Notes: The price homogeneity restriction was imposed and tested, and in both cases passed. *t*-statistics in parentheses.
Variable definitions and data sources: see Riedel (1988) and Athukorala and Riedel (1991).

Table 3.1 presents the estimates of the long-run price and income elasticities of export demand derived from estimating both the standard form and the inverse of the export demand equation.

The results, in both cases, for the standard form of the export demand equation and for its inverse present very different pictures. The standard form suggests that both countries face a price-inelastic demand in world markets, but both are favoured by high income elasticities of export demand. The inverse form suggests that both are price takers, facing an infinitely elastic demand in world markets. There is little to choose between the two sets of results on the basis of goodness-of-fit criteria; both yield values for $\bar{R}^2$ above 0.95.

Is an income elasticity value between 4 and 7 reasonable for the kinds of goods Hong Kong and Korea export, indeed for any kind of exports? Is it possible that the optimal tax on exports in these two countries is greater than 100 per cent? On the face of it, this is what the conventional price elasticity estimates imply. How were Hong Kong and Korea able to avoid falling terms of trade, since both experienced exceptionally high export growth rates? (The volume of Hong Kong's manufactured exports grew at an annual rate of 26 per cent over the period of estimation, while the volume of Korea's machinery exports grew at a rate of 23 per cent.) One possibility is that they are small countries and do not influence the prices of their exports, no matter how much they export. Another is that export expansion took the form of increasing the number of varieties exported, rather than increasing the volume of any goods which were already being exported. We have econometric evidence in support of the small country hypothesis,[3] but nothing more than conjecture to support the product diversification hypothesis, which of course does not rule it out.

3 Methodological issues

The evidence cited above suggesting that Hong Kong and Korea are small countries in world trade has recently been challenged on methodological grounds (Muscatelli, Srinivasan and Vines (MSV), 1992). Using Riedel's data for Hong Kong, they argue that the small country results Riedel obtained from estimating the inverse export demand function are biased and invalid due to misspecification of the short-run dynamics.

MSV eschew the partial adjustment mechanism and instead apply cointegration techniques. After finding that the Hong Kong data series are I(1), i.e., integrated of order one, they employ the two-stage procedure for modelling separately long-run equilibrium relationships and short-run dynamics of export supply and demand. The long-run equilibrium relationships are first estimated using the fully modified ordinary least-squares method (FMOLS) proposed by Phillips and Hansen (1990).[4] Utilizing the residual estimates from these long-run equations to represent short-run deviations from the steady-state situation, parsimoniously parameterized error-correction models (ECMs) for export demand and supply were then simultaneously estimated using the full-information maximum likelihood (FIML) method.

MSV's estimation results for the long-run export demand equation are very similar to those Riedel (1988) obtained by applying two-stage least-squares (TSLS) techniques to a partial adjustment specification of these equations. However, their long-run export demand elasticities estimated for the price-dependent export demand function (with the zero price homogeneity restriction imposed) diverge markedly from Riedel's estimates for the inverse demand function. From this evidence, MSV (p. 1475) argue that 'even an economy such as Hong Kong's may face a low price elasticity of demand for its exports and so be demand-constrained in its export markets'.

MSV's attempt to refute the small country findings of Riedel (1988) and Athukorala and Riedel (1991) is unconvincing, however, since MSV did not formally test the small country assumption by imposing a zero coefficient restriction on the export quantity variable (QX) and world income variable (YW) in estimating the inverse export demand equation. They did, however, impose the zero price homogeneity restriction, which potentially could bias the coefficient estimates of the other regressors, even when such a restriction is statistically accepted at a given probability level (Warner and Kreinin, 1983).[5]

A more appropriate way to test the small country assumption, using the Phillips–Hansen procedure, is to estimate the inverse export demand function in unrestricted form and then subsequently test zero coefficient restrictions, first, on the export quantity variable and, second, on the export quantity and world income variables jointly. Going a step further, one could employ Johansen's (1988) maximum likelihood estimation procedure, which also takes explicit account of short-run dynamics in estimating the cointegrating vector and has the added advantage of being a maximum likelihood procedure which, unlike Phillips–Hansen's OLS technique, is not sensitive to the method of normalization adopted.[6]

Table 3.2 reports the results of both of these approaches applied to the Hong Kong data used in Riedel (1988) and subsequently in MSV (1992). Using the Phillips–Hansen procedure to estimate the unrestricted equation, we find that the coefficients on *QX* and *YW* are not statistically significant (at least at the 10 per cent level), though they do have the expected signs. The zero coefficient restriction on *QX* is supported by the Wald test (equation (2)), and the coefficients of *PW* and *YW* show remarkable resilience to the imposition of this restriction. The joint zero restriction on the coefficients of *QX* and *YW* is also data acceptable (equation (3)), and interestingly the world price variable (*PW*) alone explains over 90 per cent of the total variation in the export price variable (*PX*). The coefficient of *PW* is less than unity, reflecting perhaps measurement errors (see note 5 above), but the magnitude of the difference from unity is well within two standard errors of the coefficient estimate. The results based on the Johansen procedure basically tell the same story. The zero coefficient restrictions on *QX* and jointly on *QX* and *YW* are data acceptable.

In short, the inference that Hong Kong is a price taker in world markets, based on OLS estimation of the inverse export demand function (Riedel, 1988) is equally supported by the more robust Phillips–Hansen and Johansen methods, contrary to what is argued in MSV (1992). The same holds as well for Korean exports of machinery and transport equipment. As is shown in Table 3.3, the findings presented in Athukorala and Riedel (summarized in Table 3.1) emerge even stronger when the model is estimated using the Phillips–Hansen and Johansen procedures with the appropriate coefficient restrictions imposed.[7] As these findings indicate, the choice between the price-dependent and quantity-dependent versions of export demand specifications is critically important, even when applying the Phillips–Hansen cointegration procedure.

Table 3.2 Phillips–Hansen and Johansen estimates of the price-dependent export demand function for Hong Kong†

(a) Phillips–Hansen estimates††

1 Unrestricted

$$PX = 0.18 - 0.18QX + 0.82PW + 1.16YW$$
$$(3.22)^{**}(0.98) \quad (5.49)^{*} \quad (1.58)$$
$$R^2 = 0.91 \quad DF = 3.26 \quad PP = 3.72$$

2 With zero restriction on QX

$$PX = 0.19 + 0.72PW + 0.81YW$$
$$(3.17)^{**}(4.50)^{*} \quad (1.19)$$
$$R^2 = 0.92 \quad W(1) = 1.77 \quad DF = 3.10 \quad PP = 3.53$$

3 With joint zero restriction on QX and YW

$$PX = 0.24 + 0.90PW$$
$$(4.70)^{*}(12.87)^{*}$$
$$R^2 = 0.91 \quad W(2) = 3.71 \quad DF = 3.54 \quad PP = 4.08$$

(b) Johansen estimates†††

Unrestricted

$$PX = 0.52PW + 0.53QX - 1.13YW$$

2 With zero restriction on QX

$$PX = 0.75PW + 0.51YW$$
$$LR(1) = 2.46$$

3 With joint zero restriction on QX and YW

$$PX = 0.86PW$$
$$LR(2) = 5.86$$

Notes: See Appendix A.
Variable definitions and data sources: see Riedel (1988)

4. Implications

Price and income elasticities of export demand are central to many economic issues. They are a key consideration in adopting a strategy of economic development and in setting trade policy. Yet the consensus view about the values of these estimates is clearly open to challenge. Everyone seems to recognize that conventional income elasticity estimates are implausible, but many retain faith in conventional price elasticity estimates which come out of the same regression equations. It is our contention, however, that the conventional price elasticity estimates are equally implausible. They are entirely inconsistent with the experience

Table 3.3 Phillips–Hansen and Johansen estimates of export demand functions for Korean machinery and transport equipment

(a) Phillips–Hansen estimate

Unrestricted (quantity-dependent export demand function)

$$QX = \underset{(9.02)}{23.29} - \underset{(2.01)}{1.43}PX + \underset{(2.32)}{1.53}PW + \underset{(10.22)}{5.93}YW$$

$R^2 = 0.95$ DF $= -3.79$ PP $= -4.18$

2 Unrestricted (price-dependent export demand function)

$$PX = \underset{(2.16)}{3.32} + \underset{(0.06)}{0.003}QX + \underset{(22.61)}{1.01}PW - \underset{(1.94)}{0.81}YW$$

$R^2 = 0.94$ DF $= -3.72$ PP $= -4.12$

3 With zero restriction on QX

$$PX = \underset{(5.92)}{3.05} + \underset{(20.12)}{1.013}PW - \underset{(4.75)}{0.69}YW$$

$R^2 = 0.93$ W(1) $= 3.13$ DF $= -3.27$ PP $= -3.48$

4 With zero restriction on QX and YW jointly

$PX = 0.96$PW

$R^2 = 0.88$ DF 3.07 PP $= 3.24$

(b) Johansen estimates

1 Unrestricted

$PX = 1.04PW - 0.01XD - 0.88YW$

2 With zero restriction on QX

$PX = 1.04PW - 1.01WY$

LR(1) $= 0.161$

3 With zero restriction on QX and YW jointly

$PX = 1.03PW$

LR(1) $= 3.42$

Notes: See Appendix B.
Variable definitions and data sources: see Athukorala and Riedel (1991)

of countries which achieved very rapid export growth without suffering deteriorating terms of trade.

Export diversification and product quality upgrading are important and well-documented phenomena, especially in countries with rapid export growth, where comparative advantage and relative factor prices change rapidly. However, there is no empirical evidence to support the argument, put forward by Krugman and MSV among others, that countries can achieve rapid export growth, without suffering declining terms of trade, only by increasing the number of product varieties exported and not by increasing the volume of any existing varieties because of price-inelastic demand. The evidence presented here, for Hong Kong and Korea, suggests instead that the number of small countries in world trade is likely to be much greater than is indicated by previous econometric evidence.

Appendix A Notes to Table 3.2

† The sample period is from 1977Q1 to 1984Q2.
t-ratios of regression coefficients are given in brackets with significance levels (one-tailed test) denoted as: * = 1%, and ** = 5%.

†† W = Wald test for coefficient restriction. Five per cent significant levels for the χ^2 test are W(1) = 3.84 and W(2) = 5.99. DF = Dickey–Fuller test for residual stationarity. PP = Phillips–Perron test for residual stationarity. In all cases, the residual non-stationarity hypothesis is rejected by both DF and PP at the 5 per cent level or better.

††† The cointegration likelihood ratio (LR) tests (based on the maximum eigenvalue and the trace of the stochastic matrix, respectively) suggested the existence of two cointegrating vectors. Only the one with the largest latent root is reported. Given that the estimates are based on quarterly unadjusted data, the VAR length was set at 4. LR = Likelihood ratio test statistic on coefficient restriction. The LR(n) statistics are asymptotically χ^2 variates under the null hypothesis. Five per cent critical values are, LR(1) = 3.84 and LR(2) = 5.99.

Appendix B Notes to Table 3.3

The Cointegration likelihood ratio (LR) tests (based on the maximum eigenvalue and the trace of the stochastic matrix respectively) suggested the existence of a unique cointegration vector. Given that the estimates are based on quarterly unadjusted data, the VAR length was set at 4. Definitions of test statistics are as explained in notes to Table 3.2.

Appendix C Unit root tests

Table 3A.1 Unit root tests for data series employed for the estimation of demand functions for Korean exports of machinery and transport equipment

Variable	Test for I(0)**		Test for I(1)**	
	DF	PP	DF	PP
PX	−1.66	−2.98	−5.76	−9.59
XD	−2.00	−0.55	−6.63	−5.83
PW	−1.20	−1.40	−5.65	−5.54
YW	−0.77	−0.82	−6.56	−7.21

Notes:

* DF = Dickey–Fuller test; PP = Phillips–Perron test. For both tests the text statistic reported is the -*t*-ratio on a_1 in the following auxiliary regression

$$y_t = a_0 + a_1 y_{t-1} + a_2 t + \sum_{t=1}^{p} b_j y_{t-j} + e_t$$

where *y* is the variable under consideration, *t*, time trend, and *e* stochastic error term. In estimating the regression, the lag length (*p*) on the lag-dependent variable was determined to ensure residual whiteness of the estimated equation. Note that in all cases *t* was included in the auxiliary regression to allow for the possibility that for most economic time series the main competing alternative to the presence of unit root is a deterministic linear time trend (Phillips and Perron, 1988).

** The null hypothesis of non-stationarity is not rejected at the 5 per cent level or better for any of the variables.

*** The null hypothesis of non-stationarity is rejected at the 5 per cent level or better in all cases.

NOTES

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1 The '45-degree rule' is observed even in estimates at a more disaggregated level. For example, Bushe *et al.* (1986) found wide variation in their estimates of *n* for machinery exports of the US, Germany and Japan (1.45, 2.46 and 4.93, respectively), which it turns out corresponds closely to differences in the rates of growth of their machinery exports over the period of estimation (5.6, 8.6 and 22.6 per cent, respectively).

2 Houthakker and Magee use annual data from 1951 to 1966, while Goldstein and Khan use quarterly data from 1955 to 1970.

3 The small country hypothesis has also been confirmed, using a different method of analysis, in the case of Taiwan exports of footwear. See Bee-Yan Aw (1993).

4 This is a single-equation semi-parametric least-squares and instrumental variable method which permits direct estimation of the long-run relationships through a two-step method, whereby the data are subjected in the first step to a non-parametric correction for serial correlation and endogeneity.

5 It is true that the small country cases imply a one-to-one correspondence between the prices received by a country for its exports and the world market prices of the same commodity. However, in practice it is difficult if not impossible to obtain actual price series, and instead proxies for the two series have to be used. Given the potential for measurement error, we argue against arbitrarily imposing the price-homogeneity restriction at the outset.

6 The small sample properties of estimates based on Johansen's method have not yet been systematically assessed. However, being a maximum-likelihood procedure, point estimates generated by this method may be biased for small samples. By contrast, the Phillips–Hansen procedure has been found to

perform adequately in small-scale models with as few as fifty sample observations. Therefore, we use the Johansen results only as a check on the more appropriate Phillips–Hansen results.

7 Before applying the Phillips–Hansen and Johansen procedures, we first examined the time-series properties of the Korean data, employing the Dickey–Fuller and Phillips–Perron tests. As shown in Appendix C, in terms of both tests, all the series in the Korea data set are likely to be I(1) processes; they are non-stationary in level form and stationary in first difference form.

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