IS THERE PERSISTENCE IN THE GROWTH OF MANUFACTURED EXPORTS ?

Evidence from Newly Industrializing Countries

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ABSTRACT

Applying cointegration techniques in a panel data setting, we document persistent growth of manufactured exports from certain developing countries. To complement the investigation of persistence (measured by country 'fixed effects'), we analyze asymmetry in income elasticities: for all developing countries, the decline in exports with world income contraction is sharper than is the rise on the upswing; the decline is, however, especially pronounced for countries with low or negative persistence. The results are consistent with long-term buyer-supplier relationships that create "insiders" and "outsiders" in manufactured goods trading. Exports are also influenced the transactional infrastructure (proxied by telecommunications penetration).

(JEL Classification: F13, F14)

I.. Introduction

A country's export demand function relates its export volume to the relative price of its products and to the incomes of international buyers. Price and income elasticities of demand estimated from such functions are used both for predicting exports and prescribing effective export strategies.

The focus on price and income elasticities has, however, led to the neglect of an important empirical regularity: a strong persistence in the growth rate of a country's exports. Persistence can arise from a slow adjustment to short-term demand fluctuations, lasting typically for several quarters. Such inertia is of limited interest to us. In this paper, we are concerned with a persistence of much longer duration, implying the influence of institutional features that exert long-lasting effects on export growth. The paper also is distinguished by the application of cointegration techniques in a panel data setting--in doing so, obtain additional insights into the dynamics of exports from different country groups.

Evidence on persistence can be seen in different versions of export demand functions. When the variables are expressed in levels, export demand functions tend to systematically over- or under-estimate export levels: in other words, the "residuals" (actual minus estimated exports) have a high degree of positive serial correlation, reflected in Durbin-Watson statistics of the order of 0.75 (see, for example, Krugman and Baldwin 1987, Landesmann and Snell 1989, and Bhalla 1989). This same characteristic of export growth is seen more sharply when the variables of the demand function are represented as rates of growth: in addition to growth explained by price and world income changes, a non-zero, country-specific growth rate (fixed effect) is observed.

More often, persistence in export growth rates is obscured due to the use of ad hoc procedures when estimating export demand functions. First, long-term persistence is misread as a short-term adjustment to excess supply and demand conditions and is accounted for (incorrectly, in our judgement) by the inclusion of lagged export volume as an "explanatory" variable (for a recent example, see Marquez and McNeilly 1988).

Second, persistent evolution of export volumes is subsumed in high income elasticities of demand. For some industrial countries (notably Japan) and many developing countries, income elasticities are in the range of 2.5 to 5.0, i.e., a one percent increase in world income increases their exports by 2.5 to 5 percent (see Marquez and McNeilly 1988 for a review of selected studies). Recently, Muscatelli, Srinivasan, and Vines (1992) estimated Hong Kong's income elasticity of demand to be 4.2. Most authors are generally uncomfortable when reporting such high elasticities. Muscatelli, Srinivasan and Vines (1992), for example, note that the high elasticities are due to: "a failure of conventional models of export flows (including our own) to identify important forces causing shifts in export demand: 'income effects' thus probably subsume a variety of other non-price factors." We will show that while their instinct on shifts in demand are right, their interpretation of high income elasticities is probably faulty.

Finally, Helkie and Hooper (1988) provide a more explicit accounting of persistence, using the stock of capital in the exporting country as a proxy for secular changes in the capability to supply an increasing range of products. Their defense for "this unabashedly ad hoc adjustment is that the existing price indexes do not adequately capture the price effects of the introduction of new product lines." Similarly, Krugman and Baldwin (1987) add a time trend variable to their export demand equation to account for long-term changes.

Our purpose is to cast a spotlight on the long-term persistence found in the data, examine its robustness, seek statistical proxies that may account for the persistence, and provide an interpretation of the observed patterns. In pursuing this investigation, we believe that we have identified a much richer set of export determinants than are implied in the traditional models that focus on income and price elasticities and short-run adjustment.

Our first effort is to identify and examine sources of errors and biases that may lead to exaggerated estimates of the persistence effect. A specific concern is the existence of errors in the measurement of relative price. Aw (1992) and Feenstra (1992) have taken the approach that such errors are minimized when the demand equation is estimated for narrowly defined products rather than for manufactured goods as a whole. Although measurement errors obviously exist, and the choice of instruments used to account for the errors has an influence on estimated price elasticities, these considerations are not sufficient to explain away the long-term country-specific persistence.[1]

The observed persistence reflects a diffusion of demand from industrialized to newly industrializing economies and could be considered an evolution by developing country exporters from "outsider" to "insider" status. An outsider is a marginal supplier; an insider is a supplier with whom the buyer has a long-term relationship in which both parties have made (tangible and intangible) investments. Insiders are part of an extensive network of buyer-supplier relationships and draw on this capital to maintain high growth rates.

We provide indirect evidence of an "insider-outsider" phenomenon in world markets for manufactured goods by examining asymmetries in the income elasticity of demand for different groups of countries. Specifically, we find that the magnitude of the export response depends upon whether the buyers' incomes rise or fall. When world income rises, exports rise relatively uniformly for different country groups; the decline in exports with world income contraction is larger and especially sharp for certain countries. Suppliers facing high elasticities on the down-side are marginal to the buyer. When the distinction between the rise and fall of world income is not made, high (average) elasticities are often incorrectly interpreted as a sign of successful export performance.

Countries that have profited from the shift to insider status have not been passive beneficiaries. Rather, they have invested in improving their transactions infrastructure, making

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1 Benhabib and Jovanovic (1991) also observe persistence over 15 to 25 years in the growth rates of per

them easier to do business with. The development of a country's telecommunications network appears to be a partial proxy for the ability to deliver time- and communication-sensitive services that are relevant for developing country exports of such goods as garments, shoes, bicycles, consumer electronics, and auto parts.

The paper is organized as follows. In section II, we describe how the degree of persistence has varied across (and within) countries over time. We also discuss and evaluate issues relating to mismeasurement and the choice of proper instruments. In Section III, we examine asymmetries in the income elasticity of demand. The use of telecommunications as a proxy for the quality of a country's transactional infrastructure is described in Section IV. The conclusion evaluates the evidence and comments on its policy relevance.

II. Patterns of Persistence

Our task is two-fold. First, we need to establish a framework within which to estimate persistence. Second, given the relative novelty of the exercise, we need to test the robustness of our findings.

We estimate an export demand function in first-differences (i.e., the relevant variables are measured as rates of growth). Persistence is defined as the underlying growth rate after price and income effects are accounted for. Since we are working with a panel data set (time-series information for a panel of countries), our measure of persistence is the "fixed-effect" for each country.

To test the robustness of this framework, several steps are undertaken. We begin by ensuring that measured persistence for a specific country is not a side-effect of forcing the price and income coefficients to be the same for all countries in the panel. Next, since variables are measured in first-differences (or rates of growth), we examine the possibility of misspecification

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capita incomes in a wide range of countries.

on account of omitting long-run influences from the export demand equation. Further, choice of the starting and ending dates of the panel could influence the estimates: we, therefore, examine whether the persistence measure is relatively stable over time for any one country. Finally, export demand analysis has traditionally been plagued by errors in measuring the price variable and the effects of such measurement errors are determined.

We begin with the following demand equation:

$$\Delta \log E_{it}^{d} = \gamma_i + \alpha_0 \Delta \log(P_{i,t}^{x} / P_t^{w}) + \alpha_1 \Delta \log(P_{i,t-1}^{x} / P_{t-1}^{w}) + \beta \Delta \log Y_{it}^{w} \tag{1}$$

Equation 1 is designed to estimate a set of parameters from pooled observations for a number of countries and several years. Each variable is defined for a country i and a time t. E^d is the demand for a country's exports, P^x is the price of the country's exports, P^w is the price of exports by competitors (proxied by a world price index) and Y^w is the world income relevant to the country (the weighted sum of the purchasing countries' GDP, where the weight is the average share of each purchasing country in the total exports of the country in question for the sample period).[2] The coefficients _ and ß are the price and income elasticities of demand, respectively.[3] The symbol _ indicates that the equation is specified in first differences (which approximates to the rate of growth when variables are expressed in their logarithmic value). The first-differences export demand equation is preferred to the equation in levels, because the variables of the export demand equation are not stationary in levels, but they are stationary in first-differences as further described in section II.C and Appendix B.

Of specific interest is _$_i$, which in our pooled cross-section time-series setting summarizes, for each country i, the effect of country features that we do not observe. These

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[2] We choose to use the average export share for the sample period, rather than the export share for the corresponding year to avoid the possibility of introducing endogeneity into our world income variable.

[3] Riedel (1988a) has argued that the results are substantially different when, in contrast to the procedure adopted here, price is used as the dependent variable and export volume is the independent variable. Muscatelli, Srinivasan, and Vines (1992) show, however, that the normalization (or the choice of the dependent variable) does not matter once serial correlation and endogeneity are accounted for. Riedel's results, therefore, appear to arise from the non-stationarity of the variables, leading to a "spuriously" strong correlation between the country's export price and the world price and eliminating all other partial correlations.

unobserved country features include variables that are difficult to observe, such as the strength of international marketing relationships between suppliers and their international buyers, or can in principle be observed but can be measured only imperfectly, such as the quality of a country's infrastructure. These features are a potential cause of export growth persistence, and hence we begin our description of the data by treating _i as our measure of persistence. As constructed, _i remains unchanged over time; however, by considering overlapping slices of time, we are also able to follow the evolution of _i.

The questions of interest are: first, are the _i's different from zero, i.e., is there persistence in export growth rates; and, second, are the _i's different from each other, or do the influences causing persistence vary by country?

If the _i's are different from zero and from each other, then the unobserved differences across countries apparently have a significant influence on export growth. Benhabib and Jovanovic (1991) note that persistence in per capita income growth rates across countries could either reflect country-specific features or all countries could be influenced by the same stochastic forces but the specific realization for different countries could vary and cause long-term differences. They favor the latter interpretation for its parsimony. While empathizing with this view, we choose to focus on the specific country correlates of persistence rather than attempting to identify the common stochastic structure of knowledge and institutional evolution.

It is common in such estimations to allow for lags in the response of exports to the variables influencing them. We allow throughout for lags in response to price changes. Hence the previous year's price (with the subscript "t-1") is included as an explanatory variable. Landesmann and Snell (1989) and Krugman and Baldwin (1987) show empirically that lags in the world income variable do not have much explanatory power. Landesmann and Snell argue that this is to be expected since changes in relative prices require shifting to new buyers and hence imply a lag, whereas changes in world income do not require such shifts and hence lags are not likely. Our estimations confirmed this result and we do not report them here.

A. Testing the Specification

We first present the simplest credible specification. Here all countries are pooled together, no allowance is made for long-run dynamics, but the issue of simultaniety is dealt with.

The export demand function is estimated for a set of 20 developing countries, from 1972 to 1985. The data for developing countries are pooled when estimating the demand function. The advantage in pooling the data is that we have sufficient degrees of freedom to estimate the coefficients with some accuracy; the disadvantage is that the price and income elasticities are assumed to be equal across the countries. We show in the next section that the basic finding of persistent country-specific fixed effects remains unaltered even when price and income elasticities are allowed to vary across the countries in the sample. Results for 13 developed countries over the same period are also used wherever relevant.

To allow for the possibility of simultaneous determination of export volume and relative prices, we use the two-stage least squares procedure (2SLS), where the instruments used for the endogenous relative price variable are: lagged exports ($\Delta E_{t-1,t-2}$), lagged relative export prices ($\Delta(P^X/P^W)_{t-1,t-2}$), current world prices ($\Delta P^W{}_t$), current and lagged wages ($\Delta W_{t,t-1}$), current and lagged world income ($\Delta Y^W{}_{t,t-1,t-2}$), and lagged imports of capital goods ($\Delta K_{t-1,t-2}$) -- all variables are expressed in logs.

In the first equation of table 1, $\_{i}$ are free to take on any value, allowing for the possibility that $\_{i}$ differ by country. This is our most general model. The hypothesis that the $\_{i}$ are equal to zero is rejected very strongly (p-value of the Wald statistic is 0.005).[4] The hypothesis that all $\_{i}$ are equal, though not necessarily zero, is also rejected with a p-value of 0.016.

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[4] In the presence of heteroskedasticity the least squares standard errors are biased. In our estimations we use a consistent estimator of the covariance matrix (White 1980). As a result the Wald test for the joint significance of the coefficients has a Chi-square distribution rather than an F distribution.

Thus the evidence on country-specific fixed effects terms, and hence on the persistence of growth rates in specific countries, is strong. Differences in fixed effects between countries are also evident; as we shall see below, two groups of countries have very different fixed effects and also face different buyer behavior.

The second column in table 1 shows that the 20 developing countries as a group experienced a statistically significant persistent growth of 4.4 percent a year, over and above that explained by relative price and world income changes. It is sobering to reflect that during this period the average rate of growth of manufactured exports from these countries was 12 percent; thus about one-third of export growth depended upon factors not conventionally accounted for. When the second and third columns in table 1 are compared, accounting for persistent growth rates substantially lowers the income elasticity of demand from 4.1 to 3.1.

For individual countries, such as Turkey, Indonesia, and Republic of Korea, the proportion of growth explained by the persistent country-effect was much larger than the average of one-third for all countries (see column 4 in table 2). At the other extreme, a few countries that had negative underlying persistent growth rates could have doubled their export growth if they had lost their handicap. Another way to assess the importance of country-specific effects is to subtract them from the actual growth rate to arrive at the growth rate that would have occurred if the underlying persistent country-specific export growth rate had been zero (column 3). Though the actual growth rates vary substantially, the growth rates net of country-specific fixed effects are much closer to each other. In other words, if the countries with high fixed effects terms did not have their unobserved advantages, and the countries with low fixed effects terms did not have their unobserved disadvantages, the export growth rates of different countries would have been fairly close! The implication is that the degree of relative price changes or the choice of specific export destinations had a much smaller bearing on export performance than the factors that caused the persistent growth rates.

A similar experiment with the developed countries yielded interesting results. The persistent growth rates for these countries are relatively small, generally between -2 and 3.6 percent, and these could not be considered statistically different from zero.[5] As expected, Japan has a positive persistent growth rate, but it is small (less than 3 percent). Somewhat surprisingly, Germany has a small negative persistent growth rate. As we discuss below, the relatively small size of the persistent growth rate for developed countries suggests the secular shifts that successful developing countries can benefit from diminish as the country secures its position as an insider in international markets.

B. **Country Groups**

One potential problem with our estimates of $\_i$ is that the price and income elasticities have been constrained to be equal across countries. Hence, if the countries with rapidly growing exports had larger elasticities, averaging across countries would result in high positive estimated fixed effect coefficients; similarly, it would not be surprising if countries with negative fixed effect coefficients are losing market shares as a result of below-average income and price elasticities.

Ideally, the export demand function should be estimated for each country. However, the limited degrees of freedom make the estimates imprecise as well as unstable. To overcome this limitation we adopt two strategies. First, we allow the price elasticities for specific countries to differ from the average -- for example, we allow the price elasticity of Turkey and Indonesia, which have the highest fixed effects estimates, to differ from that of other countries in Group I (by creating dummy variables for each of these countries and interacting the dummy with the relative price). The results show that income and price elasticities for these countries are not

[5] The residual growth rates for developed countries are not reported in a separate table, as they are not the main focus of the paper.

statistically different from the elasticities for other developing countries (with a p-value of 0.76), and that large fixed effects remain.

The second strategy was to split the countries into two groups -- those with positive fixed effects (Group I) and those with zero or negative fixed effects (Group II). Using the Wald statistic with heteroskedasticity-robust estimate of the covariance matrix, we tested whether splitting the countries in this fashion is supported by the data. The data strongly reject the restrictions imposed by pooling all developing countries in the sample, with a Wald statistic of 14.0 and a p-value of 0.007, thus supporting the split. Certain countries on the margin were not easy to classify; however, the exact composition of the two groups did not alter the results. To be precise, the general observations from the regression results remain unaltered; the interpretation of the performance of the specific countries on the margin, however, does change.[6]

The following differences between the two groups emerge. First, the average fixed effect for Group I is 9.4 percent and significantly different from zero with a t-value of 4.69, whereas the average fixed effect for Group II is -1.0 percent and statistically insignificant with a t-value of -0.38. The price elasticities for the two groups (the sum of the current and lagged values) are quite similar. The point estimate of the income elasticity for Group II is actually higher than that for Group I. Though the difference between the income elasticities of the two groups is not statistically significant, the larger point estimate for Group II is unexpected and we return to this issue below.

Table 3 shows that country-specific fixed effects are significantly different from each other in Group I, the p-value of the Wald statistic is 0.009. As can be expected, the fixed effects

[6] We tested again whether the assumption of equality of coefficients within groups could be maintained. A referee suggested that the equality of price elasticities was to be expected across countries based on a reading of the previous literature and it was especially important to test for equality of income elasticities. We found that we cannot reject equality within the country groups chosen. Most importantly for this paper, the p-value for the null hypothesis that the income elasticities are equal for Group I countries is 0.39. Equality is maintained for other groups too. Thus, while variation in income elasticities is higher across countries than is the case with price elasticities, the main differences is between developed countries and the rapidly growing developing countries (most of whom are in our Group I). Once countries are separated into groups, the variation within groups is much less.

terms obtained for individual countries are now different from the ones obtained from pooling all developing countries. However, the orders of magnitude and relative rankings are very similar (compare the first and the last columns in table 2). The persistent growth rates of Indonesia and Turkey are 19 and 17 percent, respectively, whereas Korea's is 11 percent. When equation 1 is estimated for Group I countries, Portugal's fixed effects term becomes positive (in contrast to the result when it is estimated for all developing countries, see Table 2).

Group II countries have negligible country-specific fixed effects when considered as a group, though some have individually negative persistent growth rates. India, for example, records -4 percent. Venezuela has a relatively high 3 percent; however, the time pattern of the growth rate is erratic, in part because its exports tend to be dominated by petroleum-related products. Accordingly we keep it in Group II.

Note also from table 3 that the income elasticity of demand is lower for the products from developed countries than the ones from developing countries -- this is commonly observed in the literature and attributed to the inclusion of high-growth countries in the developing country sample. However, Table 3 shows that the difference between developed and developing countries' income elasticities originates in part from the high elasticity for the Group II countries (3.53), which have low export growth rates and, typically, negative fixed effects. The paradox of high income elasticities in Group II countries is discussed in Section III, leading to a new interpretation of conventionally estimated income elasticities.

C. **Non-Stationary Variables and Cointegration**

Based on unit root tests, we now show that the specification of the export demand equation in first differences (equation 1) is a reasonable one. Unit root tests, which will be summarized below, verify that variables in the export demand equation are not stationary in levels, but are stationary in first-differences. However, a first-differenced equation does carry the possibility of misspecification. Specifically, the model in first differences implies, in essence, a short-run relationship between a country's exports, its relative export prices and the world income it faces. Thus, it does not take into account any long-run equilibrium relationship between the variables.

When the variables measured in levels are non-stationary, the estimated equation is spurious if the residuals from the equation (the difference between the actual exports and predicted exports) are non-stationary. However, a long term relationship is said to exist between the variables (they are said to be cointegrated) if the residuals from a levels equation regression are stationary; and in that case, the estimates from the levels regression (even though the variables are non-stationary) are no longer considered spurious. The residuals obtained from such a cointegrating relationship are thought of as deviations from a long-run relationship.

If the model is being estimated in first-differences, correct specification requires that the cointegrating residuals, lagged one period, be included in the first-differences model. If, however, the residuals from the levels equation are non-stationary, there exists no long-run equilibrium relationship and the model in first-differences is correctly specified (or, at least, is not misspecified on account of a neglect of the long-run equilibrium relationship).

We tested whether the variables of the export demand equation (exports, relative price of exports and world income) are cointegrated. This was done by first testing whether the variables follow a unit root process, that is, whether they are non-stationary in levels, but stationary in first differences. If all or some of the variables have unit roots, then the next step involves testing whether the residuals from the levels equation are stationary. If the variables themselves have unit roots but the residuals from the levels equation are stationary, then we conclude that these variables are cointegrated.

If our data were only in a time series form, it would be relatively easy to test for unit roots using a test procedure, such as the Augmented Dickey-Fuller (ADF) test. The ADF test for a time series $\{y_t\}$ regresses the first difference Δy_t on the lagged level y_{t-1}, and p-lagged first differences and the appropriate deterministic variables. ADF then looks at the significance of the coefficient on the lagged levels variable. If it is significant then ADF rejects the presence of a unit root in the time series $\{y_t\}$.

Due to the presence of the cross-section variation in panel data, it is incorrect to apply the standard ADF or other time series unit root tests to panel data. Testing for unit root for each country separately, on the other hand, will not produce reliable results, because unit root tests have very limited power against alternative hypotheses in small samples. In a recent paper, Levin and Lin (1993) have extended unit root tests for time-series data to the panel data setting. Being a residual-based test the Levin-Lin test is similar in principle to Augmented Dickey-Fuller test. However, it is more complicated to compute the Levin-Lin test statistic because of the required transformation and normalization of the data. We provide a short summary of the Levin-Lin test in Appendix A. Other than the Levin and Lin study, we know of no other application of cointegration techniques to a panel data setting.

The results (presented in Table B1) lead to the rejection of the null of no-cointegration (among the variables of the export demand equation) only for Group I developing countries. For Group I countries all three variables follow a first-order integrated process: they are non-stationary in levels but stationary in first differences. Since the cointegrating residuals are stationary, we can conclude that for Group I developing countries there exists a long-run equilibrium relationship among the variables of the export demand function. For other country groups, all variables, except world income variable faced by developed countries, follow a first-order integrated process. Yet, we are unable to reject the null of no-cointegration because the cointegrating residuals are also non-stationary for these groups. The failure to reject the null of no-cointegration for these country groups implies that our export growth equation (equation 1) is not misspecified for these groups. Thus, in the rest of the paper, we present results which include the cointegrating residuals for Group I countries but not for other country groups. We should note that though, the results obtained by consideration of cointegration are more accurate, as a practical matter, they vary very little from results obtained without cointegrating residuals.

Besides the technical merit of correcting for misspecification, certain interesting insights flow from this analysis. For Group I countries, long-term influences are important in two distinct ways. Strong persistence of growth exists as measured by the country-specific fixed-effects. They also are the only group where a long-term relationship exists between the levels of exports, prices, and world income. When the cointegrating residual is added to the first-differences equation, the sign on that variable is negative and highly significant. Thus, deviations in previous periods from expected performance lead to a "correction" in the next period. In contrast, Group II countries have no persistence or long-term relationship between the variables of the export demand function. Thus, the evolution of exports in Group II countries follows from one-year to another--there is no influence of history.

D. Does the extent of persistence change over time?

To study changes in the country-specific fixed effects over time, we create seven-year overlapping "windows" in our sample period. The first window covers 1972-1978, the second: 1973-1979, the third: 1974-1980, and so on. We thus have eight windows. For each window we estimate the export demand equation. For each country, therefore, we obtain eight $_i$s. (See Figure 1, where it should be noted that year refers to the final year of the window.)

Despite fluctuations, most countries in Group II (Argentina, Colombia, Chile, India, Pakistan, Yugoslavia, and Israel) had low fixed effects throughout the period (Figure 1a) ; in contrast, most countries in Group I (South Korea, Singapore, Brazil, Malaysia, Thailand, Philippines, Spain, and Greece) had relatively high fixed effects (Figure 1b).

However, both increasing and decreasing trends are also discernible, indicating that the groups are not closed. Indonesia, Turkey, Portugal, Venezuela, and Mexico have steadily increased the size of their fixed effects (Figure 1c). Significant realignments occurred from 1981 to 1983, during a severe downturn in global economic activity. In these years, some of the East Asian newly industrializing economies (such as Korea, Singapore, and Taiwan) experienced rapid wage growth. Countries that increased their fixed effects coefficients during those years have continued to increase them. A number of countries suffered a sharp decline in fixed effects coefficients during that period and have not recovered: most of these were countries that already had low fixed effects -- India, Israel, Argentina, Chile, Colombia, and Yugoslavia; however, Greece and Philippines also suffered.

Further, in the early 1980s, estimated fixed effects coefficients of countries that performed very well in the 1970s, including South Korea, Singapore, Brazil, Malaysia and Thailand began to decline (Figure 1b). If the high fixed effects term represents the transition from outsider status to the ranks of the insiders, a decline in the fixed effect toward zero represents their maturity as insiders.

We draw three inferences from these observations. First, the growth of exports due to unobserved factors tends to persist over time within a country. Second, when underlying cost conditions change, and for instance, Group I countries become expensive producers, new entrants are likely to gain. Third, these shifts take place over time but can be accentuated by downturns in the world economy. During such periods international buyers seek new suppliers. Firms and countries that are well-positioned in such years stand to make large gains.

E. **Mismeasurement and Incorrect Instruments**

Since the country-specific persistent influences could be merely a reflection of errors in measuring the relevant price and income variables, it is necessary to take into account the proposition that if all variables were correctly measured, the observed persistence would disappear. The implication would be that price and/or income elasticities are much higher than typically estimated. A similar argument would hold if the simultaneous determination of export prices and volumes was not fully accounted for. The use of incorrect instruments for relative price in the export demand function would lower the (absolute) value of the price elasticity of demand.

A common problem in estimating export demand functions is that the price variable is not measured correctly. The unit values typically used, as is the case here, do not account adequately for changes in the composition of exports. Thus, for countries that shift toward products with higher prices per unit of product sold (from t-shirts to televisions), the unit value index understates the price increase (See Alterman 1991, who finds that this has been the case for U.S. imports from some developing countries).

Other factors will lead to an overstatement of the price change by the unit value index. If new products include an increasing fraction of a country's exports, the true price index for the enlarged bundle of goods will be lower than the conventionally measured index (see Feenstra 1992). The effect of the increased bundle of goods is identical to unmeasured quality improvements or greater "taste" for that country's goods in world markets. Feenstra (1992) has attempted to construct price indices that reflect the introduction of new products and the exit of old products in the goods supplied by developing countries. As noted, Helkie and Hooper (1988) use a more direct (though more approximate) approach by including in their demand function the capital stock of the supplying country as a proxy for the ability to supply new products.

In our estimates, the measurement problem is alleviated by the use of instrumental variables and by data transformation. In correcting for simultaneity, we use wage rates and other proxies for production conditions as instruments. In general, the solution for simultaneity is the same as that for measurement error, and we have, in principle, corrected for simultaneity. The question is whether our correction is adequate. Specifically, have we adequately accounted for influences on the supply of exports? If not, the persistence being picked up in the demand function could well be the result of ignoring supply factors rather than a feature of the demand function.

We do not believe that it will be possible to fully resolve the question of whether the observed persistence derives from supply or demand factors. However, additional results suggest that while we have not fully accounted for all supply influences, efforts at refining the export supply equation are not likely to have much power in eliminating the persistence effects observed.

To test the sensitivity of our results to the choice of instruments, we experimented with dropping instruments individually or in groups. The overall conclusion is that neither the elasticities nor the fixed effects coefficients change significantly. Only when the world price was dropped from the list of instruments, the estimated price elasticity of the Group I countries increased, with no significant effect on the country-specific constant terms.

These results are similar to those obtained by Feenstra (1992). In a more sophisticated correction of price changes, he finds that the quality- adjusted price for many developing countries rose more slowly than conventional estimates suggest. The correction leads, in his case, to a higher price elasticity of demand and a lower income elasticity, although the changes are limited in magnitude.

A second approach to dealing with measurement errors is through the transformation of data. At least since Griliches and Hausman (1986), it has been known that specific transformations of panel data can be used to minimize the influence of measurement errors; it is also the case, however, that certain transformations of the data can exacerbate measurement errors. We measure the variables as rates of change (first differences of log values), a procedure that is generally considered to increase the "noise-to-signal" ratio, if the variable under consideration is serially correlated. However, Griliches and Hausman (p. 100) note that if the measurement error, rather than the variable itself, is serially correlated, then first-differencing helps to reduce the noise and to increase the signal. In our situation, we can expect the measurement errors to be highly positively correlated, since quality and compositional changes in exports are not random effects that vary from year to year, but represent changes over time. This is at least partially borne out by the figures in Alterman (1991), in which unit value indices are compared with "true" price indices: the errors show significant positive serial correlation. Thus first-differencing is likely to be an effective method of reducing measurement errors.

Griliches and Hausman (1986) propose a test to determine whether the presence of measurement errors is a source of bias in parameter estimates. The test involves a statistical comparison of the GLS (random effects) and the fixed effects (within) estimators.

Adapting their framework and allowing for the possibility that the unit values reflect export prices only partially, we write the percentage change in the export unit value (ΔI_{it}) as the sum of the percentage change in the unobserved export price (ΔP_{it}) and the measurement error (v_{it}): $\Delta I_{it} = \Delta P_{it} + v_{it}$. When the export demand equation (equation 1) is estimated with ΔI_{it} and $\Delta I_{i,t-1}$ rather than the true price variables, the residual term will incorporate $\_{0}\, v_{it} + \_{1} v_{i,t-1}$, which will be correlated with ΔI_{it} and $\Delta I_{i,t-1}$. If the measurement errors are negligible and there are no other sources of correlation between the residuals and the unit values, both models (GLS and fixed effects) will be unbiased and consistent; more importantly the parameter estimates from the two models (GLS and fixed effects) will be asymptotically equal. However, when the correlation between the composite residual term and the observed export unit value is statistically significant, both models will produce biased parameter estimates and these estimates will not be close to each other. As a result, it is possible to test for the statistical significance of measurement errors indirectly via a test for the equivalence of the GLS and the fixed effect estimates using the Hausman specification test (Griliches and Hausman 1986).

P-values for the Hausman test (which has a $\_^{2}$ distribution under the null hypothesis) are reported in Table 3 for both the first difference and the levels estimation. The test results show that the first-differences model is not plagued by the use of unit value indices. In contrast, however, we cannot reach the same conclusion for the levels estimation. Hausman tests for all country groups have very small p-values, which implies that the levels estimation produces biased estimates due to measurement errors; the highest p-value is obtained for the developed countries. This result is consistent with the fact that the composition of exports from developed countries is more stable over time compared to exports from developing countries, and that the quality of the data from developed countries is higher.

III. Asymmetries in Income Elasticity: Insiders and Outsiders

Traditionally, the effect of non-price factors is thought to be captured by the response of exports to changes in world income (summarized in _, the income elasticity of demand). A high income elasticity of demand is considered a measure of superior quality, although as noted, the relationship does not seem obvious from table 3. The income elasticity of demand for products from developing countries is much higher than that from developed countries. Should that be read to imply that developing countries export higher quality products or are able to rapidly expand product variety (Krugman 1989 and Muscatelli, Srinivasan, and Vines 1992)? Within developing countries, income elasticity is much higher for the lagging Group II countries than for the dynamic Group I countries, which is contrary to what would be expected if income elasticity were a good measure either of product quality or of expanding product variety.

The paradox is resolved when we consider the possibility that exports respond asymmetrically to changes in world demand. We test the proposition that income elasticity is different when world income rises than when it falls in the following equation[7]:

$$\Delta \log E_{it}^{u} = \gamma_i + \alpha_0 \Delta \log(P_{i,t}^{x} / P_t^{w}) + \alpha_1 \Delta \log(P_{i,t-1}^{x} / P_{t-1}^{w}) + \beta^{+} \lfloor \Delta \log Y_{it}^{w} \rfloor^{+} + \beta^{-} \lfloor \Delta \log Y_{it}^{w} \rfloor^{-}$$

where:

[7] More refined non-linear responses could be tested but the results presented are striking enough. The years in which world income fell were: 1974, 1975, and 1982.

$$[\Delta \log Y^{w}]^{+} = \begin{cases} \Delta \log Y^{w}, & \text{if } \Delta \log Y^{w} \geq 0 \\ 0 & \text{otherwise.} \end{cases}$$

$$[\Delta \log Y^{w}]^{-} = \begin{cases} \Delta \log Y^{w}, & \text{if } \Delta \log Y^{w} < 0 \\ 0 & \text{otherwise.} \end{cases}$$

The effects of world income change are sharply asymmetric (table 4). In years of rising income, the elasticities (_$^{+}$) for the different groups of countries are fairly close -- between 1.55 and 1.84.[8] The response to a decline in world income, however, is much higher and more heterogeneous.

The difference between income elasticities when incomes rise and fall is small and statistically insignificant in the case of developed countries, with a p-value of 0.22. The difference is substantial for Group I countries but not significant at the 5 percent significance level, and large and significant for Group II countries with a p-value of 0.023. When world income rise by 1 percent, exports of Group II countries increase by almost 2 percent (_$^{+}$=1.84); but when world income falls by 1 percent, exports from these countries decline by 14 percent (_$^{-}$ =14.0). Thus, in times of rising incomes, all countries gain equally in terms of export growth. When world income declines, however, the Group II countries are hit the hardest; since they do not enjoy any special advantage when the upturn occurs, they lose market share over time.

If product characteristics were the main factor in determining elasticity, one would expect that when incomes fall, expenditure on luxury products would register the sharpest decline. Thus we would conclude that _$^{-}$ should be the highest for developed countries, instead of which, it is actually the lowest. Similarly, if the Group II countries are supplying basic products that account

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8 For Group II countries, _$^{+}$ are barely statistically different from zero at the 10 percent level of significance; however, it is not statistically different from the corresponding elasticity for Group I and developed countries at the 5 percent level of significance. The data continue to support the estimation of demand equations separately for Group I and Group II countries, rather than for all developing countries. The Wald has a p-value of 0.009.

for a low share of the incomes of consumers, their exports would hardly be affected by a decline in world income.

The ordering of _¯ can be better explained as the result of long-term buyer-supplier relationships. Buyers and sellers invest substantially in making these long-term commitments (Egan and Mody 1992). Breaking or abandoning a relationship is therefore expensive. If the long-term relationships are not used, their value depreciates rapidly and new relationships must be rebuilt. Buyers in industrial countries, who generate about two-thirds of the world demand for manufactured exports, have strong relationships with other industrial countries (Hakansson 1987); relationships are weakest in Group II countries, and are easily broken. The marginal supplier is the first to lose an order.

This line of reasoning is closely related to the Roberts and Tybout (1992) analysis of sunk costs in exporting. Those firms that have exported are treated as having incurred the sunk costs, raising the probability of exporting in the current period and in the future. Our findings suggest that in addition to exporters incurring sunk costs, buyers also make sunk investments in the establishment of long-term relationships with their suppliers. Both buyers and sellers, therefore, have an incentive to maintain their relationship, contributing to persistence in export growth. Persistence in trade flows has also been examined in the context of U.S. trade by Krugman and Baldwin (1987), and their interpretation is similar to ours. They argue that export prices react only with long lags to changes in exchange rates and that trade volumes react only with long lags to export prices. They use the "Book of the Month Club" analogy to explain these lags: once buyers subscribe to a particular club, relative price changes do not cause them to abandon the club unless these price changes persist.

A similar pattern of persistence is found in labor markets. "Insiders" in the labor market are experienced workers who are hard to fire because they are more productive, have legal contracts that are expensive to buy, and work as a team. "Outsiders" are costly to identify and train. A consequence of insider power is that wage and employment levels tend to persist. The degree of persistence is not unlimited and if wage differentials between insiders and outsiders become too large, a reshuffling can occur (see Lindbeck and Snower 1988).

Our results show that conventional measures of income elasticity that do not disentangle the asymmetry discussed can lead to erroneous interpretations. Landesmann and Snell (1989), for example, find that the income elasticity of demand for manufactured exports from the United Kingdom has been rising, which they assume implies greater non-price competitiveness. It is curious, however, that during this period (late 1981 and 1982), the U.K.'s share of exports in world trade fell quite precipitously and declined even further in the next five years (Landesmann and Snell 1989, figure 3). Our interpretation is that U.K. exports have performed poorly during global downturns. The arithmetic average of a low elasticity when world income expands and very high elasticity when income falls can be quite large. In this case high income elasticity suggests that the U.K.'s non-price competitiveness fell during the last decade. Our measure of non-price factors, $\_{i}$, shows a clear decline in the case of England.

IV. Telecommunications Penetration: A Proxy for Transactional Quality?

There are no simple physical correlates of the process by which new buyers are drawn to engage in long-term relationships with a supplier. We demonstrate in this section that a country's physical transactional infrastructure (as proxied by its telecommunications infrastructure) is important and accounts partially for the observed fixed-effects. But much of the observed persistence remains after allowing for the influence of telecommunications, lending further credence to the hypothesis postulated in the previous section that transactional quality of a more intangible nature, such as that embedded in long-term buyer-supplier relationships, is needed to make the jump from an outsider to insider status.

The number of telephone lines per capita is used as a proxy for the ability to communicate easily with buyers and to respond rapidly to their requests. Table 5 reports the regressions when the telecommunications variable is added to the demand equation. To examine the possibility of endogeneity, both the contemporaneous growth of telecommunications and growth in the previous year are considered.

For Group I countries, we find the contemporaneous growth rate of telecommunications penetration had no effect on export growth; but telecommunications growth in the previous year did have a positive effect on the growth of manufactured exports. The elasticity of export growth with respect to the previous year's growth in telecommunications is 0.61, which is significantly different from zero at the 5 percent significance level.

When telecommunications availability is introduced as an independent explanatory variable, the fixed effects become somewhat smaller, but remain significant. As before, we use the Wald statistic to test for the significance of the persistence. Recall that when the telecommunications variable is not included as a right-hand side variable, the Wald test strongly suggested the significance of the constant term (or the average persistent growth rate) as well as the significance of the individual fixed effects coefficients. When the telecommunications variable is introduced directly into the demand equation, however, the p-values for both of the Wald tests rise to 0.045. Thus, country-specific fixed effects continue to be important at the 5 percent significance level.

Two observations are relevant in this regard. Table 6 shows that countries with high export persistence have high average growth rates of telecommunications networks. Hence a simple comparison does indicate a correlation between telecommunications and manufactured export growth. The regression results indicate that the partial correlation (after accounting for price and income effects) also holds. However, regardless of the exact specification, the inclusion of telecommunications as an explanatory variable lowers the the magnitude of the fixed effects by, on average, only two percentage points; the Korean, fixed-effect falls, for example, from 11 to 9 percent. Thus, while telecommunications growth is strongly related to export growth in Group I countries, it accounts for only a portion of the persistence.

For Group II countries (with low and negative) fixed effects, table 5 shows that the contemporary telecommunications penetration has a positive effect on export growth and the coefficient is significant at the 10 percent level. A contemporaneous link could well be the result of more rapid exports resulting in greater telecommunications investment. If that is the case, the contemporaneous coefficient will be biased upward. To test for that possibility, we also examine the relationship between the lagged value of telecommunications penetration and export growth. Lagged telecommunications growth appears to have no effect on export growth in this set of countries; the coefficient is negative and insignificant with a t-statistic of -0.35. The evidence is, therefore, very suggestive of the possibility that indeed the causality runs from export growth to telecommunications investment.

To summarize, while the direction of causality in either case remains uncertain, Group I countries have had high export and telecommunications growth rates, whereas Group II countries have experienced low growth in both respects. It is tempting to infer that a certain acceleration in growth, or a "push," in either of the two variables is required to move from the low-growth to the high-growth cycle. Although such a push could be successful in situations where the capability of exporters to respond consistently with internationally marketable goods is high, it could also lead to waste.

Note that the approach adopted in this section is similar to that of Helkie and Hooper (1988) who take the stock of capital in the economy as a proxy for the supply of a larger variety of goods. We focus on just telecommunications because it is likely to be more germane to export activity. Moreover, for the same reason, we prefer to think of telecommunications variable in the demand equation as a measure of transactional quality rather than as a measure of ability to supply an increased variety of goods.

V. Conclusions

Growth rates of manufactured exports from developing countries tend to be persistent, although not immutable. Countries with high persistent growth rates are beneficiaries of a secular shift, or a diffusion, in demand, which reflects their changing status from marginal to long-term suppliers. This descriptive account is consistent with the hypothesis that international buyers invest in long-term relationships to ensure product quality and reliable delivery, and use these relationships to transfer production and marketing knowledge to their suppliers. The resulting network of relationships creates social capital that makes it costly to change buying patterns. Hence the stability in buying relationships is broken mainly when production conditions change significantly. The limited number of countries that are able to change their status from outsiders to insiders experience explosive growth. Once their insider status is established, growth tends to level off. The evidence also shows that countries that appear to have acquired such social capital have further benefited from physical investment in their transactional infrastructure.

The results here do not support the notion that demand for developing country exports is infinitely price elastic; neither do they imply elasticity "pessimism." A price elasticity of about 1 is observed, large enough that policy efforts to change relative prices should have a significant effect on export growth. The paper's main point is that price-related measures, such as devaluation or export subsidies, will have limited effects unless the country is also a reliable supplier and invests in adequate transactional infrastructure. Riedel (1988b) points to Turkey as an example of the benefits of trade liberalization and export subsidies; but it is worth noting that Turkey's exports have also been supported by a massive effort to improve the predictability of suppliers and expand the country's telecommunications network.

Table 1: Export Demand Functions for Developing Countries

	Dependent Variable: Rate of Growth of Export Volume (Δlog E_t)		
	Country-specific _i	γ_i = _j	$\gamma_i = 0$
$\Delta\log (P^X/P^W)_t$	- 0.95 (- 3.62)	- 1.01 (- 3.46)	- 1.08 (- 3.66)
$\Delta\log (P^X/P^W)_{t-1}$	- 0.20 (- 1.42)	- 0.21 (- 1.58)	- 0.12 (- 0.87)
$\Delta\log Y^W{}_t$	3.10 (6.81)	3.10 (6.38)	4.14 (12.45)
Constant (_)	--	0.044 (2.67)	--
No. of observations	253	253	253
D-W statistic	2.06	1.85	1.79
R-bar squared	0.22	0.18	0.16
p-value (Wald-1)	-	0.016	-
p-value (Wald-2)	-	-	0.005

Notes:
1. p-value (Wald-1) = marginal significance level for H_0: _i = _j.
2. p-value (Wald-2) = marginal significance level for H_0: _i = _j =0.
3. t-statistics in parenthesis, based on standard errors robust to general cross-section and time-series heteroskedasticity (White 1980).
4. i and j refer to countries.
5. In all estimations reported instrument set includes: $\Delta E_{t-1,t-2}$, $\Delta Y^W{}_{t,t-1,t-2}$, $\Delta P^W{}_t$, $\Delta(P^X/P^W)_{t-1,\ t-2}$, $\Delta W_{t,\ t-1}$, $\Delta K_{t-1,\ t-2}$. All variables are expressed in logs.

Table 2: Quantitative Importance of Country Specific Fixed Effects

	Country-specific growth rate (pooled) (1)	Actual growth rate (2)	Growth rate without country-specific effects (3) = (2) - (1)	Relative importance of country-specific effect (4) = (1)/(2)	Country-specific growth rate (separate groups) (5)
Brazil	6.6	16.5	9.9	0.40	9.5
Greece	2.1	10.4	8.3	0.20	4.9
Indonesia	16.9	26.3	9.4	0.64	18.5
South Korea	8.9	17.0	8.1	0.52	11.2
Malaysia	10.4	15.4	5.0	0.68	11.9
Philippines	14.3	18.7	4.4	0.76	16.2
Portugal	-0.6	7.3	7.9	-0.08	2.0
Singapore	6.2	12.6	6.4	0.49	8.1
Spain	1.4	9.9	8.5	0.15	4.3
Thailand	6.7	13.7	7.0	0.49	8.7
Turkey	16.2	21.7	5.5	0.75	16.9
Argentine	-0.5	2.8	3.3	-0.18	-1.7
Chile	-0.1	7.5	7.6	-0.02	-1.1
Colombia	-0.25	3.8	6.3	-0.67	-3.9
India	-2.3	3.7	6.0	-0.61	-3.6
Israel	3.3	11.7	8.4	0.28	2.6
Mexico	-3.2	3.1	6.3	-1.04	-4.5
Pakistan	1.3	8.7	7.4	0.15	0.2
Venezuela	3.6	14.4	10.8	0.25	2.7
Yugoslavia	0.6	5.6	5.0	0.11	-0.7

Note: Country-specific growth rates (column 1) are the corresponding fixed-effect terms from the pooled estimation of equation 1 for 20 developing countries in our sample. Country-specific growth rates in Column 5 are the fixed-effects terms from separate estimations for Group I and II countries reported in Table 3 below.

Table 3: Persistence Effects in Different Country Groups

Dependent Variable: Rate of Growth of Export Volume ($\Delta \log E_t$)

	Developing Countries			Developed Countries
	Group I	Group II	All	
Lagged Cointegration Residuals	-0.12 (-1.99)	--	--	--
$\Delta \log (P^X/P^W)_t$	- 0.52 (- 1.63)	- 0.93 (- 2.76)	- 0.95 (- 3.62)	- 0.28 (-3.38)
$\Delta \log (P^X/P^W)_{t-1}$	- 0.29 (- 2.4)	- 0.11 (- 0.5)	- 0.20 (- 1.42)	- 0.16 (1.79)
$\Delta \log Y^W_t$	2.22 (3.8)	3.53 (4.52)	3.10 (6.81)	1.78 (8.67)
No. of observations	141	112	253	169
D-W statistic	1.84	2.14	2.06	2.15
R-bar squared	0.21	0.24	0.22	0.38
Hausman - difference	0.76	0.26	0.91	0.94
Hausman - levels	2×10^{-5}	0.0	0.001	0.022
p-value (Wald-1)	0.009	0.99	0.016	0.22
p-value (Wald-2)	$3*10^{-6}$	0.99	0.005	0.27

Notes:
1. p-value (Wald-1) = marginal significance level for H_0: $\_i = \_j$.
2. p-value (Wald-2) = marginal significance level for H_0: $\_i = \_j =0$.
3. t-statistics in parenthesis, based on standard errors robust to general cross-section and time-series heteroskedasticity (White 1980).

Table 4: Asymmetries in Income Elasticity

Dependent Variable: Rate of Growth of Export Volume ($\Delta\log E_t$)				
	Developing Countries			Developed Countries
	Group I	Group II	All	
Lagged Cointegration Residuals	-0.12 (-1.81)	--	--	--
$\Delta\log (P^x/P^w)_t$	- 0.64 (- 1.92)	- 0.88 (- 2.69)	- 1.02 (- 3.93)	- 0.28 (- 3.31)
$\Delta\log (P^x/P^w)_{t-1}$	- 0.30 (- 2.45)	- 0.10 (- 0.48)	- 0.20 (- 1.52)	- 0.14 (1.55)
$[\Delta\log Y^w{}_t]^+$	1.65 (2.23)	1.84 (1.78)	1.64 (2.65)	1.55 (5.14)
$[\Delta\log Y^w{}_t]^-$	7.08 (2.0)	14.00 (2.97)	12.68 (3.97)	3.65 (2.45)
No. of observations	141	112	253	169
D-W statistic	1.87	2.18	2.12	2.10
R-bar squared	0.21	0.27	0.24	0.38
p-value (Wald-1)	0.006	0.96	0.007	0.17
p-value (Wald-2)	0.0001	0.90	0.0001	0.19
p-value (Wald-3)	0.17	0.023	0.002	0.22

Notes:
1. p-value (Wald-1) = marginal significance level for H_0: $\_{i} = \_{j}$.
2. p-value (Wald-2) = marginal significance level for H_0: $\_{i} = \_{j} = 0$.
3. p-value (Wald-3) = marginal significance level for H_0 : $ß^+ = ß^-$.
4. t-statistics in parenthesis, based on standard errors robust to general cross-section and time-series heteroskedasticity (White 1980).

<u>Table 5</u>: Telecommunications as a Proxy for Service Quality - Developing Countries

Dependent Variable: Rate of Growth of Export Volume (Δlog E_t)				
	Group I		Group II	
Lagged Cointegration Residuals	-0.13 (-1.94)	-0.12 (-1.81)	--	--
Δlog $(P^x/P^w)_t$	- 0.65 (-2.04)	-0.63 (-1.87)	- 0.90 (- 2.52)	- 0.87 (- 2.27)
Δlog $(P^x/P^w)_{t-1}$	- 0.26 (- 2.01)	-0.34 (-2.88)	- 0.15 (- 0.57)	- 0.10 (- 0.4)
Δlog $Y^W{}_t$	2.24 (3.76)	2.33 (3.88)	3.51 (4.15)	3.27 (3.85)
Δlog T_t	0.17 (0.65)	-- --	0.85 (1.63)	-
Δlog T_{t-1}	-- --	0.61 (2.34)	-	- 0.19 (- 0.35)
No. of observations	138	137	100	103
D-W statistic	1.85	1.84	1.97	2.07
R-bar squared	0.21	0.24	0.28	0.23
p-value (Wald-1)	0.044	0.045	.99	.99
p-value (Wald-2)	0.018	0.044	.86	.99

<u>Notes</u>: 1. p-value (Wald-1) = marginal significance level for H_0: $\_{i} = \_{j}$.
2. p-value (Wald-2) = marginal significance level for H_0: $\_{i} = \_{j}$ =0.
3. t-statistics in parenthesis, based on standard errors robust to general cross-section and time-series heteroskedasticity (White 1980).

Table 6: Telecommunications Penetration Rates

	Average 1971-85	1971	1985	Average growth rate	Group
South Korea	9.67	2.10	25.50	0.15	I
Singapore	24.65	6.80	44.24	0.11	I
Turkey	4.21	1.62	9.07	0.10	I
Malaysia	4.66	1.61	9.08	0.09	I
Yugoslavia	8.10	3.59	NA	0.07	II
Thailand	0.96	0.43	NA	0.09	I
Brazil	5.28	2.09	9.30	0.09	I
Mexico	6.20	3.12	NA	0.08	II
Greece	26.96	11.97	41.34	0.07	I
Indonesia	0.33	0.17	0.52	0.07	I
Spain	27.52	13.46	39.59	0.06	I
Israel	30.02	17.40	46.89	0.06	II
India	0.37	0.22	0.57	0.06	II
Philippines	1.20	0.65	NA	0.06	I
Pakistan	0.42	0.79	0.68	0.05	II
Portugal	13.65	8.50	20.48	0.05	I
Colombia	5.85	3.77	8.04	0.04	II
Argentina	9.10	6.81	11.60	0.03	II
Chile	5.03	4.01	6.51	0.03	II
Venezuela	6.13	3.70	8.19	0.02	II
All	9.51	4.64	17.6	0.159	

Note: Number of telephone sets of all kinds per 100 inhabitants.

Appendix A: Definition of Variables and Data Sources

P^x : Unit value index for manufactured exports, World Tables (WT), 1980=100, US dollar based.

P^w : Unit value index for worldwide exports of manufactured goods, GATT: International Trade, 1987-1988, 1980=100.

W : Nominal wage rate, (= RW * GDP deflator),
RW : real earnings per employee in manufacturing, WT, 1980=100;
GDP deflator : WT, 1980=100, local prices.

e : Nominal exchange rate, index, WT, conversion factor (annual average).

E : Exports of manufactured goods, WT, constant 1980 US dollars.

Y^w : World GDP, defined as

$$Y_i^w = \sum_{j=1}^{94} s_i^j Y^j$$

where Y^j is the GDP of the jth partner of the country i, in constant 1980 US dollars, and s^j_i is the average share of partner j in the total exports of i throughout the period:

$$s_i^j = \frac{\sum_t E_t^{ij}}{\sum_j \left(\sum_t E_t^{ij} \right)}$$

and E^{ij} : Manufactured goods exports from ith country to its jth partner, constant 1980 US dollars. Manufacturing sector covers all products from SITC 5 to SITC 8 minus SITC 68. Source: United Nations, Comtrade Database.

K : Capital good imports, constant 1980 US dollars.
Data from Comtrade database (= 71 + 72, SITC) in current prices, US dollars, are deflated by the price index for worldwide capital good imports from UN Statistical Bulletin, 1970-87.

T : Telephone stations (sets) of all kinds per 100 inhabitants. Data series 9.1, from The Yearbook of Common Carrier Telecommunications Statistics, various issues.

Appendix B: Levin-Lin Test for Cointegration in Panel Data

Levin-Lin test is an extension of the ADF test to pooled cross-section time-series data setting. It evaluates the null hypothesis that the variable in question carries a unit-root for each individual (in our case, country) versus the alternative that it is stationary in levels for all individuals. Unlike the standard test statistics for time series data, panel-based unit root test statistic has limiting normal distribution. In addition, it displays super-consistency: "In contrast to results for stationary panel data, the convergence rate of the test statistic is higher with respect to the number of time periods than with respect to the number of individuals in the sample" (Levin and Lin 1993, p. 3).

To test for cointegration, first it has to be determined whether the variables are stationary in levels. If they are, then no further step is required. If, however, the variables follow a first order integrated, I(1), process, then it is necessary to estimate the cointegrating relationship in levels and test whether or not the residuals from this regression are non-stationary. If cointegration residuals are stationary, that is I(0), then the variables involved are cointegrated. Note that this two-step cointegration test is still valid even if one or more of the variables in question are stationary. (See Campbell and Perron 1991.)

The Augmented Dickey-Fuller test for a time series $\{y_t\}$ regresses the first difference Δy_t on the lagged level y_{t-1}, and p-lagged first differences and the appropriate deterministic variables, denoted by DV={1,t}:[9]

$$\Delta y_{it} = \alpha_{oi} + \alpha_{1i} t + \delta_i y_{it-1} + \sum_{k=1}^{p} \beta_{ik} \Delta y_{it-k} + \varepsilon_{it} \qquad \text{(B1)}$$

In its most general form Levin-Lin procedure tests whether the constant term (α_{0i}), the time trend coefficient (α_{1i}) as well as the coefficient on y_{it-1} (δ_i), in the above ARMA process is zero for each individual (country, in our case).

There are four major steps in Levin-Lin unit root test procedure. Depending on the presence of deterministic variables in the cointegrating relationship the final step may require some additional calculations.

1) First the influence of time-specific effects are eliminated by subtracting cross-section averages from each variable y_{it}. The resulting variable is denoted y_{it}.

2) In this step the regression of equation (B1) is divided into two auxiliary regressions and estimated for each individual.

$$\Delta y_{it} = \alpha_{0i}^{1} + \alpha_{1i}^{1} t + \sum_{k=1}^{p} \gamma_{ik} \Delta y_{it-k} + e_{it} \qquad \text{(B1-1)}$$

$$y_{it-1} = \alpha_{0i}^{2} + \alpha_{1i}^{2} + \sum_{k=2}^{p} \phi_{ik} \Delta y_{it-k} + u_{it-1} \qquad \text{(B1-2)}$$

The estimated residuals (innovations) from these regressions, $\hat{e}_{it}$ and $\hat{u}_{it-1}$, are the orthogonalized first differences and orthogonalized lagged levels with respect to p-lagged first differences and deterministic variables.

3) In order to control for heterogeneity across individuals, the two variables, $\hat{e}_{it}$ and $\hat{u}_{it-1}$, are normalized by the standard error of the regression of $\hat{e}_{it}$ on $\hat{u}_{it-1}$ for each individual i. Let's denote the normalized residuals with $\tilde{e}_{it}$ on $\tilde{u}_{i,t-1}$.

4) If there were no deterministic variables in equation (B1), the final step of the Levin-Lin test would be to regress normalized orthogonal first differences of step 3, $\tilde{e}_{it}$, on normalized orthogonal lagged levels, $\tilde{u}_{i,t-1}$, for the panel as a whole.

$$\tilde{e}_{it} = \delta\, \tilde{u}_{i,t-1} + \tilde{\varepsilon}_{it} \qquad \text{(B2)}$$

[9] The original Dickey-Fuller test does not include *p* lagged first differences on the RHS.

Levin-Lin procedure tests whether or not δ in equation (B2) is significantly different from zero. If, based on the t-statistic, this estimated coefficient is significantly different from zero, then the variable y is stationary.

When the estimated cointegration equation contains deterministic variables, such as a constant term and a time trend, Levin-Lin test requires two adjustments to the t-statistic.[10] These adjustments are necessary because, with deterministic variables present the panel unit root t-statistic does not have a standard limiting distribution. First, the ratio of the long-run standard deviation of y_{it}, (the variable after the time-specific effects are removed in the first step) to its short-run standard deviation (which is nothing but the standard deviation of the regression of $\hat{e}_{it}$ on $\hat{u}_{it-1}$ of step 3 is calculated. Under the null hypothesis of a unit root without drift, the long-run variance of y_{it} is equal to the variation of Δy_{it} at zero frequency, which is the weighted sum of sample covariances of $\{\Delta y_{it}\}$. With drift, however, the trend in y_{it} is removed before calculating the long-run variance. It is already mentioned above that the short-run standard deviation is nothing but the regression standard error estimated in the third step. The ratio of the long-run and the short-run standard deviations is first calculated for each individual and then the average is calculated for the whole panel. In the second and the last step of the adjustment of the t-statistic, adjustment factors for the mean and the variance of the t-statistic are obtained through Monte Carlo simulations.

Based on specifiction tests we found that that the ARMA process (equation B1) contains a constant term and a time trend for the variables in levels, but not for the variables in first differences. Consequently, we calculated the mean and variance adjustment factors for the variables in levels only. They are presented in Table B2.[11] Using the mean and variance

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[10] In our tests we used the procedure recommended by Levin and Lin (1993) to select the model specification, which is to start from a maximum number of lagged first diffences and deterministic variables in equation (B1). We reduced the lag-length unless the last additional lagged difference term is significant, kept the deterministic variables if they are significant at the chosen lag-length.

[11] For each variable in the export demand equation and for each country group we obtained different mean and variance adjustment factors, because there is no uniform chosen lag length in equation (B1).

adjustment factors along with the standard deviation ratio, we obtain the t-statistic with a standard normal limiting distribution for each variable in levels.[12] For the variables in first differences, we directly use the t-statistic for δ in equation B2 with no adjustment.

12 Mean and variance adjustment factors, reported in Table B2, are based on 10,000 replications.

<u>Table B1</u>: Levin-Lin Unit Root t-statistic
(marginal significance levels in parenthesis)

	Developing Countries			Developed Countries
	Group I	Group II	All	
Cointegration Residuals	-2.56 (0.01)	-0.55 (0.58)	-1.03 (0.30)	-1.42 (0.15)
log Exports	-0.07 (0.95)	0.26 (0.80)	0.41 (0.68)	-0.38 (0.7)
log relative export price $(P^X/P^W)_t$	0.18 (0.85)	-0.54 (0.59)	-0.17 (0.87)	- 0.23 (0.82)
log World Income	1.51 (0.13)	0.90 (0.37)	1.45 (0.15)	2.96 (0.003)
Δ [log Exports]	-4.15 ($3.0*10^{-5}$)	-4.25 (2.0*10-5)	-6.84 (0.0)	-6.61 (0.0)
Δ [log Relative Export Prices]	-6.48 (0.0)	-8.45 (0.0)	-6.96 (0.0)	-6.84 (0.0)
Δ [log World Income]	-2.78 (0.005)	-2.62 (0.009)	-3.94 ($8.0*10^{-5}$)	-5.60 (0.0)
log Telecom	3.83 (0.0001)	-0.34 (0.74)	--	--
Δ [log Telecom]	-2.578 (0.009)	-3.64 (0.0003)	--	--

Table B2: Monte-Carlo simulated adjustment factors for the mean and standard deviation of the t-statistic when deterministic variables are present

		Developing Countries			Developed Countries
		Group I	Group II	All	
Cointegration Residuals	mean	-0.34749	-0.65774	-0.62982	-0.49464
	variance	1.9734	4.0665	4.7215	2.2412
log Exports..	mean	-0.27997	-0.45881	-0.38968	-0.40804
	variance	7.0099	2.7318	3.9116	3.3821
log relative export price $(P^x/P^w)_t$	mean	-0.35862	-0.5177	-0.39879	-0.4091
	variance	4.2897	2.4334	3.8336	3.6373
log World Income	mean	-0.35318	-0.34194	-0.33732	-0.39757
	variance	4.0732	4.2452	4.7066	3.3705

Note: Based on 10,000 replications for each group and variable.

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