Devaluation, Relative Prices, and International Trade: Evidence from Developing Countries

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Evidence from Developing Countries

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Devaluation is an integral part of adjustment in many developing countries, particularly relied upon by countries facing large external imbalances. A devaluation can only reduce trade imbalances if it translates to a real devaluation and if trade flows respond to relative prices in a significant and predictable manner. However, a recent strand in the empirical trade literature has questioned the existence of a stable relationship between trade flows and its traditional determinants. This paper re-examines the relationship between relative prices and imports and exports in a sample of 12 developing countries. [JEL F11, F14, F31, F32]

DEVALUATIONS HAVE often been used by developing countries to reduce large external imbalances, correct perceived "overvaluations" of the real exchange rate, increase international competitiveness, and promote export growth. The 50 percent devaluation in early 1994 by the CFA franc zone countries stands out as a recent example of such a policy (see Ostry (1994)).[1] However, a devaluation can only accomplish these tasks if, in the first place, it translates into a real devaluation and,

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[1] The CFA franc zone countries are: Benin, Burkina Faso, Cameroon, Central African Republic, Chad, Congo, Côte d'Ivoire, Equatorial Guinea, Gabon, Mali, Niger, Senegal, and Togo. The Comorian Franc was also devalued at that time by 33 percent.

second, if trade flows respond to relative prices in a significant and predictable manner.

With regard to the response of the real exchange rate to a nominal devaluation, the empirical literature appears to agree that in most devaluation episodes the real exchange rate does indeed respond significantly to a nominal exchange rate shock, at least in the short run.[2] Examining the behavior of the real exchange rate in the aftermath of 29 devaluation episodes, Edwards (1989) finds that, in most instances (the real effects in chronic high-inflation countries appear to be much less), there are significant real effects one year after the devaluation; the effects, however, appear to erode completely beyond the third year. Kiguel and Ghei (1993) further conclude that, in low-inflation economies with a tradition of a fixed exchange rate, the real effects of the devaluation may be even longer-lived than was suggested by the earlier work of Kamin (1988) and Edwards (1989).

The next question is whether trade flows systematically respond to the change in relative prices produced by the devaluation and, if so, what is the order of magnitude of the response. The earlier literature that modeled trade in developing countries (see, for instance, Khan (1974), Rittenberg (1986), Bond (1987), and Marquez and McNeilly (1988)) commonly found evidence that relative prices play a significant role in the determination of trade flows, buttressing policies of devaluation as a way to correct trade imbalances. Their evidence often came in the form of significant t-statistics on the relative price variable in static or "long-run" specifications of import demand or export supplies and, hence, calls to mind the work on the inference problems inherent with variables that have unit roots.

More recent empirical work (see Rose (1990) and (1991); and Ostry and Rose (1992)), however, has suggested that, once the time-series properties of the variables are properly taken into account in the estimation, there is little evidence that relative prices have a significant and predictable impact on trade.[3] While Rose (1990) does not model imports and exports separately (as is done in some of the earlier literature), using data for 30 developing countries he finds that changes in the real exchange rate do not have a significant effect on changes in the balance of trade.

[2] This is also the case for industrial countries; see Mussa (1986).

[3] In particular, Rose (1990) finds that there is little evidence of a systematic relationship between changes in the terms of trade and changes in the current account for various developing countries. Changes in the internal terms of trade induced by commercial policies are examined in Ostry and Rose (1992), who find negligible effects on the trade balance, for both developed and developing countries.

The latter conclusion would, of course, imply that a devaluation is likely to be ineffective in its "expenditure-switching" role and, therefore, in achieving its main goals of reducing trade imbalances and stimulating export growth.

In light of such conflicting evidence and policy implications, this paper re-examines the relationship between relative prices and the imports and exports for a sample of developing countries along the lines of the earlier literature on this subject (see Khan (1974), among others). The analysis, however, is conducted in light of the inference problems that arise when the variables used to estimate behavioral relationships are nonstationary. The paper connects the growing literature on estimating relationships among variables that are nonstationary, including the contributions of Engle and Granger (1987); Johansen (1988) and (1991); and Stock and Watson (1989) to a practical policy problem endemic to developing countries. The cointegration approach to estimating preference parameters employed here is found elsewhere in the recent empirical literature (see, for instance Ogaki (1992) and, for an application very similar to ours, Clarida (1994)). This approach provides reliable estimates of the long-run price and income elasticities of developing countries' import demand as well as industrial countries' demand for developing country exports.

By analyzing whether and to what extent the imports and exports of developing countries respond to relative price changes, conclusions can be drawn as to the effectiveness of the often-used devaluation policies. In addition, the empirical results presented in this paper can also be used to evaluate the efficacy of various commercial policies aimed at altering the relative price of traded and nontraded goods.[4]

Section I discusses the theoretical underpinnings of import and export determination in the context of an intertemporal optimizing version of the simple imperfect substitutes model that has dominated this literature (see, for instance, Goldstein and Khan (1985)). In the developing country, utility-maximizing consumers choose between a nontraded domestic good and an imported good. Similarly, in industrial countries, households choose among the domestically produced good and the export of the developing country. Section II first establishes the time-series properties of the variables used in the analysis and then applies the cointegration tests of Johansen (1988) and (1991) to determine if the specifications suggested by theory adequately define the steady-state behavior of imports and exports. With these relationships in hand, Section III presents

[4] On this, see Ostry and Rose (1992).

an estimator (see Stock and Watson (1990)) that is nuisance-parameter free to tackle the issue of whether or not relative prices affect trade flows in a significant and systematic way and to obtain reliable parameter estimates. Section IV pools into regional blocks the time-series data of the various countries in order to highlight some of the stylized facts that characterize trade flows among developing and industrial countries and the distinct patterns that prevail among geographical regions. The final section summarizes the key results and reviews some of the policy implications.

I. A Simple Model of Developing Country Foreign Trade

The modeling of foreign trade relationships has a long history, as illustrated by Goldstein and Khan (1985). In effect, there is a remarkable degree of consensus in the profession on the empirical form of the demand for imports and exports. The standard approach to specifying and estimating trade equations, the model most prevalent in the empirical trade literature, is the "imperfect substitutes model." The central assumption of that model is that neither imports nor exports are perfect substitutes in consumption for domestic nontraded goods. The assumption of imperfect substitutability has found broad empirical support. For instance, Ostry and Reinhart (1992), who estimate the intratemporal elasticity of substitution between traded and nontraded goods for a broad panel of developing countries, find that this parameter is in the 1.0–1.5 range in all the regions considered in the study, implying gross substitutability.[5] Similar results were found for individual developing countries (see Ogaki, Ostry, and Reinhart (1995)). This section describes such an economy.

The simple continuous-time model of a representative utility-maximizing household described below is meant to be illustrative, as it yields representations of the demand for imports that are quite common in the trade literature (see, for example, Khan and Ostry (1992)). The model outlined below describes a small open exchange economy populated with identical agents that possess perfect foresight. These agents have inherited an outstanding stock of internationally traded debt; since there is perfect capital mobility, the residents in this economy take the world

[5] Gross complementarity would imply an intratemporal elasticity of substitution below unity, while perfect substitutability would imply an infinite intratemporal elasticity of substitution. In addition, if imports and exports were perfect substitutes, one would not observe two-way trade, at least in a static framework.

Combining $u(.)$ with the budget constraint, and introducing the costate variable, μ_t, leads to a Hamiltonian of the form

$$\max U = \int_{t=0}^{\infty} [\alpha \ln(h_t) + (1-\alpha)\ln(m_t)]\exp(-\beta t) + \mu_t[q_t + x_t(p^x/p)_t - r_t^* A(p^x/p)_t - h_t - m_t(p^m/p)_t]. \quad (3)$$

The first order conditions yield the familiar relationships between consumption of the home and imported goods that hold at each point in time:

$$h_t = [\alpha/(1-\alpha)]m_t(p^m/p)_t. \quad (4)$$

Equation (4) equates the intratemporal marginal rate of substitution between importables and nontradables to the relevant relative price. Dynamics place consumption of the importable along the optimal path given by the Euler equation:

$$\dot{m}_t = m_t(r_t^* - \beta). \quad (5)$$

Equation (5) is analogous to equation (4), as it relates the marginal rate of substitution between current and future consumption to the relevant intertemporal price, the world real interest rate.

Industrial Countries' Demand for Developing Country Exports

The optimization problem faced by consumers in the industrial countries directly parallels the foregoing simple framework. The representative infinitely-lived household consumes a nontraded and an imported good, denoted as h_{t^*} and x_t, respectively. The imported good is the export of the developing countries. There are endowments of the home good, q_{t^*}, and the export good, m_t (imported by developing countries), which is not domestically consumed. Households in the industrial countries are assumed to be net lenders, who receive interest income; they can consume or accumulate the asset. The representative consumer problem and solution are summarized by equations (6)–(9),

$$\max U = \int_{t=0}^{\infty} [\alpha\ln(h_t^*) + (1-\alpha)\ln(x_t)]\exp(-\beta t), \quad (6)$$

$$s.t.\ \dot{A} = q_t^* + m_t(p^m/p^*)_t + r_t^* A(p^x/p^*)_t - h_t^* - x_t(p^x/p^*)_t, \quad (7)$$

where the preference parameters are assumed to be the same as those of households in developing countries.[10] As before, the home good serves as the numeraire, $(p^x/p^*)_t$ defines the price of the imported good (which is exported by the developing country) to the home good. The total endowment in terms of the home good is defined as $y_t^* = q_t^* + m_t(p^m/p^*)_t$. The first order conditions yield relationships between consumption of the home and imported goods that hold at each point in time:

$$h_t^* = [\alpha/(1-\alpha)]x_t(p^x/p)_t, \tag{8}$$

while dynamics are given by the Euler equation:

$$\dot{x}_t = x_t(r_t^* - \beta). \tag{9}$$

The Steady State

The dynamics of imports in developing and developed countries are given by the Euler equations (equations (5) and (9), respectively). However, our primary interest in the analysis that follows is to employ cointegration analysis to examine the "long-run" steady-state relationships that describe import demand.[11] Market clearing conditions for the home goods markets ($h_t = q_t$ and $h_t^* = q_t^*$) determine the relative prices of the nontraded goods. A steady-state solution requires that the subjective rate of time preference equal the world rate of interest ($\beta = r^*$); the latter ensures that there is no saving (dissaving) in the steady state. We solve the budget constraints (equations (2) and (9)) to obtain an expression that links imports to their price relative to the home good and to permanent income, which in our nonstochastic framework is defined as the endowment of the exportable plus or minus interest incomes. Its log-linear versions for developing and developed countries are given by equations (10) and (11):

$$\log(m_t) = \log\{[x_t - r^*A](p^x/p)_t\} - \log(p^m/p)_t \tag{10}$$

$$\log(x_t) = \log\{[m_t\, p_t^m + r^*Ap_t^x]/p_t^*\} - \log(p^x/p^*)_t. \tag{11}$$

[10] The assumption of identical preferences only serves to simplify the discussion. In the empirical analysis that follows, we provide country specific estimates of the parameters of interest. Only when the data are pooled by region do we restrict the preferences to be the same across countries within each region.

[11] Cointegration analysis has been employed by Ogaki (1992) and by Clarida (1994) to examine the first order condition linking the consumption of imports to consumption of the domestic good, equations (4) and (8), since this relationship must hold at each point in time. Estimation of that first order condition, if cointegration obtains, yields estimates of the intratemporal elasticity of substitution between importable and home goods, assumed to be minus unity for the Cobb-Douglas case.

This is a nonstochastic version of the "long-run" relationship that describes the behavior of imports, and what will be termed a cointegrating equation that is estimated in the following section. The model can be made to accommodate a stochastic element by either reconsidering the maximization problem under uncertainty, or by assuming that some or all the underlying concepts are measured with error.

Thus, a simple theoretical framework provides a number of testable implications. First, it suggests that a scale variable (such as permanent income or wealth) and relative prices are both necessary and sufficient to define the long-run behavior of imports. This would argue against the inclusion of any other variables in an ad hoc manner. Second, it assigns a predictable and well-defined role for relative prices in affecting trade flows. Third, the simple Cobb-Douglas utility function employed predicts that the income and price elasticities are one and minus one, respectively. The sections that follow will test all these hypotheses for a variety of developing countries and for an aggregate "industrial country" bloc.

II. Empirical Analysis

The structural model outlined in the previous section links the steady-state consumption of the imported good to real permanent income and relative import prices. In this simple two-good setting, the relevant deflator for import prices is the price of nontraded goods. However, because of data limitations, the analysis that follows uses imports as a proxy for consumption of importables, and consumer prices and real GDP to proxy the price of nontraded goods and permanent income, respectively. The extent to which these variables imperfectly proxy the underlying concepts introduces a measurement error that is likely to vary across countries and across time. The only assumption that underlies the estimation is that such measurement errors are stationary processes with well-defined variances. Industrial countries' consumption of developing countries' exports similarly depends on permanent income and the relative price of the exportable. The data used are annual and cover the period 1968–92. Details of the data and sources are presented in the Appendix; the countries included in the sample are listed in the tables that follow.

Time-Series Preliminaries

We establish the time-series properties of the relevant variables through the standard unit root tests: the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests (Dickey and Fuller (1981)). These tests deal with a variable, z_t, that admits a simple autoregressive representation,

$$z_t = \rho z_{t-1} + e_t ,$$

where e_t is a random error term drawn from an unchanging and independent distribution. The two tests examine the null hypothesis that the series z_t is I(1), that is, if the absolute value of ρ equals or exceeds 1, versus the alternative hypothesis that the absolute value of $\rho < 1$. Because shocks do not die out, a nonstationary series has no well-defined asymptotic variance. It is simple to show that the unconditional variance of z depends on the variance of e and the coefficient ρ, as in $\sigma_z^2 = \sigma_e^2/(1-\rho)$. So, if $\rho = 1$, the variance of z is unbounded.

Unless otherwise noted, it was found that the null hypothesis of a unit root could not be rejected for the level of the variable but was rejected for the first difference of the variable. In other words, the results suggest (subject to the usual caveats about the low power of the unit root tests) that the variables in question are I(1) processes. In most instances, real imports, real GDP, the ratio of real exports to real GDP, and relative prices are integrated of the same order.

The next task is to determine across variables if these shocks coincide in a way predicted by economic theory. For instance, is a permanent increase in income associated with a permanent increase in imports, as predicted by the simple theoretical model? The strategy is to determine if one or more linear combinations of these variables is drawn from a stationary distribution. If that holds, then the individually integrated variables are said to be cointegrated. If so, then it can be concluded that these variables define steady-state trade relationships and our simple theoretical model finds support in the data.

Cointegration

In much of the earlier literature, estimates of the preference parameters (i.e., the price and income elasticities) were frequently obtained by applying ordinary least squares (OLS) to a specification that was often very similar to the import equation, (10), and to the export equation, (11). These specifications of the demand for imports and exports usually yielded parameters in accordance with the model's priors, the scale variable entering positively while relative prices entered negatively. Most often, the estimates were statistically significant. However, as Granger and Newbold (1974) first showed, two nonstationary variables may appear to have a relationship only because they have similar time-series properties. Indeed, this could be the case here since, as discussed, all the variables of interest are nonstationary. What we set out to do in the next two subsections is to ensure that inference regarding relative prices is not clouded by a spurious element.

A large body of econometric literature (see Banerjee and others (1986)) tells us that even if cointegration obtains, inference problems still remain in OLS estimation. OLS provides consistent, but inefficient, estimates of the true parameters; biases arising from serially correlated errors and simultaneity problems are present and could be quite large for samples as small as those considered here (see Campbell and Perron (1991)). Under such circumstances, standard errors and t-statistics do not provide an adequate measure of statistical significance.

Lack of cointegration is even more problematic, since the OLS estimates are no longer consistent. Further, failure to obtain cointegration may reflect a fundamental misspecification in the model, possibly arising from the omission of one or more variables. Again, *no* valid inferences can be drawn. In either case, such problems raise questions about the findings of the earlier literature.

To re-examine the role of relative prices in light of these developments and assess if the implied theoretical model is capable of describing the data, we proceed in two steps. First, we test for cointegration; this tells us whether the long-run behavior of import demand is adequately specified. Second, in the next subsection we employ an estimator that is free from nuisance parameters and, hence, provides reliable estimates of the price and income elasticities and allows us to test whether relative prices significantly affect trade.[12]

The cointegration test most commonly employed in the literature is that suggested by Engle and Granger (1987). However, a more powerful test that allows for the detection and estimation of the number of cointegrating vectors was developed by Johansen (1988) and (1991) in the context of a vector autoregression model (VAR).[13] This is the test employed here.

In the Johansen (1988) and (1991) procedure, maximum likelihood is applied to an autoregressive representation of the form given by equation (12).

[12] A similar approach was taken by Hoffmaister (1992), who examines the behavior of exports and imports for Costa Rica, and by Milesi-Ferretti (1994), who analyzes these issues for South Korea.

[13] The difference in the power of the rank tests when compared with the Engle-Granger (1987) test is analyzed in Kremers, Ericsson, and Dolado (1992), who present both theoretical and Monte Carlo evidence in favor of the specifications employed in the rank tests.

$$\begin{bmatrix} \Delta m_t \\ \Delta y_t \\ \Delta(p_m/p)_t \end{bmatrix} = \Gamma(L) \begin{bmatrix} \Delta m_{t-1} \\ \Delta y_{t-1} \\ \Delta(p_m/p)_{t-1} \end{bmatrix} + \Pi \begin{bmatrix} m_{t-1} \\ y_{t-1} \\ (p_m/p)_{t-1} \end{bmatrix} + \begin{bmatrix} e_t^m \\ e_t^y \\ e_t^{p_m/p} \end{bmatrix}, \tag{12}$$

where $\Gamma(L)$ is a 3x3 matrix of polynomials in the lag operator, which shifts a series back in time, that is, $Ly_t = y_{t-1}$.

The intuition is as follows: a stationary variable, such as Δm_t, cannot depend on a set of variables that are individually blowing up (such as m_{t-1}, y_{t-1}, etc.). Statistically, this implies that the coefficients on the lagged variables appearing on the right-hand side of equation (12) should be insignificantly different from zero unless that set of variables is cointegrated.[14] That is to say, a linear combination of these I(1) variables produces a stationary process. The lagged first differences of the dependent variables included in the right-hand side ensure that any serial correlation in the residuals is corrected.

These tests thus focus on the properties of the matrix of coefficients, Π. In the absence of cointegration Π is a singular matrix (its rank, $r = 0$). Hence, in our case, the rank of Π could be anywhere between zero, if no cointegrating vector exists, and three, the number of variables in the system. The λ-Max tests the null hypothesis of r cointegrating vectors versus the alternative hypothesis of $r + 1$ cointegrating vectors.[15] If the largest eigenvalue of Π (λ-Max) exceeds the critical value tabulated under the null hypothesis, we can reject the null hypothesis in favor of the alternative. The trace test has the same null hypothesis as the λ-Max test; however, the alternative hypothesis is the rank of Π is $n - r$, where n represents the number of variables in the system. If the trace of Π exceeds the critical value the null hypothesis is rejected.

Tables 1 and 2 present the results of these two tests and their attendant critical values. The null hypothesis tested is that there is no cointegrating vector, $r = 0$.[16] In the case of developing countries' import demand (Table 1), we can reject the hypothesis of no cointegration (using either or both tests) for 10 of the 12 countries in our sample. The results for industrial countries' demand for developing countries' exports (summarized in Table 2) are somewhat less conclusive. The rank tests detect cointegration in 8 of the 12 countries. No cointegrating vector was found for Brazil, Congo, Costa Rica, and Indonesia. In the case of industrial

[14] The relevant joint test is an F-test.

[15] For a concise discussion of these tests, see Campbell and Perron (1991).

[16] The critical values are adjusted using the small sample correction suggested by Cheung and Lai (1993), which is equal to $0.1 + 0.9T/(T - nj)$, where T is the number of observations, n is the number of variables in the system, and j is the number of lags in the VAR.

Table 1. *Testing for Cointegration: Developing Countries' Import Demand, 1970–92*

Country	Maximum likelihood rank tests (null hypothesis $r = 0$)	
	λ-Max	Trace
Africa		
Congo	30.134	41.063
Kenya	23.163	35.527
Morocco	27.592	38.553
Asia		
Hong Kong	29.732	40.552
Indonesia	20.178	29.354
Pakistan	50.414	58.986
Sri Lanka	41.956	62.842
Latin America		
Argentina	25.132	37.457
Brazil	24.036	35.678
Colombia	35.389	47.920
Costa Rica	22.214	34.787
Mexico	24.783	38.985
Critical values for $p - r = 3$		
90 percent	22.32	36.13
95 percent	24.84	39.41

Notes: We test for no cointegrating vector; hence, p, the number of variables, is 3, and r, the number of cointegrating vectors, is 0. The critical values are taken from Osterwald-Lenum (1992) for Case 1* and adjusted using the small sample correction factor suggested by Cheung and Lai (1993).

country demand for developing country exports, the lower incidence of cointegration may simply reflect the fact that for some developing countries the demand for their exports is increasingly coming from other developing countries (see Muscatelli, Stevenson, and Montagna (1994) and Milesi-Ferretti (1994)). This would imply equation (11) is misspecified. In most instances, however, the simple relationships suggested by the theoretical framework seem to find fairly broad support in the data.

III. The Role of Relative Prices: Empirical Evidence

Given the prominence of devaluation in adjustment programs, and particularly since a recent strand in the empirical trade literature has called into question whether relative prices have any effect on trade balances (for instance, see Rose (1990)), our next goal is to assess whether

Table 2. *Testing for Cointegration: Industrial Countries' Demand for Developing Countries' Exports, 1970–92*

Country	Maximum likelihood rank tests (null hypothesis $r = 0$) λ-Max	Trace
Africa		
Congo	22.143	34.273
Kenya	25.981	36.024
Morocco	25.133	37.641
Asia		
Hong Kong	24.985	36.434
Indonesia	19.125	27.541
Pakistan	23.445	36.974
Sri Lanka	25.965	38.397
Latin America		
Argentina	27.325	38.916
Brazil	17.743	23.981
Colombia	29.329	39.912
Costa Rica	18.631	25.983
Mexico	29.180	41.821
Critical values for $p - r = 3$		
90 percent	22.32	36.13
95 percent	24.84	39.41

Notes: We test for no cointegrating vector; hence, p, the number of variables, is 3, and r, the number of cointegrating vectors, is 0. The critical values are taken from Osterwald-Lenum (1992) for Case 1* and adjusted using the small sample correction factor suggested by Cheung and Lai (1993).

and, if so, to what extent trade flows respond to relative prices. Hence, this section focuses on obtaining estimates of the price and income elasticities and tests hypotheses about these parameters.

To that end, we adopt Stock and Watson's (1989) specification, which deals with the biases introduced in the cointegrating regressions by simultaneity and serial correlation in the errors. By eliminating these nuisance parameters, we can obtain reliable estimates of the long-run relationship among these variables. The nonlinear specification reproduced below in equation (13) was estimated for imports and exports, respectively, and the results are summarized in Tables 3 and 4:

$$m_t = \beta_0 + \beta_1 y_t + \beta_2 (p_m/p)_t + \sum_{i=1}^{k} \delta_{1i} \Delta y_{t-i} + \sum_{i=1}^{k} \delta_{2i} \Delta (p_m/p)_{t-i} + e_t \quad (13)$$

In 11 of the 12 countries, relative import prices were significant with the anticipated sign; the exception was Morocco. The price elasticities

Table 3. *Stock and Watson Estimates of Developing Countries' Import Demand, 1970–91*

Country	Constant	p_m/p	y	R^2
Africa				
Congo	−12.42	−0.156	1.359	0.859
	(1.675)	(0.025)	(0.143)	
Kenya	1.960	−0.650	0.095	0.675
	(0.809)	(0.340)	(0.391)	
Morocco	−4.716	0.275	1.204	0.940
	(2.368)	(0.279)	(0.650)	
Asia				
Hong Kong	−1.247	−1.280	1.402	0.985
	(0.623)	(0.362)	(0.049)	
Indonesia[a]	−9.704	−0.927	1.620	0.950
	(1.036)	(0.170)	(0.106)	
Pakistan	−4.046	−0.398	1.150	0.941
	(0.418)	(0.147)	(0.083)	
Sri Lanka	−6.668	−0.304	1.976	0.852
	(0.793)	(0.158)	(0.249)	
Latin America				
Argentina	−1.377	−0.467	1.092	0.404
	(1.380)	(0.147)	(0.583)	
Brazil	13.791	−0.553	2.759	0.850
	(1.364)	(0.147)	(0.320)	
Colombia	1.184	−1.363	1.138	0.901
	(2.520)	(0.537)	(0.121)	
Costa Rica[a]	−0.381	−0.747	0.975	0.519
	(0.279)	(0.263)	(0.257)	
Mexico	−3.360	−0.393	0.893	0.884
	(3.128)	(0.143)	(0.388)	

Notes: Standard errors are in parentheses. Description of the data and their sources are in Table A1.
[a] These results are reported but are not reliable, as cointegration did not obtain.

range from −0.156 to −1.363. Estimates of industrial countries' demand for developing country exports show that prices are significant in seven of the nine countries where cointegration obtains (Kenya and Mexico are exceptions). To examine the robustness of these results, we also tested the significance of relative prices in the context of a VAR, Johansen framework. We compared the unrestricted system given by equation (12) with the restricted version reproduced below in equation (14).

$$\begin{bmatrix} \Delta m_t \\ \Delta y_t \end{bmatrix} = \Gamma(L)\begin{bmatrix} \Delta m_{t-1} \\ \Delta y_{t-1} \end{bmatrix} + \Pi\begin{bmatrix} m_{t-1} \\ y_{t-1} \end{bmatrix} + \begin{bmatrix} e_t^m \\ e_t^y \end{bmatrix} \tag{14}$$

Table 4. *Stock and Watson Estimates of Industrial Countries' Demand for Developing Country Exports, 1970–91*

Country	Constant	p_x/p^*	y^*	R^2
Africa				
Congo[a]	−13.189	−0.320	2.056	0.909
	(1.062)	(0.088)	(0.209)	
Kenya	−5.868	0.188	1.352	0.503
	(8.301)	(0.179)	(1.793)	
Morocco	−8.963	−0.357	1.852	0.957
	(0.767)	(0.103)	(0.164)	
Asia				
Hong Kong	−19.360	−0.544	4.410	0.994
	(0.960)	(0.165)	(0.222)	
Indonesia[a]	−7.201	−0.015	2.022	0.944
	(0.652)	(0.052)	(0.129)	
Pakistan	−7.172	−0.970	1.454	0.935
	(0.897)	(0.244)	(0.209)	
Sri Lanka	−4.937	−0.607	0.889	9.971
	(0.369)	(0.057)	(0.079)	
Latin America				
Argentina	−5.280	−0.415	1.359	0.901
	(1.052)	(0.099)	(0.222)	
Brazil[a]	−9.527	−0.148	2.447	0.940
	(1.011)	(0.157)	(0.221)	
Colombia	−7.254	−0.522	1.626	0.914
	(0.993)	(0.111)	(0.210)	
Costa Rica[a]	−5.189	−0.486	1.071	0.936
	(0.357)	(0.136)	(0.078)	
Mexico	−13.704	0.312	3.379	0.949
	(1.036)	(0.173)	(0.206)	

Notes: Standard errors are in parentheses. Description of the data and their sources are in Table A1.
[a] These results are reported but are not reliable, as cointegration did not obtain.

These results are presented in Table 5. In the case of imports, the χ^2 tests comparing the restricted system (true under the null hypothesis) with the unrestricted system, which includes relative prices, also indicate that for imports the restriction excluding prices was rejected by the data in 8 of the 10 countries where cointegration obtained. For exports, the restricted system was rejected in 7 of the 8 countries where cointegration obtained. The cumulative evidence from the test results for these countries appears to indicate that relative prices play a significant role in the determination of imports, exports, or both.

Hence, in the majority of cases, income and relative prices are sufficient to define a steady state, that is, these variables are cointegrated with

Table 5. *Can Relative Prices Be Excluded? 1970–92*

Country	$\chi^2(1)$ Imports	$\chi^2(1)$ Exports
Africa		
Congo	8.570	12.620
	(0.00)	(0.00)
Kenya	5.910	4.340
	(0.02)	(0.04)
Morocco	1.560	1.840
	(0.21)	(0.08)
Asia		
Hong Kong	2.990	0.230
	(0.08)	(0.64)
Indonesia	10.140	1.520
	(0.00)	(0.22)
Pakistan	36.590	2.720
	(0.00)	(0.10)
Sri Lanka	2.220	12.220
	(0.04)	(0.00)
Latin America		
Argentina	0.500	14.340
	(0.48)	(0.00)
Brazil	8.140	3.290
	(0.00)	(0.07)
Colombia	20.150	2.230
	(0.00)	(0.04)
Costa Rica	8.120	3.360
	(0.00)	(0.07)
Mexico	6.700	12.610
	(0.01)	(0.00)

Note: Probability values are in parentheses.

imports in a way predicted by theory. Second, when the model is compared with a restricted model that excludes relative prices, the data reject this restriction in the majority of cases. Not surprisingly, where the model fares the worst is in its predictions of the income and price elasticities; the joint hypothesis of (1, −1) income and price elasticities was rejected in the overwhelming majority of cases.[17] This result could be due

[17] These results are available from the author upon request. However, the results presented in Tables 3 and 4 already provide estimates of the unrestricted system. In the case of developing countries' import demand, the unit income elasticity could not be rejected for half of the countries. For industrial countries' demand for developing country exports, the unit income elasticity holds in four of the eight cases in which cointegration obtains.

be due to the presence of measurement error in the scale and relative price variables, as these imperfectly proxy the underlying concepts. Just as likely, however, the data could be rejecting the Cobb-Douglas specification in favor of a more general specification. For instance, in the case of a CES utility function, the relative price elasticity will depend on the intratemporal elasticity of substitution, and is not restricted to equal minus unity (see Ostry and Reinhart (1992) who find evidence favoring a CES specification).

In general, the countries in the sample appear to meet the static Marshall-Lerner condition for stability, as changes in relative prices do produce long-run reallocation of trade flows. However, Backus, Kehoe, and Kydland (1992) show in the context of an intertemporal model of international trade that the sign of the relation between the terms of trade and the trade balance will depend on the elasticity of substitution between the imported and home goods rather than in the fulfillment of the static Marshall-Lerner condition. In these models what remains essential is that consumption respond to price changes, a condition for which we find ample empirical evidence.

IV. Regional and Aggregate Evidence

For most of the countries in the sample, only annual data for the variables of interest are available. As such, sample sizes are limited to 25 observations or fewer. The usual small sample handicaps can, however, be circumvented by pooling together countries within a geographical region. Grouping together countries within each region not only increases the efficiency of the estimates of the parameters of interest, but also helps highlight the broader stylized facts that may be obscured in the country-specific analysis. The following section explores these regional trade patterns in greater detail.

The model sketched in Section I assumed, for the sake of simplicity, that preferences were identical across countries, industrial and developing alike. Yet country-specific parameter estimates, as Tables 3 and 4 attest, tend to vary over fairly broad ranges. However, the individual country estimates make it difficult to discern if differences in responses to prices and income follow any broader regional pattern, or if the assumption that developing and industrial countries are alike is met by the data.

To attempt to address these issues we pooled countries within regions and across regions. Several caveats are in order. First, the individual country cointegration tests revealed that, in some of the countries (albeit

Table 6. *Regional and Aggregate Estimates: Fixed Effects Specification, 1970–91*

Country	Relative price	y	R^2
Developing countries' import demand			
Latin America	−0.357	0.964	0.569
	(0.070)	(0.079)	
Asia	−0.403	1.386	0.904
	(0.073)	(0.045)	
Africa	−1.363	1.138	0.901
	(0.537)	(0.121)	
All Countries	−0.531	1.219	0.737
	(0.052)	(0.041)	
Industrial countries' demand for developing country exports			
Latin America	−0.192	2.069	0.780
	(0.084)	(0.115)	
Asia	−0.398	2.494	0.777
	(0.090)	(0.140)	
Africa	−0.266	1.253	0.472
	(0.099)	(0.171)	
All Countries	−0.324	2.052	0.718
	(0.053)	(0.078)	

Notes: Standard errors are in parentheses. The Stock and Watson (1989) estimator was used.

a minority) that are pooled to make up the regional estimates, no cointegration was established. To date, little is known about unit root testing and cointegration tests with panel data. The recent work of Levin and Lim (1993) suggests that there are gains from pooling, as the power of these tests increases. With these caveats in mind, we present the panel estimates in Table 6. These estimates were obtained using the fixed effects estimator, which allows the intercept to vary across countries while imposing the restriction that the slope coefficients are the same and by correcting, along the lines of Stock and Watson (1989), for any potential simultaneity and serial correlation that may be present.

When all developing countries are pooled, and similarly, when the demands for developing country exports are grouped into a single panel, the Houthakker and Magee (1969) results re-emerge. The income elasticity of industrial countries' demand for imports is 2.05, compared with an income elasticity of 1.22 for developing countries' demand for imports.[18] Hence, if developed and developing countries grow at the same

[18] The conditions sketched in the theoretical model have the developing countries as the net debtors and the industrial countries as the net lenders, who receive

rate, industrial countries' trade balance would deteriorate over time. However, this "behavioral" discrepancy, which tends to favor developing countries, does not apply uniformly to all regions. While the Houthakker-Magee result characterizes Asian and Latin American trade patterns quite well, it does not apply to the case of Africa. Industrial countries' demand for African exports has an income elasticity of 1.25, about one half of what it is for Asia and well below the 2.07 for Latin America. Indeed, it is not significantly different from the income elasticity of African import demand. As discussed in Baban and Greene (1992) and Reinhart and Wickham (1994), the high primary commodity content of African exports is a probable explanation for this result.[19] The income elasticity estimates for industrial countries closely resemble those obtained by Clarida (1994), who employs a similar estimation strategy for the United States, and are somewhat lower than those found by Marquez (1989).[20]

Not surprisingly, the panel estimates confirm what the country-specific results showed, namely, that relative prices play a significant role in affecting trade flows. Both industrial and developing countries' demand for imports (irrespective of the region considered) respond to relative prices as predicted by theory. However, with the exception of African import demand, relative price elasticities are well below unity, suggesting that large relative price swings are necessary to produce an appreciable reallocation of trade flows. These estimates are smaller (in absolute value) than Clarida's (1994) estimates for the United States and about in line with the Marquez (1989) estimates.

V. Conclusions

An older empirical literature on trade commonly found evidence that relative prices play a significant role in the determination of trade flows. These results, in turn, lent support to policies of devaluation as a means of correcting trade imbalances and promoting export growth. However,

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interest income. Hence, when GDP is used to proxy permanent income, it introduces systematic biases in the income elasticities. Specifically, since GDP excludes factor payments abroad, it overstates developing country income and understates industrial country income. This systematic error, other things equal, would bias upward the industrial countries' income elasticity, with the opposite being true for developing countries.

[19] For estimation of the demand for specific primary commodities, see World Bank (1994).

[20] Clarida (1994) estimates place the U.S. permanent income elasticity in the 2.12–2.21 range. Marquez (1989), who estimates these parameters for an aggregate of industrial countries, estimates it to be around 2.6.

some of the recent studies that have taken into account the time-series properties of these variables have arrived at a very different conclusion, namely, that no systematic relationship between trade balances and relative prices is discernible from the data.

This paper has re-examined the role of relative prices in affecting trade and therefore, implicitly, the effectiveness of devaluation policies in light of the recent time-series literature that deals with variables that have unit roots and no well-defined limiting distributions. Several empirical regularities emerge. First, the analysis suggests that, in accordance with standard microeconomic theory, income and relative prices are, more often than not, both necessary and sufficient to pin down steady-state trade flows. However, the "traditional" specification appears to fare better when modeling developing-country demand for imports than when applied to industrial-country demand for developing-country exports. The latter may suggest that a fruitful area to investigate is intra-developing-country trade. Second, it is found that, for the majority of cases, relative prices are a significant determinant of the demand for imports and exports. Third, while relative prices have a predictable and systematic impact on trade, price elasticities tend to be low, in most instances well below unity. The latter suggests that large relative price swings are required to have an appreciable impact on trade patterns. Finally, while industrial-country income elasticities are well above their developing-country Asian and Latin American counterparts, suggesting that in a scenario of balanced growth the developing country trade balance should improve, this is not the case for Africa. The high primary commodity content of African exports probably accounts for this result.

APPENDIX

Description and Sources of Data

All data are annual and cover the period 1968–92. The source is the IMF's *World Economic Outlook*.

Variable definitions:

m_t = Nominal imports deflated by import unit values.
y_t = Real gross domestic product (in the domestic currency).
$(p_m/p)_t$ = Import unit values (converted to domestic currency) deflated by consumer prices.
x_t = Nominal exports deflated by export unit values.
y_t^* = Real gross domestic product of industrial countries (in U.S. dollars).
$(p_x/p^*)_t$ = Export unit values deflated by industrial countries' consumer prices (in U.S. dollars).

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