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Terms of trade fluctuations and economic growth in developing economies*

Parantap Basu and Darryl McLeod

Fordham University, Bronx, NY 10458, USA

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The effect of terms of trade fluctuations on capital accumulation is investigated in a simple open economy stochastic growth model. Imported inputs make domestic capital more productive, but export prices are uncertain. The model's output process has a random walk component so even transient price shocks have permanent effects on output levels. The size of the random walk component depends on the country's trade share, the supply response of exports and other structural parameters. Also, more variable export prices generally reduce expected domestic investment. These results are consistent with the estimated variance ratios and impulse response functions for a number of LDCs.

1. Introduction

Recent fluctuations in primary commodity prices have renewed interest in the question of how terms of trade movements affect economic growth. This is a familiar if controversial issue in development economics.[1] This paper uses an open economy growth model and some standard time series tests to explore the link between export prices and output growth. We first verify some old and establish some new 'stylized facts' regarding the time series

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[1]In the early 1950s Raul Prebisch, H.W. Singer, and W. Arthur Lewis initiated an enduring debate regarding the effect of raw material price trends on growth in developing countries [see for example the essays by these authors and their discussants in Meier and Seers (1984)]. When international buffer stock schemes were proposed in the 1970s, research shifted toward the welfare consequences of price instability in particular markets [Newberry and Stiglitz (1981) or Adams and Klein (1978) survey these studies]. Most recently, the coincidence of oil price shocks and industrial country business cycles has revived earlier growth and investment concerns [as in Bruno and Sachs (1985)].

stochastic properties, including the appropriate detrending method. This section uses several of these tests to determine the 'size' of the permanent random walk or trend component in LDC terms of trade and GDP series.[5]

Table 1 provides Dickey–Fuller tests for 21 countries' terms of trade over the period 1950–1987. The unit root null can be rejected at the 10% level for eleven of nineteen developing countries. The aggregate non-oil LDC price series also appears to be trend stationary. Rejection of the unit root hypothesis in this instance is both unusual and fortunate. It is unusual because tests of similar price series typically fail to reject [see Nelson and Plosser (1982) or Perron (1989a)]. It is fortunate because the presence of a deterministic trend in primary commodity prices has long been assumed by those debating its direction [see Spraos (1980) or Cuddington and Urzua (1989) for recent reviews]. Since unit root tests lack power, making the random walk the null makes rejection less likely.[6] These results can thus be taken as fairly strong evidence of trend reversion.

Rather than choose between the extremes of trend stationarity and a pure random walk, a more natural characterization of macroeconomic processes might attribute their movement to a *combination* of cyclical and long term forces. Along these lines, Beveridge and Nelson (1981) propose decomposing a series into a stochastic trend which follows a random walk and a stationary cyclical component. One way to gauge the importance of the non-stationary component is to compare its variance to the total for the series itself.[7] Cochrane (1988) provides a convenient nonparametric estimator of the 'size' of the random walk trend component: $1/k$ times the variance of the kth difference. Comparing this variance to that of the first difference yields the variance ratio,

$$V_k=(1/k)[\operatorname{var}(y_t-y_{t-k})/\operatorname{var}(y_t-y_{t-1})]\tau, \tag{2.1}$$

[5]Tests for trend reversion require fairly long time series data which are rarely available for LDCs. Primary commodity prices are easier to find, but there is still the problem of structural breaks and splices in long historical series [Cuddington and Urzua (1989) and the appendix discuss this problem].

[6]These tests have low power when the actual ρ is near 1 and for small samples: see Cochrane (1988). Longer series are available for some table 1 countries. However, as Perron (1989b) emphasizes, long samples also increase the probability of a structural shift in trend or intercept due to major events such as wars, depressions, etc. These structural breaks also create a bias in favor of the unit root. Perron modifies the unit root test to include one exogenous break or crash. We applied his test to table 1 countries for which longer series are available. Mexico (1937), Brazil (1948), Chile (1974) and Venezuela (1973) all appear to have structural breaks which if accounted for using Perron's procedure, lead to rejection of the unit root. Thus, if the break hypothesis is accepted, the number of countries for which trend stationary terms of trade cannot be rejected rises to 15.

[7]Cochrane (1988) shows that this ratio applies to a number of decompositions of this type. The variance ratio may also be interpreted as an indicator of the relative stability of long vs. short run growth rates [see Cogley (1990)]. Cogley also finds the relative stability of the long term trend growth rate emphasized by Cochrane does not hold outside the United States.

Table 1

Terms of trade 1950–1987: Unit root tests and persistence measures.[a]

$$y_t=\mu+\rho y_{t-1}+b(t-T/2)+\sum_{j=1}^{m}\beta_j(y_{t-j}-y_{t-1-j})$$

								Variance ratios		
	m	T	ρ	$t(\rho=1)$	b	t_b	Φ_3[b]	V_4	V_8	V_{12}
Argentina	1	38	0.67	2.5	−0.77	−2.4	3.3	0.51	0.47	0.41
Bolivia	2	37	0.40	3.2*	0.69	1.7	5.6*	0.54	0.45	0.33
Brazil	1	38	0.26	6.3**	−2.68	−6.2	20.3**	0.46	0.35	0.38
Chile	2	37	0.80	2.2	−0.86	−2.5	3.1	0.70	0.90	1.00
Colombia	5	34	0.74	2.1	0.55	1.5	2.2	0.85	0.94	0.80
Costa Rica	1	38	0.42	4.7**	−0.99	−4.6	12.2**	0.70	0.37	0.21
Dominican Rep.	1	38	0.60	2.9	−0.21	−1.3	4.4	0.65	0.39	0.25
Ecuador	1	38	0.51	3.4*	−1.54	−3.0	5.9*	0.58	0.36	0.23
Guatemala	1	38	0.61	3.2*	−0.99	−3.2	5.5	0.75	0.51	0.34
Honduras	1	38	0.45	4.0**	−0.73	−3.9	8.7**	0.53	0.29	0.21
Mexico	5	34	0.50	2.5	−0.60	−2.2	3.3	0.60	0.40	0.35
Panama	1	38	0.70	2.5	0.23	1.4	3.1	0.59	0.59	0.47
Peru	1	38	0.54	3.2*	−0.91	−3.0	5.2	0.29	0.27	0.29
El Salvador	1	38	0.50	3.9**	−0.77	−3.2	8.5**	0.62	0.39	0.21
Uruguay	1	38	0.59	3.2*	−1.53	−3.1	5.3	0.53	0.45	0.42
Venezuela	4	35	0.83	1.9	0.80	1.3	1.8	1.10	1.28	1.51
Philippines	1	38	0.67	2.7	−1.01	−2.6	3.6	0.43	0.45	0.47
Sri Lanka	2	37	0.54	3.4*	−1.53	−3.0	5.8*	0.61	0.39	0.39
India	2	36	0.48	3.4*	−0.12	−0.8	5.7*	0.65	0.45	0.22
Thailand	2	33	0.76	2.2	−0.42	−1.9	2.5	0.62	0.78	0.86
Average for 20 LDCs								0.67	0.55	0.50
IMF non-oil developing countries	1	35	0.48	3.3*	−0.23	−2.2	5.5	0.52	0.53	0.53
								(0.23)	(0.34)	(0.52)[c]
Industrial countries:										
United States	4	35	0.84	2.2	−0.19	−2.2	2.9	1.08	1.30	1.41
								(0.47)	(0.81	(1.1)
IMF industrial countries	4	35	0.82	2.0	−0.06	−1.1	2.2	0.94	1.05	1.21
								(0.41)	(0.65)	(0.97)

[a]All series are logs of price indicies. The IMF non-oil LDC index runs from 1951–1986. See the appendix for data sources.

[b]For the distribution of the statistic Φ_3 (a test of the joint hypothesis $\rho=1$ and $b=0$), see Dickey and Fuller (1981). The augmented Dickey–Fuller procedure was used to correct for serial correlation in ε_t with m chosen from $(0, T^{1/3})$ using the Akaike criteria.

[c]The standard errors shown in parenthesis are $(4k/3T)^{0.5}\hat{V}_k$.

*The null of $\rho=1$ can be rejected at the 10% (*) or 5% level (**).

where $\tau=[T/(T-k+1)]$ adjusts for small sample bias. This ratio approaches one for a pure random walk and zero for a trend stationary series.

The last three columns of table 1 show variance ratio estimates for 4, 8 and 12 year differences. These results generally corroborate the unit root

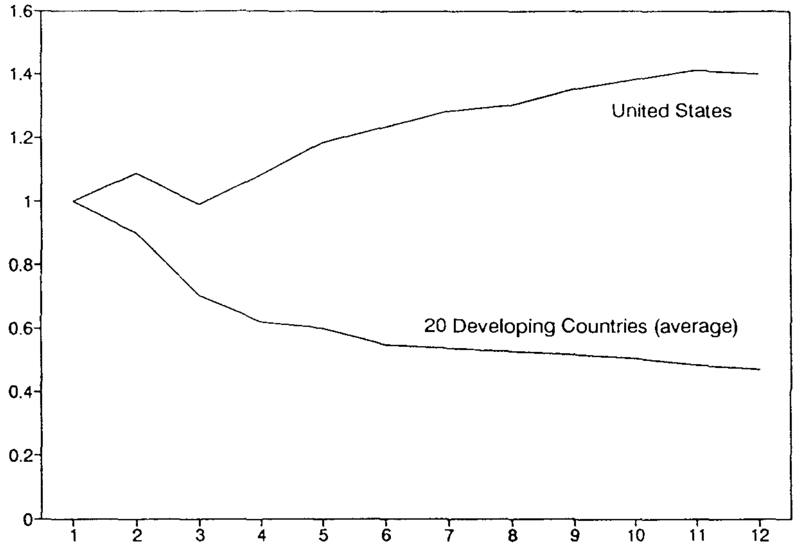

Fig. 1. Terms of trade variance ratios (annual).

tests. For all but four countries (Venezuela, Thailand, Colombia and Chile) variance of the random walk component is less than half that of the series. This contrasts with the higher ratios obtained for the United States. Fig. 1 plots the average V_k for the 20 developing countries along with that of the U.S. A higher variance ratio indicates positive autocorrelation in growth rates. Note that after the initial shock LDC export prices begin to revert to trend almost immediately, whereas the U.S. price innovations are amplified over time. The U.S. ratio settles to about 1.4 as opposed to an average of about 0.6 for the twenty LDCs.[8] The aggregate non-oil developing country index tells the same story: its stochastic trend has only about half the variance of the series.

Table 2 provides the same set of tests for real GDP. Here the pattern is just the reverse of that for prices: LDC output exhibits a larger random walk component and a higher degree of persistence than U.S. GNP. The unit root

[8]If innovations to growth rates are positively correlated the variance of the random walk component may be greater than that of the series so that V_k will exceed one, as it does for the U.S. terms of trade. More intuitively, V_k can be written as a weighted sum of autocorrelations among growth rates ρ_j,

$$V_k = 1 + 2\sum_{j=1}^{k-1} (k-j)/k\rho_j.$$

A positive ρ implies that a jump in today's growth rate signals higher growth tomorrow, so output diverges even more from its previously forecast level. This is the definition of persistence used throughout this paper [see Campbell and Mankiw (1987)].

Table 2

Output (GDP) levels 1950–1987 unit root tests and persistence measures.[a]

$$y_t = \mu + \rho y_{t-1} + b(t - T/2) + \sum_{j=1}^{m} \beta_j(y_{t-j} - y_{t-1-j})$$

	m	T	ρ	$t(\rho=1)$[b]	Φ_3[b]	Variance ratios		
						V_4	V_8	V_{12}
Argentina	1	38	0.92	0.9	1.08	0.82	1.08	1.12
Bolivia	4	35	0.80	3.8**	7.29**	2.19	2.92	2.59
Brazil	4	35	0.81	1.8	1.94	1.86	1.93	1.37
Chile	2	37	0.76	2.4	3.10	1.25	0.62	0.74
Colombia	3	36	0.83	2.1	2.67	1.73	1.72	1.79
Costa Rica	3	36	0.96	0.5	3.47	1.28	1.43	1.16
Dominican Rep.	3	36	0.79	1.5	1.40	0.94	0.87	0.64
Ecuador	4	35	0.86	1.5	1.42	1.47	1.75	1.69
Guatemala	5	34	0.86	1.4	2.28	2.49	2.90	2.95
Honduras	2	37	0.72	2.4	3.04	1.17	0.66	0.47
Mexico	4	35	0.96	0.3	1.43	1.69	1.36	1.52
Panama	3	36	0.89	1.4	2.32	1.52	1.32	1.74
Peru	2	37	0.90	1.4	2.36	1.27	1.22	1.28
El Salvador	2	37	0.95	1.1	1.91	2.64	3.44	2.86
Uruguay	4	35	0.51	2.9	4.34	1.03	0.54	0.36
Venezuela	3	36	0.96	0.8	2.34	2.40	2.92	2.22
Philippines	5	34	0.64	2.5	4.19	2.38	1.79	1.09
Sri Lanka	3	35	0.72	3.2*	7.05*	0.96	1.06	1.08
Thailand	4	34	0.70	3.1	4.95	1.36	1.54	1.23
Average for 19 LDCs						1.64	1.72	1.58
Latin America GDP (ECLA Series)						1.97	1.72	2.52
						(0.54)	(0.92)	(1.12)[c]
United States	1	38	0.71	3.0	5.47	0.95	0.65	0.45
						(0.42)	(0.40)	(0.35)

[a]All series are log *GDP* (*GNP* for the U.S.). See the appendix for for data sources.
[b]See footnote b of table 1.
[c]The standard errors shown in parenthesis are $(4k/3T)^{0.5}\hat{V}_k$.
*The null of $\rho=1$ can be rejected at the 10% (*) or 5% level (**).

can only be rejected for Bolivia.[9] The variance ratios also suggest that long term growth rates are typically more variable than short term rates in LDCs. Fig. 2 compares the variance ratio for 19 developing countries with the well documented U.S. pattern. Whereas U.S. output begins to revert to trend almost immediately, the LDC series has as many as five positive autocorrelations before it settles to about 1.5.

Overall, these tests suggest that developing countries' terms of trade are characterized by substantial short term fluctuations around a more stable long term trend, while their output movements are dominated by long swings

[9]Since the standard errors of V_k are large these estimates need to be corroborated with other evidence, such as unit root tests. Note that the variance ratios for Bolivia and Sri Lanka seem to indicate less trend reversion (in most other countries two tests are consistent).

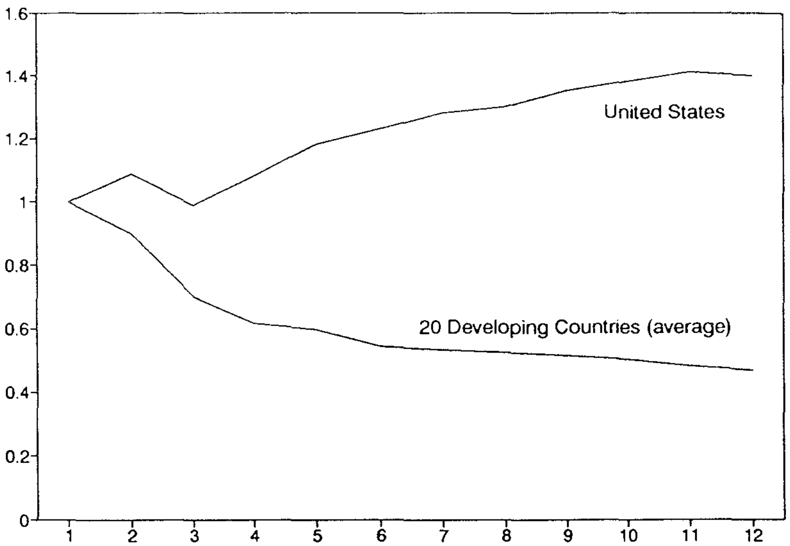

Fig. 1. Terms of trade variance ratios (annual).

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Latin America GDP (ECLA Series)						1.97 (0.54)	1.72 (0.92)	2.52 (1.12)[c]
United States	1	38	0.71	3.0	5.47	0.95 (0.42)	0.65 (0.40)	0.45 (0.35)

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[b]See footnote b of table 1.
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Overall, these tests suggest that developing countries' terms of trade are characterized by substantial short term fluctuations around a more stable long term trend, while their output movements are dominated by long swings

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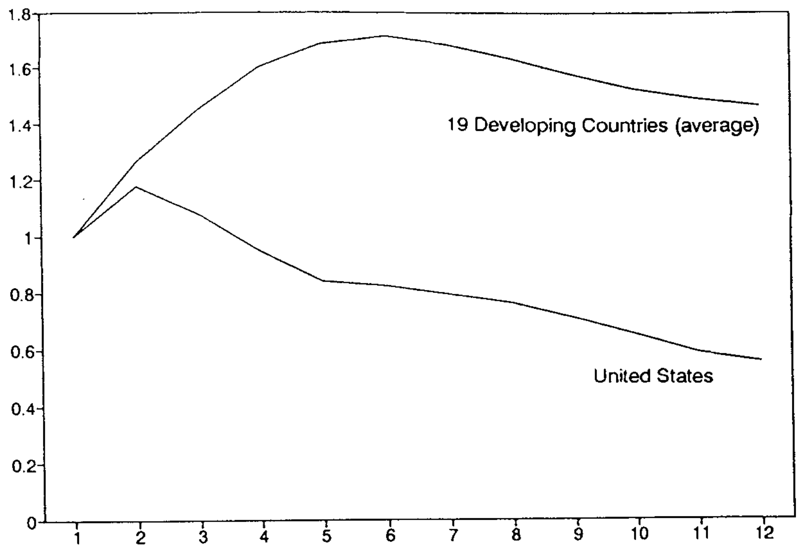

Fig. 2. Variance ratios for output (annual).

in average growth rates. As discussed above, one implication of this asymmetry is that transient terms of trade shocks or 'booms' may have long lasting effects on output levels and average growth rates. This idea can be developed more explicitly in a growth model, a task to which we now turn.

3. Growth with uncertain export revenue

Despite a considerable literature on terms of trade fluctuations, there have been few attempts to examine these issues in a stochastic growth model. The open economy version developed here stresses the role of intermediate imports [as in Bardhan (1970) and Bruno and Sachs (1985)] and a balance of payments constraint (as in the 'two gap' literature). The country's decisions have no effect on export or import prices. Output consists of a single good that can be consumed, invested or exported in exchange for intermediate inputs. An important assumption is that domestic capital and imports are complements: there are constant returns to capital and intermediate inputs together, but diminishing returns to each input separately. This assumption along with the fact that all inputs are produced or are freely available in world markets make the model one of a class of 'endogenous' growth models in which the steady state growth rate depends on the country's propensity to save and in this case, on the probability distribution of world export prices and local supply shocks.

The social planner's problem then is to solve the program,

$$\max \mathrm{E}_0 \sum_{t=0}^{\infty} \beta^t U(C_t), \qquad 0<\beta<1, \tag{3.1}$$

subject to:

$$Y_t=\gamma_t K_{t-1}^{\alpha} v_t^{(1-\alpha)}, \qquad 0<\alpha<1 \quad \text{(production function)} \tag{3.2}$$

$$K_t=I_t+(1-\delta)K_{t-1}, \quad 0<\delta<1 \quad \text{(capital accumulation)} \tag{3.3}$$

$$Y_t=C_t+I_t+X_t \qquad \text{(national product)} \tag{3.4}$$

$$p_t X_t=v_t \qquad \text{(trade balance)} \tag{3.5}$$

$$X_t=\lambda_t K_{t-1} \qquad \text{(export constraint)} \tag{3.6}$$

where $U(C_t)$ is the instantaneous utility function and Y_t is national output to be consumed (C_t), exported (X_t) or invested (I_t).[10] The price of exports, p_t, is uncertain while world import prices are fixed at unity. In line with the evidence just presented, p_t follows an exponential trend subject to serially uncorrelated 'global' shocks ε_t, so that $p_t=\mu^t\varepsilon_t$ (we usually set $\mu=1$). The economy also experiences local supply shocks to the productivity parameter γ_t. Since this is a small economy, it is reasonable to assume ε_t and γ_t are uncorrelated. The absence of serial correlation is a simplifying assumption.

The trade balance eq. (3.5) does not allow for borrowing or accumulation of foreign assets. These complications are examined for a similar model in Basu and McLeod (1990). Given the plausible assumption that terms of trade improvements raise home country wealth, a richer portfolio balance of payments specification does not affect the main results derived below.

Exports may be capacity constrained, requiring specialized capital or time to build (e.g. planting trees). This situation is represented in equation (3.6) where export supply is limited to a share λ_t of the previous period's capital stock. It may seem more natural to think of exports as a share of GDP, but if λ_t is chosen optimally in each period, the solution based on (3.6) is equivalent to choosing the optimal export share in each period (or the optimal level of exports). The supply constrained social planner takes λ as given and chooses an optimal I_t and C_t having observed that year's p_t and γ_t.

In what follows, we contrast the implications of these two export supply

[10]The absence of an exogenous labor supply constraint can be interpreted in the Harrod–Domar tradition of labor abundance or by assuming K includes both physical and human capital, as in the 'AK' growth model of Barro (1990).

specifications. If domestic output is easily transformed into tradables, the country chooses λ_t optimally after observing the local and global shocks to γ_t and p_t. Note the decision sequence. At date t, p_t and γ_t are revealed. In the benchmark flexible export case, the country solves a temporal allocation problem to determine λ_t, setting its marginal revenue product of intermediate input equal to their price $(1/p_t)$. Generally, the choice of I_t and C_t is an intertemporal problem depending on rational expectations of p_{t+1} and γ_{t+1}.

To solve the model use 3.1–3.6 to obtain the resource constraint,

$$C_t + K_t = \psi_t K_{t-1}, \tag{3.7}$$

where $\psi_t = [1 - \delta + \gamma_t(p_t\lambda_t)^{(1-\alpha)} - \lambda_t]$. If exports are predetermined, λ_t is replaced by $\bar{\lambda}$. The growth program for both economies then reduces to the maximization of (3.1) subject to (3.7).

The extra step in the flexible export case is the optimal choice of λ_t. Differentiating ψ_t with respect to λ_t and setting the result to zero yields

$$\lambda_t = [\gamma_t(1-\alpha)]^{1/\alpha} p_t^{v}, \tag{3.8}$$

where $v = (1-\alpha)/\alpha$. Substituting this expression for λ_t into (3.7) yields the resource constraint for the flexible export supply case,

$$C_t + K_t = \phi_t K_{t-1}, \tag{3.9}$$

where $\phi_t = \{1 - \delta + \alpha\gamma_t^{1/\alpha}[(1-\alpha)p_t]^{v}\}$.

We can now solve the intertemporal problem for both economies assuming logarithmic utility.

Proposition 1. If $U(C_t) = \log(C_t)$ the equilibrium process for output is given by eq. (3.10) for the flexible export supply case and by (3.11) if the supply of exports is predetermined as share of K_{t-1}.

$$Y_{t+1}/Y_t = \beta\phi_t(\gamma_{t+1}/\gamma_t)^{1/\alpha}(p_{t+1}/p_t)^{v} \quad \text{(flexible export case)} \tag{3.10}$$

$$Y_{t+1}/Y_t = \beta\bar{\psi}_t(\gamma_{t+1}/\gamma_t)(p_{t+1}/p_t)^{(1-\alpha)} \quad \text{(constrained export case)} \tag{3.11}$$

where $\bar{\psi}_t$ is just ψ_t with λ_t replaced by $\bar{\lambda}$. Proposition 1 follows directly from the optimal accumulation rule for logarithmic utility,[11]

$$K_t = \beta\phi_t K_{t-1} \tag{3.12}$$

[11] The propositions to follow also hold for constant relative risk aversion utility functions as long as the interest elasticity of savings is positive.

for the flexible export case (replace ϕ_t with $\bar{\psi}_t$ to obtain the fixed export share economy). Eq. (3.10) is obtained by eliminating v_t in (3.2) and using (3.8) to replace λ_t. Dividing the resulting expression for Y_{t+1} by Y_t and using (3.12) to eliminate K_t and K_{t-1} yields (3.10). Eq. (3.11) is derived similarly but with $\bar{\lambda}$ replacing λ_t.

In both cases, output follows a (log) difference stationary process driven by a nonlinear combination of price innovations ε_{t+1} and ε_t as well as by the local supply shocks γ_t. A one time export price shock has permanent effects on the *level* of output but only a temporary effect on the growth *rate*. With a positive shock, for example, the growth rate increases the year of the shock and then falls the next year. However, the subsequent fall never offsets the initial increase so the output level and *average* growth rate are left permanently higher (even though the economy returns to its pre-shock growth *rate* two periods later). This random walk component or lack of full trend reversion is described more formally in Proposition 2.

Proposition 2. In both the fixed and flexible export share economies a one time (transient) improvement in export prices permanently raises both the level and the average growth rate of output from the period of the shock forward (compared to its pre-shock growth path and all else constant).

Proposition 2 refers to a single positive price shock. An increase in p_{t+1} raises the growth rate in period $t+1$ and decreases it in period $t+2$, but the initial rise dominates, permanently raising average growth and the level of output.[12] The year of the shock the level of output overshoots its new growth path. This 'boom' is followed by a year of negative growth before the economy returns to its new permanently higher output path. The relative increase in permanent output is greater in the flexible share economy since it can change its export level to better exploit the benefits of temporarily cheap imports. Of course a transient adverse shock reduces the level of GNP relative to its previous growth path.

In general, the more 'open' the economy (that is the greater its trade share $\bar{\lambda}$) the larger is the permanent component of each terms of trade shock. This point is illustrated in fig 3 for a fixed export share economy subject to a single 20% terms of trade shock in period 3.[13] The initial response of output does not depend on $\bar{\lambda}$, but the second year 'trend reverting' movement

[12] To see this let $z_t = (Y_t/Y_{t-1})$ and note from (3.10) and (3.11) that $\partial\log(z_{t+1})/\partial\log(p_{t+1})$ is $v = (1-\alpha)/\alpha$ in the flexible export share and $(1-\alpha)$ in the fixed share economy. The same derivatives in $t+2$ are $-v(1-a_{t+1})$ and $-(1-\alpha)[1-b_{t+1}]$ respectively, where $a_{t+1} = [\phi_{t+1} - 1 + \delta]/\phi_{t+1}$ and $b_{t+1} = [\bar{\psi}_{t+1} - 1 + \delta + \bar{\lambda}]/\bar{\psi}_{t+1}$. Since a_{t+1} and b_{t+1} are evidently less than one, the change in first year growth dominates.

[13] For the simulations in fig. 3 the parameter values for α, γ, δ and β are 0.7, 0.6, 0.1 and 0.95 respectively with $p_t = 1$ except at $t = 3$ when $p_t = 1.2$. In the high and low trade share economies $\bar{\lambda}$ is 0.2 and 0.07 respectively. For the serially correlated shock, $\rho = 0.2$.

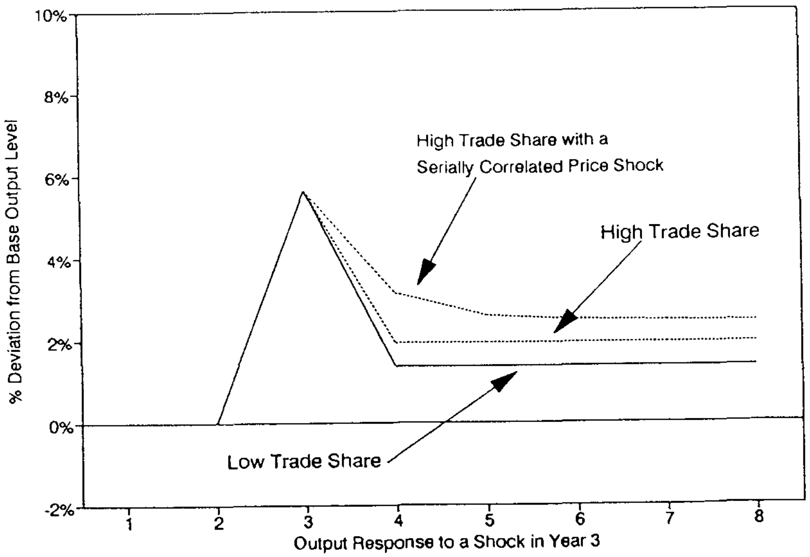

Fig. 3. Impulse response to a 20% price shock (fixed export share economy).

becomes smaller as the trade share, $\bar{\lambda}$, increases.[14] Adding a realistic dose of serial correlation to the price shock creates a smoother impulse response function, not unlike those observed for a number of countries in section 4 below.[15] Hence a combination of infrequent price shocks and a higher degree of openness provides one possible explanation of the higher LDC variance ratios reported in table 2. More closed economies are still affected by external shocks, but short term fluctuations tend to dominate output movements leading to lower variance ratios.

A change in terms of trade uncertainty also affects growth and capital accumulation. Proposition 3 summarizes the consequences of a mean preserving spread in export prices.

Proposition 3. A mean preserving spread of future export prices, p_{t+i}, $i=1,2\ldots$, lowers expected growth rate between t and $t+1$ in both economies, provided $\alpha>0.5$ for the flexible share case. The long-run average growth rate,[16] $E(Y_{t+1}/Y_t)$, may rise or fall, depending on the particular parameters of the economy. The average growth rate of domestic capital, $E(K_{t+1}/K_t)$, is unambiguously reduced by a spread of p in both economies (assuming $\alpha>0.5$).

[14]To see this note that for the fixed export share economy a rise in $\bar{\lambda}$ lowers the offsetting (mean reverting) second year shock but has no effect on the size of the first year shock (assuming $\delta+\bar{\lambda}<1$). Thus the permanent effect of the shock increases. In the flexible share economy λ_t is endogenous but an increase in γ_t, for example, raises both λ_t and the permanent effect of each terms of trade shock (by the same reasoning just applied to $\bar{\lambda}$).

[15]Serially correlated shocks cause no problems in the fixed share economy. In the flexible share case, however, past realizations of ε_t matter for the optimal choice of λ_t, greatly complicating the model's solution.

This proposition can be proved using the general result of Rothschild and Stiglitz (1971) regarding the effect of a mean preserving spread on the expected value of a concave function. Since $E_t(Y_{t+1}/Y_t)$ is strictly concave in p_{t+1} [see eq. (3.11)] a mean preserving spread in p_{t+1} reduces expected GDP growth between t and $t+1$. The same is true for the flexible share economy (3.10), provided the capital share, α, is greater than 0.5 (we assume this condition is met in the discussion which follows).

The long term rate of capital accumulation depends only on the gross marginal product of capital [ϕ or ψ, see eq. (3.12)] which is concave in p in both cases. So a spread in export prices always reduces average investment. Since the marginal product of intermediate imports is strictly convex in p, a greater spread leads to substitution of intermediate imports for domestic capital in the flexible share economy. The increased return to intermediate inputs has a positive while the fall in the expected return to capital has a negative effect on expected output growth. The net change in long run growth depends on the strength of these two opposing effects (see footnote 16). In the fixed export share economy a higher $\bar{\lambda}$ makes it more likely that an increase in price instability will reduce overall growth.[17] Also compared to the flexible share economy, a fall in growth is more likely in the fixed share economy because the optimal λ_t rises with a greater spread of p [see eq. (3.8)]. However, in both cases growth may rise or fall for plausible values of γ,δ and α so the long term consequences of export price instability remains an empirical question. The next section sheds some light on this issue as a one time serially correlated shock to export price variance (modeled as an ARCH process) tends to reduce average GDP growth in the twelve countries we tested.

The results of Proposition 3 can be contrasted with the effect of a spread in the local productivity shock, γ. In the benchmark flexible share case, the average rate of capital accumulation is a *convex* function of γ [see (3.12)], so that greater variability in local supply shocks *increases* average domestic

[16]Here 'long run average' refers to the unconditional expected growth rate. In the flexible share economy this is $\beta E(p^{\nu})E[(1-\delta)p^{-\nu}+\alpha\gamma^{(1/\alpha)}(1-\alpha)^{\nu}]$ (with γ constant). The first expectational term is due to the return to capital, while the second term reflects the marginal product of intermediate inputs. A spread of p evidently moves these expected returns in opposite directions so the stochastic steady state growth rate may rise or fall. The unconditional growth rate for the fixed export share economy is discussed below.

[17]The stochastic steady state growth rate for the fixed share economy with a constant γ is $\beta E(p^{1-\alpha})E[(1-\delta-\bar{\lambda})p^{-(1-\alpha)}+\gamma\bar{\lambda}^{(1-\alpha)}]$ where the second expectational term is again convex in p. A larger $\bar{\lambda}$ reduces the positive contribution of this second term, making it more likely that growth will fall.

investment. This highlights the difference between more variable export prices that lower the expected return to capital and more variable local supply shocks that increase it.

4. Some evidence on terms of trade and growth dynamics

This section provides some empirical evidence on the two main results of the last section: (i) that transient terms of trade shocks have persistent effects on output levels and (ii) that a mean-preserving spread in export prices may lower output growth. Cross section or panel data can be used to test (ii), but time series are required to examine growth patterns and adjustment to shocks.[18] We examine terms of trade and growth dynamics by estimating unconstrained vector autoregressive systems (VARs) for a number of developing countries. Though our main focus is on the relationship between export price shocks and output growth, we also add a time varying variance measure based on the ARCH hypothesis that recent deviations from trend are good predictors of future variability. This variance measure both helps control for changes in price dispersion and provides some indication of how changes in the spread of prices affect growth.

Unfortunately, long price and output series are only available for a limited number of countries. Using a forty-observation minimum leaves just 13 of the countries from table 1. Bolivia was dropped because its impulse response function is sensitive to even small changes in the sample period.[19] The remaining 12 countries fall into three groups. The longest series (from 1928 to 1988) are available for Mexico, Brazil, Argentina and Colombia. In principle this longer sample period should improve the estimates for this group of countries. However, long samples also increase the likelihood of 'structural breaks' due to oil price shocks, world wars, etc. In fact, at least four of the 12 countries considered here do appear to have structural shifts in their terms of trade (see footnote 6). Two steps were taken to deal with the potential bias introduced by these breaks. For Brazil and Mexico, separate 'post break' VARs were estimated. Also, the detrending method of Hodrick and Prescott (1980) was used to compute the variance proxy for all four long sample countries as well as Chile and Venezuela. Of the methods we examined, their method seemed to be the best choice for series with shifting trends or intercepts (for more discussion of this 'HP' filter, see the appendix).

[18]Nearly all of these cross section studies focus on export revenue rather than terms of trade instability. Their findings vary widely, see for example McBean (1966), Lancieri (1978), Kenan and Voivodas (1972) and Glezakos (1973).

[19]Recall that Bolivia was also an outlier in the univariate GDP tests. These problems are probably due to the tin price collapse and Bolivia's subsequent debt crisis. Peru and Ecuador had similar crises, but by using pre-1985 data fairly robust response pattern was obtained. Price shocks have strong short term output effects in Bolivia and Chile (exceptionally large for Bolivia and unusually transient for Chile): this was evident in the growth rate VARs.

For the remaining countries the exponential trend suggested by the univariate tests of section 2 and assumed in the growth model of section 3 was used to compute the variance measure.

Fig. 4 reports the impulse response functions for output following a one standard deviation terms of trade shock (both measured as log levels with terms of trade sample standard deviation in parentheses).[20] Everywhere but Chile, the output response seems to have a substantial 'permanent' component. The share of the initial shock which persists differs among countries, though most initially overshoot their new growth path to some extent. For Brazil and Mexico 1933–1989 there is little evidence of mean reversion even after 12 years. This may be due to the structural breaks in their terms of trade discussed earlier. In fact as shown in panel 1 of fig. 4, their output responses for the post break period (1950–1989) look more like those of the other countries.

The medium sample countries (Venezuela, Chile, Honduras and Ecuador) display a more uniform pattern. Except for Venezuela, growth overshoots and then returns to pre-shock levels after about three years, with about a third of the initial shock persisting after ten years (a profile similar to that reported in fig 3). Only Chile shows signs of continued trend reversion (a result consistent with the variance ratios of table 2). Venezuela moves steadily to its new long term growth path without an initial boom. The results for the short sample countries are similar, with output returning to its long run rate 3–4 years after the shock. Peru is the exception here, as it takes over five years to return to pre-shock growth rates.

Fig. 5 shows the output effect of an increase in the variance or spread of the terms of trade.[21] Note that given the terms of trade level, growth is reduced by a shock to the variance process, with most countries reverting to trend growth after three years. The exceptions are again Mexico, Brazil and Venezuela where the transition takes somewhat longer.

Table 3 provides the variance decomposition for the level of output in each country. Since the standard Choleski factorization was used to orthogonalize the system's innovations, the impulse response and variance decomposition depend to some extent on the order of the variables. Here the appropriate ordering is clearly dictated by the small country assumption: the

[20]Another approach is to detrend and difference each series before running the VAR. Sims, Stock and Watson (1990) argue that this two step procedure is generally unnecessary unless one is testing for causality or intercept levels. Differenced data also tends to accentuate high frequency as opposed to long term interaction among the variables. For these reasons the VARs for levels are reported here. The VARs for growth rates yield similar results except perhaps for Chile and Bolivia (see the previous footnote).

[21]The HP filter was used for all the long sample countries and whenever a break in trend or intercept was evident (i.e. in Chile and Venezuela) – see the appendix for details. Squared deviations from an exponential trend were used for all other countries. This variance proxy is similar to the moving average measure used by Kenan and Rodrik (1986) among others, except that here the weights on past deviations are estimated as part of each VAR.

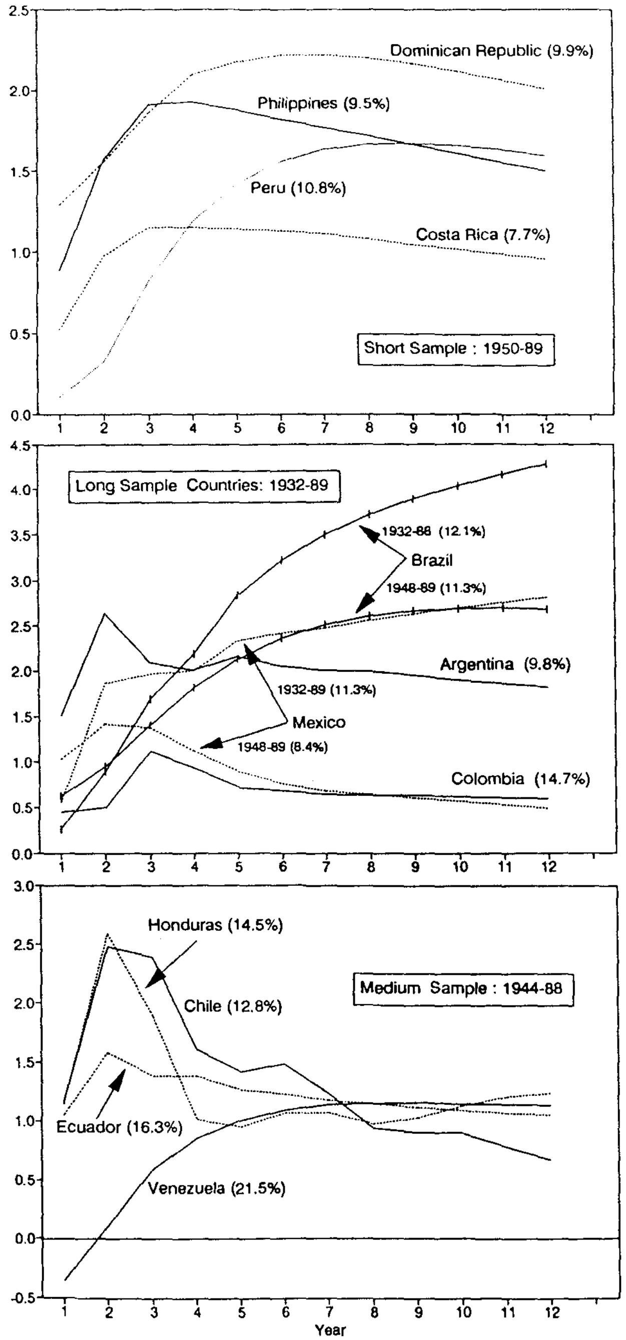

Fig. 4. Output response to a one standard deviation terms of trade shock (% change in GDP from its base level).

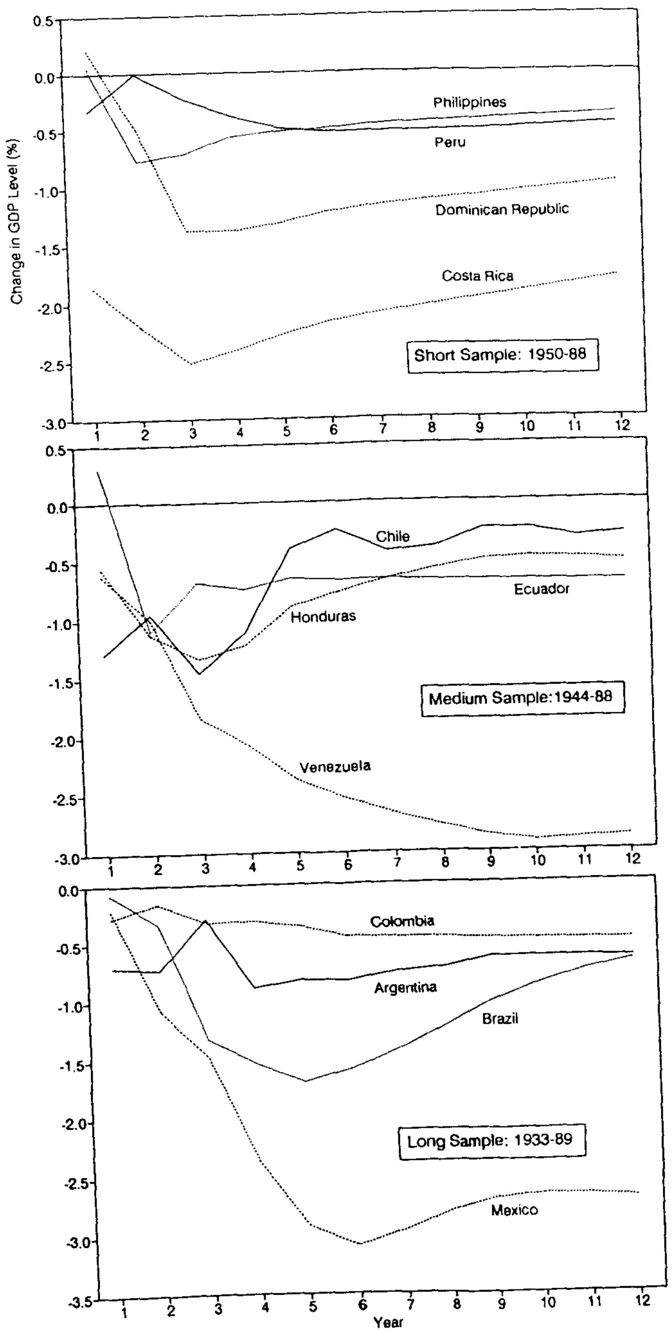

Fig. 5. Output response to an increase in terms of trade instability.

Table 3
Variance decomposition for output.

Year	Long sample countries (1933–89): Std. error	Terms of trade: Level	Terms of trade: Spread	Output	Medium sample countries (1944–89): Std. error	Terms of trade: Level	Terms of trade: Spread	Output	Short sample countries (1950–89): Std. error	Terms of trade: Level	Terms of trade: Spread	Output
	Argentina				*Chile (1947–1989)*				*Philippines*			
1	4	14	3	83	5	6	7	87	3	0	0	100
4	7	32	3	65	10	16	6	78	5	7	5	88
8	10	34	4	61	13	14	4	83	8	20	4	76
12	12	35	4	61	15	12	3	85	9	26	3	71
	Brazil				*Ecuador (1950–1984)*				*Costa Rica*			
1	3	1	0	99	3	11	1	88	4	2	25	73
4	8	12	6	81	7	17	5	77	7	7	39	54
8	14	28	7	65	9	16	5	79	10	9	41	50
12	18	36	5	59	11	15	5	80	11	10	41	49
	Colombia				*Venezuela*				*Peru (1950–1984)*			
1	2	7	2	91	3	1	4	94	3	7	1	92
4	5	12	1	87	7	2	17	81	7	21	1	78
8	7	10	2	88	12	4	24	71	10	24	1	74
12	8	8	3	89	16	4	27	68	12	25	2	73
	Mexico				*Honduras*				*Dominican Republic*			
1	4	2	0	97	3	16	4	80	5	7	0	93
4	9	16	12	72	7	26	10	63	9	9	3	88
8	14	19	23	57	9	20	8	72	14	16	5	79
12	17	22	24	54	11	19	7	75	16	18	5	77

terms of trade level is first, followed by the variance measure and output.[22] Some ambiguity arises in allocating explanatory power between the two price variables. When two variables are correlated, the first tends to be credited with most of the explanatory power. This appears to be true here, as reversing the order often increased the variance proxy's share. When the instability measure is as, or more important than the level even in second position (as in Costa Rica, Venezuela and Mexico), some linear combination of the two price variables explains this share of the variance in output.

These estimates must be interpreted with caution. It is possible that adding other country specific and international variables would affect these impulse response functions. On the other hand, the results are fairly consistent across countries and with the univariate tests of section 2. Everywhere but Colombia, terms of trade movements explain a significant fraction of output fluctuations.[23] And with the exception of Chile and perhaps Bolivia, the profile of the output response is similar to that of the model presented in section 3.

5. Concluding comments

This paper explores the time series properties of a simple growth model in which imports enhance the productivity of domestic capital but the price of exports is uncertain. The level and variability of the export prices (as well as the propensity to save) directly affect the steady state growth rate. This property is consistent with the thinking of some well known development economists, but is not shared by Solow generation models such as Findlay (1980) where terms of trade changes only affect the transition to the steady state. The long run growth rate is 'endogenous' in this model because of a constant returns technology that requires only produced capital and imports. The same structural characteristics that make this an endogenous growth model also introduce a random walk component into the output process. One implication of this stochastic trend is that even transient terms of trade shocks have permanent effects on output levels. There is some reversion to trend since part of the initial shock is reversed in the next period, but output never returns to its previous path. Perhaps most important, the degree of trend reversion depends on the basic stuctural parameters of the economy, including the degree of openness as measured by the trade share of GDP.

[22] A case could be made for putting the variance measure first. If primary commodity prices were determined in an asset pricing model, their mean would depend on their variability as in the ARCH-M model of Engle et al. (1987).

[23] The low explanatory power of Colombia's terms of trade may be due to omitted variables. Adding domestic policy variables to a pooled growth equation for 12 countries including Colombia (as well as Brazil and the Philippines), Edwards (1989) finds a highly significant terms of trade GDP elasticity of about 0.1 (typical of those reported in fig. 4). Using a similar equation, McLeod and Sheehey (1990) report a 1960–88 export price elasticity of about 0.15 for Mexico.

Though long time series data on output levels is scarce, some standard univariate tests indicate that the observed output process in many primary commodity exporting LDCs is consistent with this kind of growth process operating in a stochastic environment dominated by volatile but transitory terms of trade shocks. Of course this is not the only structural model or forcing process which could give rise to these patterns. However, VARs for twelve countries confirm the persistent effects of terms of trade shocks on output levels. They also suggest that greater terms of trade variability reduces economic growth. The long run terms of trade-output elasticities are in the range of 0.1 to 0.2 with terms of trade levels and variability explaining 20–50% of the long term variation in output levels in ten of the twelve LDCs. These estimates are in line with those obtained by Edwards (1989) and others for some of the same countries.

The relationship between export prices and growth in developing countries remains a controversial one. There is disagreement over terms of trade trends, over how export instability affects investment and over the extent to which growth is limited by export earnings. Both the model and time series tests presented here suggest that terms of trade trends and variability can have significant effects on growth and investment in small open economies. While our results certainly do not resolve the controversies in this area, they do help shift debate to firmer methodological ground and provide some new testable hypotheses regarding the link between terms of trade, fluctuations and economic growth.

Appendix: Data sources and estimation methods

The terms of trade and output data used for the unit root tests of section 2 and the VARs of section 4 were compiled from several sources. The industrial countries, India, Philippines, Thailand, Sri Lanka, the Non-oil Developing and Industrial country aggregates are from the IFS (series 74 and 75). Most of the long terms of trade series were obtained from CEPAL (1976) 'America Latina: Relacion de Precios Del Intercambio' *Cuadernos de La Cepal No. 1,* Las Naciones Unidas, Santiago, Chile and various issues of the *Economic Survey of Latin America* also published by U.N. Economic Commission on Latin America. These series cover trade in goods only and end in 1965. Data for 1965–88 is from 'El Balance de Pagos de America Latina, 1950–77' (*Cuadernos de La Cepal No. 5*) and various issues of the *Economic Survey of Latin America* and the *Economic Panorama of Latin America*, 1989 (ECLAC, United Nations, NY). An exception was Mexico where the Banco de Mexico's series was used for 1970–88. The aggregate terms of trade for non-oil developing countries was computed as the ratio of IFS export price index (series 74) for non-oil developing countries divided by

the export price index for the industrial countries the alternative IFS non-oil developing country import index begins in 1960).

The GDP series for the Industrial and Asian countries came from the IFS yearbooks, various editions while the longer historical series for most of the Latin American countries (except Mexico) came from CEPAL (1978) 'Series Historicas del Crecimiento de America Latina' *Cuadernos de La Cepal No. 3* up until 1965. After 1965 the World Bank series were used through 1987: 1988 growth rates came mainly from the IFS. All of these series are available on diskette from the authors upon request.

The VARs of section 3 include three variables: log levels of terms of trade and output along with time varying price variance measure (squared deviations from trend). This last variable predicts future terms of trade variability in accordance with the 'ARCH' hypothesis that large deviations in either direction tend to be followed by more fluctuations. A lag length of three was used for the long sample countries and two for the medium and short sample groups, except for Colombia (four lags). A problem in constucting the variance measure for some of the long price series was how to detrend a series which evidently has 'structural breaks' in its trend or intercept [Cuddington and Urzua (1989) also encounter this problem with a long primary commodity price series]. Three methods present themselves. One is to ignore the break in the series and simply detrend it by differencing or fitting an exponential trend. This approach has the obvious problem of introducing a very large shock when the intercept jumps and/or persistent correlated errors around the change in trend or drift. A second method is to fit a trend which allows for a one time shift or intercept. This assumes that agents immediately recognize a particular shock as being entirely permanent and/or instantly revise their expectations regarding the underlying trend or drift the year the shift occurs. A third method is the more pragmatic Hodrick–Prescott (HP) filter which simply minimizes the sum of squared deviations while penalizing changes in slope. This slow changing trend seems to provide the best compromise between ignoring structural breaks and assuming an artificially rapid revision of expectations the year of the break. The HP filter was used whenever a structural break in the series was suspected.

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