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

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Abstract

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.

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Summary

Developing countries have often resorted to devaluations 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 an example of such a policy. However, a devaluation can accomplish these tasks only if, in the first place, it translates into a real devaluation and, 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, indeed, in most devaluation episodes the real exchange rate responds significantly to the nominal exchange rate shock, at least in the short run. As for the impact of real exchange rate changes on trade flows, an older empirical literature on trade commonly found evidence that relative prices play a significant role in the determination of trade flows. However, 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 re-examines 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" specifications appear to fare better when modeling developing country demand for imports than when applied to industrial country demand for developing country exports. This finding may suggest that a fruitful area to investigate is trade among developing countries. Second, for the majority of cases, relative prices are found to be a significant determinant of the demand for imports and exports. Third, although relative prices have a predictable and systematic impact on trade, price elasticities tend to be low, in most instances well below unity. This result suggests that large relative price swings are required to have an appreciable impact on trade patterns. Finally, 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. However, this result does not hold for Africa, most likely because of the high primary commodity content of African exports.

I. Introduction

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, secondly, 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, indeed, in most devaluation episodes the real exchange rate responds significantly to the 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 relevant issue 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 modelled 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), Rose (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

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1/ The CFA Franc Zone countries are: Cameroon, Central African Republic, Chad, Congo, Côte d'Ivoire, Equatorial Guinea, Gabon, Mali, Niger, Senegal, Togo.

2/ This is also the case for industrial countries, see Mussa (1986).

impact on trade. 1/ 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. 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 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 non-stationary. 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 1990), and Stock and Watson (1990) 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 (1992)). 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. 2/

The next section 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 non-traded 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 III first establishes the time-series properties of the

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1/ 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.
2/ On this, see Ostry and Rose (1992).

variables used in the analysis and then applies the cointegration tests of and 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 IV presents 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 V 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.

II. A Simple Model of Developing Country Foreign Trade

The modelling 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 non- traded 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. 1/ Similar results were found for individual developing countries (see Ogaki, Ostry, and Reinhart (1994)). 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 interest rate as given. As the framework outlined below does not include monetary considerations (i.e., it is a real model), it takes as given, as previously discussed, that a nominal devaluation is capable of altering the relative

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1/ 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, and one would not observe two-way trade, at least in a static framework.

price of traded and non-traded goods. Hence, the households are faced with a "real" shock.

1. Developing countries' import demand

The representative infinitely-lived household consumes a non-traded and an imported good, denoted as h_t and m_t, respectively; they have endowments of the home good, q_t and the export good, x_t, which is not domestically consumed. 1/ Thus, the total endowment expressed in terms of the home good is defined as $y_t = q_t + x_t(p^x/p)_t$. This endowment, in turn, is used to service the interest payment on the fixed stock of outstanding debt, which is denominated in terms of the export, with the remaining income used to finance consumption or accumulate the asset. 2/ The representative consumer maximizes the lifetime utility function,

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

where $\beta > 0$ is the subjective rate of time preference. We assume for simplicity that the household derives utility in each period according to a Cobb-Douglas utility function. 3/

$$s.t.\ \dot{A} = q_t + x_t(p^x/p)_t - r^*A(p^x/p)_t - h_t - m_t(p^m/p)_t, \tag{2}$$

Equation (2) defines the household's flow budget constraint. Taking the home good as the numeraire, $(p^x/p)_t$ defines the price of the export relative

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1/ The assumption that the exportable is not consumed is fairly standard in the trade literature, for a similar framework see Ostry (1988) and Ostry and Reinhart (1992).

2/ None of the results hinge on the household being a net debtor, as the next subsection shows, the analysis carries over to the case where the household is a net lender. The assumption, which follows Reinhart (1991), simply facilitates the link to the developing countries in our sample, most of which are net international debtors.

3/ The empirical evidence, however, tends to favor the more general CES utility function (see Ostry and Reinhart (1992) and Ogaki, Ostry, and Reinhart (1994)), although for many countries the intratemporal elasticity of substitution is not significantly different from unity.

to the home good, and $(p^m/p)_t$ is the relative price of imports. The world interest rate is r^* and A_t represents the stock of debt. 1/

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

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

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

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

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

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

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

2. Industrial countries' demand for developing country exports

The optimization problem faced by consumers in the industrial countries directly parallels the simple foregoing framework. The representative infinitely-lived household consumes a non-traded 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),

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1/ For the sake of simplicity, this study employs a Cobb-Douglas utility function. The empirical implications of this specification will be subsequently tested. The work of Ostry and Reinhart (1992) examines some of these issues in the context of a CES utility function, while Clarida (1992) and Ceglowski (1991) employ an addilog utility function to examine the behavior of imports.

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

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

where the preference parameters are assumed to be the same as those of households in developing countries. 1/ 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:

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

3. 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. 2/ Market clearing conditions for the home goods markets ($h_t=q_t$ and $h_t^*=q_t^*$) determines 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 insures 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

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1/ 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 is pooled by region do we restrict the preferences to be the same across countries within each region.

2/ Cointegration analysis has been employed by Ogaki (1992) and by Clarida (1992) 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.

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 version for developing and developed countries are given by equations (10) and (11):

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

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

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 an 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. Secondly, it assigns a predictable and well-defined role for relative prices in affecting trade flows. Thirdly, 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.

III. 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, due to 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 which 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 is annual and covers the 1968-92 period. Details of the data and sources are presented in Appendix Table A.1; the countries included in the sample are listed in the tables that follow.

1. Time-series preliminaries

We establish the time-series properties of the relevant variables via the standard unit root tests: the Dickey-Fuller (D.F.) and Augmented Dickey-Fuller (A.D.F.) 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 $\sigma^2_z = \sigma^2_e/(1-\rho)$. 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 there are one or more linear combinations of these variables that 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.

2. Cointegration

In much of the earlier literature, estimates of the preference parameters (i.e., the price and income elasticities) were often obtained by applying ordinary least squares (OLS) to a specification which 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 (1973) 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 sections is to ensure that inference regarding relative prices is not clouded by a spurious element.

A large body of econometric literature (see Banerjee et al. (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. Secondly, (next section) 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. 1/

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 1990) in the context of a vector autoregression model (VAR). 2/ This is the test employed here.

In the Johansen (1988 and 1990) procedure Maximum likelihood is applied to an autoregressive representation of the form given by equation (12).

1/ 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.

2/ 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 an 3x3 matrix of polynomials in the lag operator, which shifts a series back in time, i.e., $Ly_t = y_{t-1}$.

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} etc.). Statistically, this implies that the coefficients on the lagged variables appearing on the right-hand-side of (12) should be insignificantly different from zero unless that set of variables is cointegrated. 1/ That is to say, a linear combination of these I(1) variables jointly produce 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 Lambda-Max tests the null hypothesis of r cointegrating vectors versus the alternative hypothesis of r + 1 cointegrating vectors. 2/ If the largest eigenvalue of Π (Lambda-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 Lambda-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 1A and 1B 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. In the case of developing countries' import demand (Table 1A), we can reject the hypothesis of no cointegration (using either or both tests) for all of the twelve countries in our sample. The results for industrial countries' demand for developing countries exports (summarized in Table 1B) are somewhat less conclusive. The rank tests detect cointegration in nine of the 12 countries. No cointegrating vector was found for Brazil, Costa Rica, and Indonesia. In the case of industrial country demand for developing country exports, the lower incidence of cointegration may simply reflect that for some developing countries the

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1/ The relevant joint test is an F-test.
2/ For a concise discussion of these tests see Campbell and Perron (1991),

Table 1A. Testing for Cointegration: Developing Countries' Import Demand 1970-92

Country	Maximum Likelihood Rank Tests (Null Hypothesis r=0) λ Max	Trace
Africa		
Congo	27.431	31.604
Kenya	19.634	29.557
Morocco	21.551	33.853
Asia		
Hong Kong	22.213	34.952
Indonesia	16.578	27.138
Pakistan	48.314	55.281
Sri Lanka	41.956	62.842
Latin America		
Argentina	17.811	35.557
Brazil	19.332	32.771
Colombia	35.389	47.920
Costa Rica	20.283	33.683
Mexico	19.081	36.803
Critical values for p-r=2		
90%	16.85	22.76
95%	18.96	25.32

Notes: We test for at least one cointegrating vector, hence p, the number of variables, is three and r, the number of cointegrating vectors, is one. The critical values are taken from Osterwald-Lenum (1992).

Table 1B. Testing for Cointegration: Industrial Countries' Demand for Developing Countries' Exports 1970-92

Country	Maximum Likelihood Rank Tests (Null Hypothesis r=0)	
	λ Max	Trace
<u>Africa</u>		
Congo	21.103	24.722
Kenya	17.188	28.430
Morocco	19.755	31.213
<u>Asia</u>		
Hong Kong	24.985	36.434
Indonesia	11.824	19.116
Pakistan	16.429	29.173
Sri Lanka	22.267	28.93
<u>Latin America</u>		
Argentina	23.585	32.816
Brazil	10.662	15.767
Colombia	19.419	35.391
Costa Rica	12.511	18.469
Mexico	20.080	27.607
<u>Critical values for p-r=2</u>		
90%	16.85	22.76
95%	18.96	25.32

Notes: We test for at least one cointegrating vector, hence p, the number of variables, is three and r, the number of cointegrating vectors, is one. The critical values are taken from Osterwald-Lenum (1992).

demand for their exports is increasingly coming from other developing countries (see Muscatelli, Steveson, 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.

IV. 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 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 hypothesis about these parameters.

To that end, we adopt Stock and Watson's (1991) 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 2A and 2B.

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

In 11 of the 12 countries, relative import prices were significant with the anticipated sign; the exception was Morocco. The price elasticity ranges from -0.156 to -1.363. Estimates of industrial countries demand 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).

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

These results are presented in Table 3. 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

Table 2A. Stock and Watson Estimates of Developing Countries' Import Demand, 1970-91

Country	Constant	Pm/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	-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	-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 its sources are in Table A.1.

Table 2B. Stock and Watson Estimates of Industrial Countries' Demand for Developing Country Exports, 1970-91

Country	Constant	Px/P	y	R^2
<u>Africa</u>				
Congo	-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)	
<u>Asia</u>				
Hong Kong	-19.360	-0.544	4.410	0.994
	(0.960)	(0.165)	(0.222)	
Indonesia 1/	-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)	
<u>Latin America</u>				
Argentina	-5.280	-0.415	1.359	0.901
	(1.052)	(0.099)	(0.222)	
Brazil 1/	-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 1/	-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 its sources are in Table A.1.

1/ These results are reported but are not reliable, as cointegration did not obtain.

Table 3. Can Relative Prices Be Excluded?
1970-92

Country	$\chi^2(1)$ Imports	$\chi^2(1)$ Exports
Africa		
Congo	8.570 (0.00)	12.620 (0.00)
Kenya	5.910 (0.02)	4.340 (0.04)
Morocco	1.560 (0.21)	1.840 (0.18)
Asia		
Hong Kong	2.990 (0.08)	0.230 (0.64)
Indonesia	10.140 (0.00)	1.520 (0.22)
Pakistan	36.590 (0.00)	2.720 (0.10)
Sri Lanka	2.220 (0.14)	12.220 (0.00)
Latin America		
Argentina	0.500 (0.48)	14.340 (0.00)
Brazil	8.140 (0.00)	3.290 (0.07)
Colombia	20.150 (0.00)	2.230 (0.14)
Costa Rica	8.120 (0.00)	3.360 (0.07)
Mexico	6.700 (0.01)	12.610 (0.00)

Notes: Probability values are in parentheses.

for imports the restriction excluding prices was rejected by the data in 11 of the 12 countries. For exports, the restricted system was rejected in eight of the nine 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 either 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 imports in a way predicted by theory. Secondly, when the model is compared to a restricted model that excludes relative prices, the data rejects 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. 1/ This result could 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), recently 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 responds to price changes, a condition for which we find ample empirical evidence.

V. Regional and Aggregate Evidence

For most of these countries only annual data for the variables of interest are available. As such, sample sizes are limited to 25 observations or less. The usual small sample handicaps, however, can be

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1/ These results are available upon request. However, the results presented in Tables 2A and 2B 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 Malaysia no inference can be drawn). For industrial countries demand for developing country exports, the unit income elasticity holds in four out of the nine cases where cointegration obtains.

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 II assumed, for the sake of simplicity, that preferences were identical across countries, industrial and developing countries alike. Yet country-specific parameter estimates, as Tables 2A and 2B 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 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 (1992) 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 4. 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, along the lines of Stock and Watson (1990), correcting for any potential simultaneity and serial correlation that may be present.

When all developing countries are pooled, and similarly, when the demand 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 to an income elasticity of 1.22 for developing countries' demand for imports. 1/ Hence, if developed and developing countries grow at the same 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

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1/ 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 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.

Table 4. Regional and Aggregate Estimates: Fixed Effects Specification 1970-91

Country	Pm/P	y	R^2
Developing Countries' Import Demand			
Latin America	-0.357 (0.070)	0.964 (0.079)	0.569
Asia	-0.403 (0.073)	1.386 (0.045)	0.904
Africa	-1.363 (0.537)	1.138 (0.121)	0.901
All Countries	-0.531 (0.052)	1.219 (0.041)	0.737
Industrial Countries' Demand for Developing Country Exports			
Latin America	-0.192 (0.084)	2.069 (0.115)	0.780
Asia	-0.398 (0.090)	2.494 (0.140)	0.777
Africa	-0.266 (0.099)	1.253 (0.171)	0.472
All Countries	-0.324 (0.053)	2.052 (0.078)	0.718

Notes: Standard errors are in parentheses. The Stock and Watson (1990) estimator was used.

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 Babban and Greene (1992) and Reinhart and Wickham (1994), the high primary commodity content of African exports is a probable explanation for this result. 1/ The income elasticity estimates for industrial countries closely resemble those obtained by Clarida (1992), who employs a similar estimation strategy for the United States, and are somewhat lower than those found by Marquez (1989). 2/

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 in a way 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 are estimates are smaller (in absolute value) than Clarida's (1992) estimates for the United States and about in line with the Marquez (1989) estimates.

VI. 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, 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 reexamined the role of relative prices in affecting trade, and therefore, implicitly the effectiveness of devaluation policies in light of the recent the 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 appear to fare

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1/ For estimation of the demand for specific primary commodities see World Bank (1994).

2/ Clarida (1992) 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.

better when modelling 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.

Table A.1: Description and Sources of Data

All data is annual and covers the 1968-1992 period. The source is World Economic Outlook (IMF).

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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