ELSEVIER

Journal of Development Economics
Vol. 54 (1997) 323–356

JOURNAL OF
Development
ECONOMICS

Terms-of-trade uncertainty and economic growth

Enrique G. Mendoza *

Department of Economics, Duke University, Durham, NC 27708, USA

Received 4 September 1994; accepted 18 April 1996

Abstract

This paper examines a stochastic endogenous growth model in which terms-of-trade uncertainty affects savings and growth. The model explains the well-known positive link between growth and the mean rate of change of terms of trade, and predicts also that terms-of-trade variability affects growth. Increased terms-of-trade variability results in faster or slower growth depending on the degree of risk aversion, but in either case it reduces social welfare. These growth effects imply that welfare costs of macroeconomic uncertainty are much larger than first thought. Cross-country panel regressions provide strong support for the model's key predictions. © 1997 Elsevier Science B.V.

JEL classification: F41; F43

Keywords: Terms of trade; Savings; Growth; Volatility; Uncertainty

1. Introduction

Many recent empirical studies in growth theory have examined the nature of cross-country growth differentials and the economic forces that explain them. [1] These studies have produced mixed statistical evidence on the contribution of macroeconomic policies or country characteristics to explain cross-country differ-

* Corresponding author. Tel.: + 1-919-490-9037; fax: + 1-919-684-8974, e-mail: mendozae@econ.duke.edu

[1] See, e.g., Barro (1991), Barro and Lee (1993), Barro and Sala-i-Martin (1995), Razin and Yuen (1994), Easterly and Rebelo (1993), Fischer (1993), and Easterly et al. (1993).

PII S0304-3878(97)00046-1

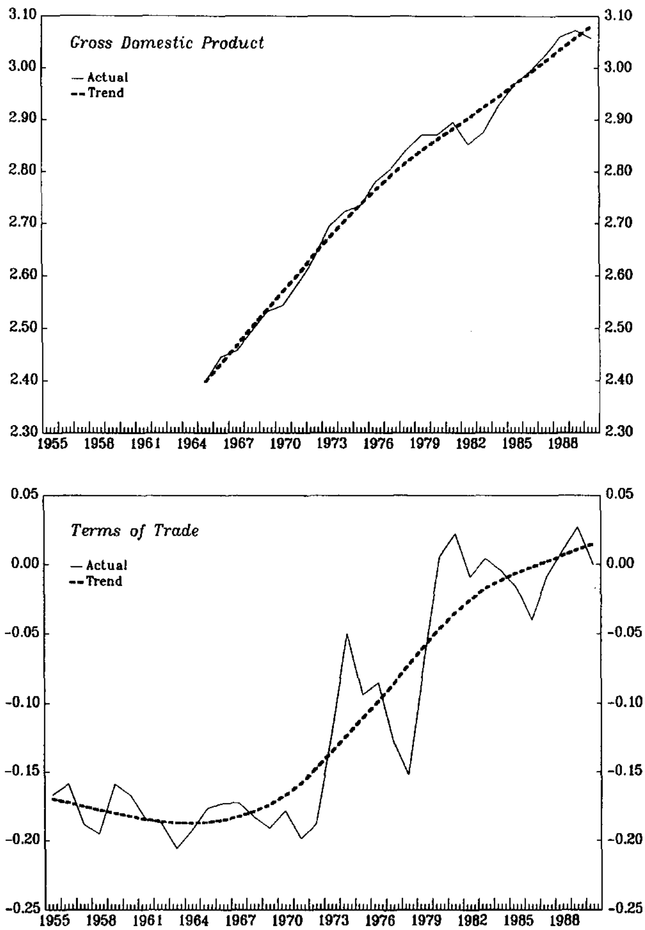


Fig. 1. Canada: GDP and terms of trade.

ences in average growth rates. The results of these studies show that, while policy indicators and country characteristics have a variable degree of significance and robustness in panel regressions, the terms of trade are typically a significant and robust determinant of economic growth. A simple visual analysis of the experience of two industrial and developing countries chosen at random (Canada and Kenya) illustrates dramatically the close relationship between sustained economic growth and rising terms of trade (see Figs. 1 and 2). [2] Canada displays sustained growth,

[2] The trends of per-capita GDP and terms of trade in the charts were produced using the Hodrick–Prescott filter with the smoothing parameter set at 100. See Section 3 for details in data sources and transformations.

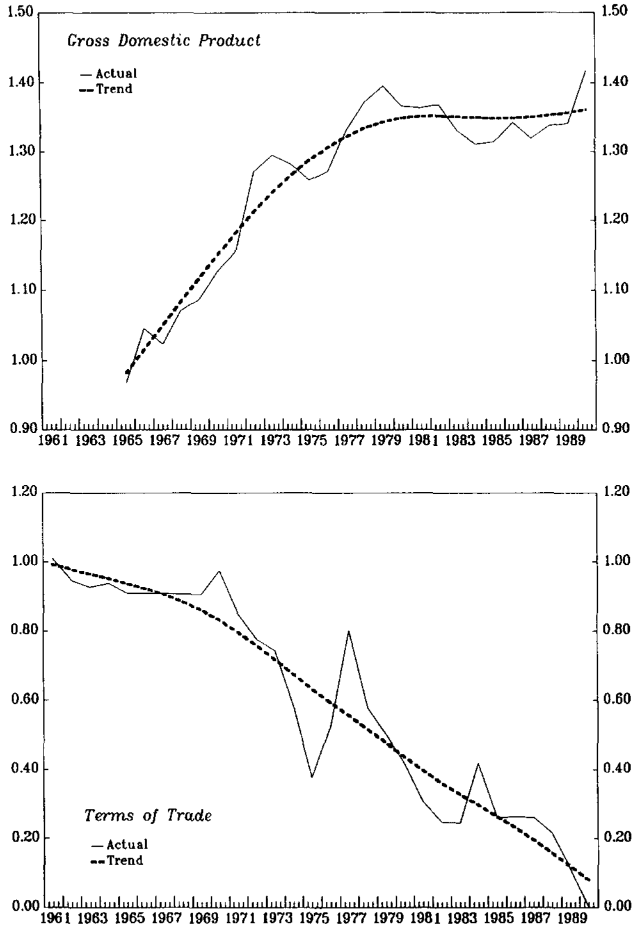


Fig. 2. Kenya: GDP and terms of trade.

with moderate deviations from trend, coinciding with a marked upward trend in the relative price of her exports, while in Kenya the opposite phenomena are observed. The opposing trends of the terms of trade for these countries reflect to a large extent the protracted and severe decline of real commodity prices over the last two decades (see Reinhart and Wickham, 1994), and hence are typical of comparisons across industrial and developing countries

In line with the above arguments, Easterly et al. (1993) find that economic policies and country characteristics, such as educational attainment and political stability, contribute little to explain the observed lack of persistence in growth performance, while terms-of-trade changes are highly correlated with growth, particularly during the 1980s. [3] Barro and Sala-i-Martin (1995) and Fischer (1993) find that country characteristics do contribute to explain growth differentials, but the terms of trade still play a key role. Barro and Sala-i-Martin's results show that the growth effects of terms of trade compare to those of educational attainment, public spending on education, human capital, and political instability. Fischer shows in addition that the terms of trade are more important at higher frequencies, since they are more significant in pooled than in between-means regressions. [4]

While the positive relationship between terms of trade and growth has been clearly identified, most empirical work does not conduct structural tests tied to a theoretical framework in which the role of terms of trade is explicitly modelled. The explanatory power of country characteristics and policy variables is related to existing growth models, whereas the terms of trade are viewed as an exogenous variable with a somewhat uninteresting role. Moreover, growth effects that could result from uncertainty and risk aversion in the presence of terms-of-trade fluctuations. as well as uncertainty with regard to other growth determinants, are generally ignored.

This paper attempts to shed some light on these issues by studying a model of savings under uncertainty that provides an interpretation for the observed positive relationship between *average* rates of change of terms of trade and average consumption growth rates. [5] In addition, the model highlights the importance of the *variability* of terms of trade as a determinant of average growth rates, and thus illustrates potentially important growth effects of uncertainty and risk. These effects in turn have important implications for the welfare costs of macroeconomic uncertainty.

The main argument that the paper makes with regard to growth effects of terms-of-trade uncertainty is a feature of growth models of varying complexity, but it remains a feature even of simple stochastic growth models. Thus, for clarity and technical simplicity, the analysis is based on a model of a small open economy that extends the savings-under-uncertainty framework of Phelps (1962) and Levhari and Srinivasan (1969). The model can be interpreted also as a stochastic version of the one-sector endogenous growth model of Rebelo (1991). In the model, the mean and variance of the terms of trade determine the savings rate and

[3] Pooled regressions of Easterly et al. (1993) show that if the terms of trade gain as a share of GDP rises by 1% per annum, annual growth rises by 0.42% in the 1970s and by 0.85% in the 1980s.

[4] This explains why a smoothed measure of terms of trade fails to explain growth in the panels of country means of de Gregorio (1992), and suggest that the variance of terms of trade may be important for explaining growth.

[5] There is a large literature on trade theory that examines the link between terms of trade and growth, as reviewed by Findlay (1984). This paper differs from the trade literature in that the link between the two variables follows from the effects of uncertainty on savings in a neoclassical setting, not from market imperfections (see, e.g., Lewis, 1954; Prebisch, 1950; Singer, 1950; and Findlay, 1980). Market imperfections are likely to be relevant for explaining growth differentials but it is also useful to determine first to what extent market forces explain the facts.

consumption growth. Growth is slower in economies in which terms of trade grow at slower rate, on average, because slow terms-of-trade growth reduces the expected real rate of return on savings—in units of imported goods—and this affects the savings rate. The variability of terms of trade also affects the savings rate and growth, with an effect that is positive or negative depending on the degree of risk aversion. If risk aversion is low, increased variability in the terms of trade, measured as a mean-preserving spread, reduces both growth and social welfare. If risk aversion is high, increased terms-of-trade variability produces faster grow but it still reduces welfare. Thus, the model predicts that the variance of the terms of trade contributes to explain growth, and hence suggests that cross-country growth regressions that include only the mean of the rate of change of terms of trade are misspecified.

Given the growth effects of terms-of-trade uncertainty, the welfare gains that result from reducing consumption instability are much larger than the negligible gains estimated by Lucas (1987). Lucas's estimates, as well as similar results obtained in international real-business-cycle models (see, e.g., Backus et al., 1992; Cole and Obstfeld, 1991; and Mendoza, 1991), abstract from the possibility that altering the variance of consumption may not only affect the amplitude of consumption fluctuations around trend, but the level of that trend as well. In the savings-under-uncertainty framework, risk-averse agents adjust their savings rate, and hence the trend level of consumption, in response to changes in the variability of the underlying process driving consumption fluctuations. A similar proposition is examined in the context of the assessment of the welfare gains of international risk-sharing by Obstfeld (1994). When international diversification affects growth, the gains from risk-sharing are significantly larger than those produced by models of business cycles around exogenous trends (see, e.g., Tesar, 1995).

The paper examines the empirical relevance of the model's key predictions using a multi-country database including 40 industrial and developing countries and using panel estimation methods. The model's closed-form solutions establish two hypotheses regarding consumption growth in competitive equilibrium: (a) a time-series linear relationship between the rate of change of terms of trade and consumption growth, and (b) a cross-section relationship such that country average growth rates depend on the mean and variance of the rate of change of terms of trade. These hypotheses are strongly supported by the data. In particular, there is a large adverse effect of terms-of-trade variability on economic growth, and the information captured by the variance of the terms of trade for explaining growth is not captured by the mean. These results are robust to the addition of the explanatory variables typically examined in the recent empirical growth literature (see Barro and Sala-i-Martin, 1995).

The paper is organized as follows. Section 2 presents the model and conducts numerical simulations to illustrate its welfare implications. Section 3 examines cross-country regularities of growth and terms of trade, and conducts econometric tests of the model's closed-form solutions. Section 4 provides some conclusions.

2. The terms of trade and growth: a basic framework

This section examines a basic neoclassical model that explains the relationship between terms of trade and growth under uncertainty. The model extends the savings-under-uncertainty framework developed by Phelps (1962) and Levhari and Srinivasan (1969) to the case of a small open economy facing terms-of-trade shocks. The model can also be interpreted as a one-sector, stochastic variant of the endogenous growth model developed by Rebelo (1991). [6]

The economy is inhabited by households that formulate optimal plans for consumption of an imported good so as to maximize expected lifetime utility:

$$U(C) = E\left[\sum_{t=0}^{\infty} \beta^t \frac{C_t^{1-\gamma}}{1-\gamma}\right]$$

$$\gamma > 0,\ 0 < \beta < 1,$$

where C_t is consumption of the imported good, β is the subjective discount factor, and γ is the coefficient of relative risk aversion (i.e., $1/\gamma$ is the intertemporal elasticity of substitution).

The production technology adopts the simple form of a perfectly durable asset, or a linear technology, that yields a stochastic return each period. [7] The return is an exportable commodity that agents exchange for imports of consumer goods in a perfectly competitive world market, where there is a large number of market participants. The relative price of imported goods in terms of exports is also subject to random disturbances. Markets of contingent claims are incomplete, and hence households cannot insure away country-specific income shocks resulting from changes in real returns or the terms of trade. Thus, households maximize utility subject to the following period-by-period resource constraint:

$$A_{t-1} \leq R_t(A_t - p_t C_t). \tag{2}$$

Given $A_0 > 0$. A_t is the stock of wealth in units of exportables, R_t is the domestic gross rate of return on savings, and p_t is the relative price of imports in terms of exports determined in world markets, or the reciprocal of the terms of trade $\mathrm{tot}_t = p_t^{-1}$. R_t and p_t are non-negative random variables such that the effective rate of return $r_t = R_t p_t / p_{t+1}$ follows a log-normal i.i.d. distribution. Thus, $\ln(r_t)$ is an i.i.d. probabilistic process with mean μ and variance σ^2, and hence the mean and variance of the r_t process are $\mu_r = \exp(\mu + \sigma^2/2)$ and $\sigma_r^2 =$

[6] Hopenhayn and Muniagurria (1993) and Obstfeld (1994) examine two-sector extensions with linear technologies.

[7] Although the paper examines a simple representation of the model, the same basic growth equation results if one allows for investment and depreciation in continuous time (see Rebelo, 1991; and Hopenhayn and Muniagurria, 1993). International borrowing and lending in one period bonds, yielding the same return as the domestic linear technology, can also be added without altering the results.

$\mu_r^2(\exp(\sigma^2 - 1))$. At each date t, p_t is known but R_t and p_{t+1} are unknown. The competitive equilibrium is defined by optimal intertemporal consumption allocations that maximize Eq. (1) subject to Eq. (2). [8] The optimality conditions are the constraint and the Euler equation:

$$U'(C_t) = \beta E\left[\frac{R_t p_t}{p_{t+1}} U'(C_{t-1})\right]. \tag{3}$$

Closed-form solutions for this model are obtained using dynamic programming techniques. The solutions are:

$$C_t^* = \lambda\left(\frac{A_t}{p_t}\right), \tag{4}$$

$$A_{t+1}^* = (1-\lambda) R_t A_t, \tag{5}$$

where:

$$\lambda \equiv \left[1 - \beta^{\frac{1}{\gamma}}\left[E\left(r_t^{1-\gamma}\right)\right]^{\frac{1}{\gamma}}\right], \quad r_t \equiv \frac{R_t p_t}{p_{t+1}}, \tag{6}$$

and lifetime welfare is:

$$V^*(A_t, p_t) = \frac{\lambda^{-\gamma}}{(1-\gamma)}\left(\frac{A_t}{p_t}\right)^{1-\gamma}. \tag{7}$$

The constant λ is the marginal propensity to consume with respect to wealth, and r_t is the real interest rate in units of importables, or the consumption-based rate of return. Under the feasibility condition that $E[r_t^{1-\gamma}] < \beta^{-1}$, consumption in each period is a positive fraction of the real value of asset holdings in units of importables, and savings (i.e., assets carried over to the following period) are a positive fraction of the gross return on initial asset holdings. Notice also that because the terms of trade are known but the return on exportables is unknown when C_t is chosen, the actual realization of R_t does not affect consumption, while the actual realization of p_t does.

It is straightforward to show that, since agents cannot insure themselves against fluctuations in r_t, increased risk in consumption-based asset returns (i.e., a mean-preserving increase in σ_r^2 due to increased variability in the rate of return on exportables or in the terms of trade) [9] leads to reduced savings and increased consumption if the coefficient of relative risk aversion is lower than 1 ($\gamma < 1$), or

[8] An analysis of a similar model for the case of tariffs and trade reforms of uncertain duration is undertaken by Calvo and Mendoza (1994).

[9] For σ_r^2 to increase while keeping μ_r unchanged, it must be the case that σ^2 increases in such a way that μ is adjusted to keep $\mu + \sigma^2/2$ constant.

the intertemporal elasticity of substitution is greater than 1 $(1/\gamma > 1)$, and that an increase in the mean return has the opposite effects. [10] These results are derived by expressing λ as a function of the mean and the mean-preserving variance of asset returns:

$$\lambda(\mu_r, \sigma^2) \equiv 1 - (\beta \mu_r^{1-\gamma})^{1/\gamma} \exp\left(-(1-\gamma)\frac{\sigma^2}{2}\right). \tag{8}$$

In order to examine the growth implications of terms-of-trade uncertainty, the analysis focuses on the probabilistic process that governs logarithmic first differences of consumption in the competitive equilibrium. This is equivalent to adopting a first-difference filter to separate the trend and cyclical components of consumption. The use of this detrending procedure is appropriate in this case because, given Eq. (4), Eq. (5) and the statistical properties of r_t, consumption growth can be expressed as:

$$\frac{C_{t+1}}{C_t} = (l - \lambda) r_{t'}, \tag{9}$$

where r_t is log-normal.

Denote the log first-difference of consumption as $C_t \equiv \ln(C_t) - \ln(C_{t-1})$, and define $\ln(r_t) \equiv \mu + \epsilon_t$, so that ϵ_t is the period-t deviation of the log of the real interest rate from its mean. Then, it follows from Eq. (9) that ΔC is:

$$\Delta C_t = \frac{1}{\gamma}[\ln(\beta) + \ln(\mu_r)] - [(1-\gamma) + 1]\frac{\sigma^2}{2} + \epsilon_{t-1'}. \tag{10}$$

ΔC is a proxy of the growth rate of consumption. [11] The first two terms in the right-hand side of Eq. (10) define the trend of ΔC_t and the error term ϵ is the cyclical component. The variance of the cyclical component is given by σ^2. However, because $\ln(r_t)$ is white-noise, business cycles in this economy do not display persistence, and the correlation between shocks to asset returns, or to the terms of trade, and fluctuations in savings or consumption is perfectly positive. These co-movements are in sharp contrast to what is observed in the data at business-cycle frequencies. [12] Thus, the strong assumptions needed to generate closed-form solutions produce implications that render this model inappropriate for business-cycle analysis.

[10] The savings rate with respect to wealth is $1 - \lambda$.

[11] The average growth rate reflected in the expected value of C_{t+1}/C_t becomes a poor proxy for the growth rate measured by the average of ΔC_{t+1} as σ^2 rises because, since r_t is log-normal, $\ln(E[r_t]) - E[\ln(r_t)] = \sigma^2/2$.

[12] For a quantitative analysis of terms of trade and business cycles, see Mendoza (1995).

The model does provide some interesting implications for the low-frequency relationship between terms of trade and growth. Eq. (9) shows that, even in a simple framework of i.i.d. log-normal shocks, the rate of consumption growth is determined by the savings rate and the rate of change of the terms of trade. In addition, Eq. (10) shows that the relationship between terms of trade and growth can be expressed simply in terms of the determinants of the savings rate: (a) the rate of time preference $1/\beta - 1$, (b) the elasticity of intertemporal substitution $1/\gamma$ (or the degree of risk aversion γ), (c) the average rate of return on savings μ_r, and (d) the inherent riskiness of domestic asset returns and the terms of trade σ^2.

It is important to note that although whether the intertemporal elasticity of substitution is greater or less than unitary determines if changes in μ_r or σ^2 have positive or negative effects on the *level* of consumption (see Eqs. (4) and (8)), an increase in μ_r always induces an increase in the *growth* rate, regardless of the size of γ (see Eq. (10)). Thus, *countries in which the terms of trade grow at a faster rate on average also experience faster average consumption growth.* Moreover, Eq. (10) also shows that a mean-preserving increase in σ^2 (i.e., an increase in risk associated, for instance, to increased variability in the terms of trade) induces a fall (rise) in consumption growth as long as $\gamma < 2$ ($\gamma > 2$). Thus, *if the degree of risk aversion is relatively low, a mean-preserving increase in terms-of-trade variability* (i.e., *an increase in risk*) *lowers the average growth rate.* Alternatively, if risk-aversion is high, growth rises as risk rises. However, since welfare in Eq. (7) is always declining in σ^2, welfare in such a high-risk, fast-growing economy would be lower than in a low-risk, slow-growing economy with the same degree of risk aversion.

The potential empirical implications of the model can be quantified by imposing some benchmark parameters in the closed-form solution Eq. (10) and exploring the results of numerical simulations. The model is calibrated to create a benchmark economy that conforms roughly to some empirical evidence for developing economies. Parameter values for the discount factor and the mean of asset returns are set to $\beta = 0.95$ and $\mu_r = 1.07$. $\beta = 0.95$ is the value that Lucas (1987) used to calculate welfare losses of business fluctuations. The econometric evidence from Ostry and Reinhart (1992) suggests that this value of β is biased downwards relative to estimates for developing countries, but it is a convenient benchmark for illustrating how the results of Lucas's welfare analysis are altered by growth effects of terms-of-trade shocks. The real interest rate at 7% is consistent with historical evidence documented by Mehra and Prescott (1985) for the mean real interest rate on risky assets in industrial countries, which is a good proxy if developing countries are small open economies.

Table 1 lists the model's equilibrium savings rate, $1 - \lambda$; the growth rate of consumption, g_c; and welfare effects of uncertainty, W, for various combinations of σ and γ, given $\beta = 0.95$ and $\mu_r = 1.07$. The table includes results for $\gamma = 2.33$, which corresponds to the GMM estimate obtained by Ostry and Reinhart

Table 1
Effects of uncertainty on savings, welfare, and growth: model simulations ($\beta = 0.95$, $\mu_r = 1.07$)

σ	(I) $\gamma = 0.5$			(II) $\gamma = 1.5$			(III) $\gamma = 2.33$			(IV) $\gamma = 5$		
	$(1-\lambda)$	g_c	W	$(1-\lambda)$	g_c	W	$(1-\lambda)$	g_c	W	$(1-\lambda)$	g_c	W
0.0	0.9657	0.0327	–	0.9448	0.0109	–	0.9412	0.0070	–	0.9376	0.0033	–
0.01	0.9657	0.0326	0.0007	0.9449	0.0109	0.0013	0.9412	0.0070	0.0019	0.9378	0.0034	0.0038
0.02	0.9656	0.0324	0.0028	0.9449	0.0108	0.0052	0.9414	0.0071	0.0075	0.9384	0.0039	0.0152
0.04	0.9653	0.0315	0.0113	0.9452	0.0105	0.0208	0.9422	0.0073	0.0305	0.9406	0.0057	0.0637
0.06	0.9648	0.0300	0.0253	0.9457	0.0100	0.0477	0.9434	0.0076	0.0709	0.9444	0.0087	0.1546
0.08	0.9641	0.0279	0.0450	0.9463	0.0093	0.0870	0.9452	0.0081	0.1318	0.9497	0.0129	0.3088
0.10	0.9633	0.0252	0.0702	0.9472	0.0084	0.1405	0.9474	0.0087	0.2186	0.9566	0.0183	0.5723
0.12	0.9622	0.0219	0.1011	0.9482	0.0073	0.2107	0.9502	0.0094	0.3402	0.9650	0.0249	1.0613
0.15	0.9602	0.0159	0.1578	0.9502	0.0053	0.3563	0.9553	0.0107	0.6217	0.9808	0.0370	3.3602

Note: σ is the standard deviation of $\ln(r)$, γ is the coefficient of relative risk aversion, $(1-\lambda)$ is the savings rate (i.e., the fraction saved of the gross return on initial assets), g_c is the mean of the log-first-difference of consumption, and W is the welfare cost of uncertainty measured as the change in a stationary consumption stream that leaves the household indifferent between a consumption path with risk σ and a risk-less consumption path.

(1992) for Latin America, and $\sigma = 0.12$, which is the standard deviation of the log of the terms of trade for Latin American countries estimated by Mendoza (1995). The various parameterizations produce consumption growth rates in the range between 0.33 and 3.7%, which includes the average growth rates of real consumption per-capita for several developing countries over the last two decades. In particular, the Latin American benchmark ($\sigma = 0.12$ and $\gamma = 2.33$) results in an average growth rate of 0.94%, in line with the observed average per-capita growth rate of consumption in that region. The savings rates range between 93.8 and 98.1%, which imply values for the marginal propensity to consume with respect to wealth ranging between 1.9 and 6.2%.

In the two cases that $\gamma < 2$, Table 1 shows that consumption growth falls as terms-of-trade variability rises. If $\gamma = 1/2$, growth falls by more than 1.5 percentage points as σ increases from 0 to 0.15, while if $\gamma = 1.5$ growth falls by only 1/2 of a percentage point. In the first case, the savings rate declines by 1/2 of a percentage point, but in the second case, the savings rate in fact rises by 0.5. This is because, as Eqs. (8) and (10) show, both growth and the propensity to consume fall as risk increases when $1 < \gamma < 2$, whereas when $\gamma < 1$ growth falls but the propensity to consume rises as risk increases. For $\gamma = 2.33$, both the savings rate and growth increase as σ rises, the former rises from 94.1 to 95.5% and the latter increases from 0.7 to 1.1%. Similar implications follow in the case that $\gamma = 5$, except that the effects are stronger—the savings rate rises from 93.8 to 98.1% and the growth rate increases from 0.3 to 3.7%. When $\gamma = 2.33$ or 5, growth is faster the more variable are the terms of trade, but since lifetime welfare in Eq. (7) is decreasing in γ, this faster growth is accompanied by reduced welfare.

The third column in each of the four panels of Table 1 reports the welfare costs resulting from the uncertainty of terms of trade and real asset returns. The social costs of uncertainty are measured as in the method of Lucas (1987) by computing percentage compensating variations in equilibrium consumption paths that render households indifferent between a risk-free consumption path and the consumption path of a risky environment, measuring risk by the size of σ. After some manipulation of Eq. (7), the welfare cost as a function of σ, given μ_r, is:

$$W(\sigma) = \left(\frac{\lambda(\mu_r, 0)}{\lambda(\mu_r, \sigma)} \right)^{-\frac{\gamma}{1-\gamma}} - 1. \tag{11}$$

Thus, the welfare costs of uncertainty in this model are a function of how uncertainty affects the propensity to consume relative to the risk-free, Pareto-optimal case (i.e., $\lambda(\mu_r, 0)$).

The figures reported in Table 1 show in general that variability in domestic asset returns or in the terms of trade is very costly. If one considers the benchmark welfare costs of 1/10 of a percentage point obtained by Lucas (1987), the results show that the savings-under-uncertainty model produces significantly larger welfare losses, except for cases in which $\sigma = 0.01$. [13] For cases in which $\sigma = 0.12$, the welfare costs range from 10.1% when $\gamma = 1/2$ to 106.1% when $\gamma = 5$. These costs are much larger than the costs of 12% standard deviation of consumption estimated by Lucas (1987) that range from 0.65 to 13.6% as γ rises from 1 to 20. Part of the large difference between these results and those of Lucas is accounted for by the smaller standard deviations in some of Lucas's experiments. However, most of the difference is due to the fact that consumption behavior here is the outcome of an optimization exercise in which uncertainty affects not just fluctuations of consumption around a trend growth rate, but that trend growth rate itself. Lucas's computations, in contrast, are based on a hypothetical consumption function that abstracts from the distortionary effects of uncertainty on the propensity to consume and on consumption growth. [14]

The welfare costs in Table 1 for the cases in which $\gamma > 2$ illustrate the quantitative implications of the model's result that increased risk can induce faster growth and lower welfare. When $\gamma = 2.33$, the welfare costs of uncertainty rise from around 2/10s of a percentage point to 62.2% as σ increases from 0 to 0.15, even though this also implies that the average growth rate of the economy rises by nearly 1/2 of a percentage point from 0.7 to 1.1%.

3. Empirical analysis

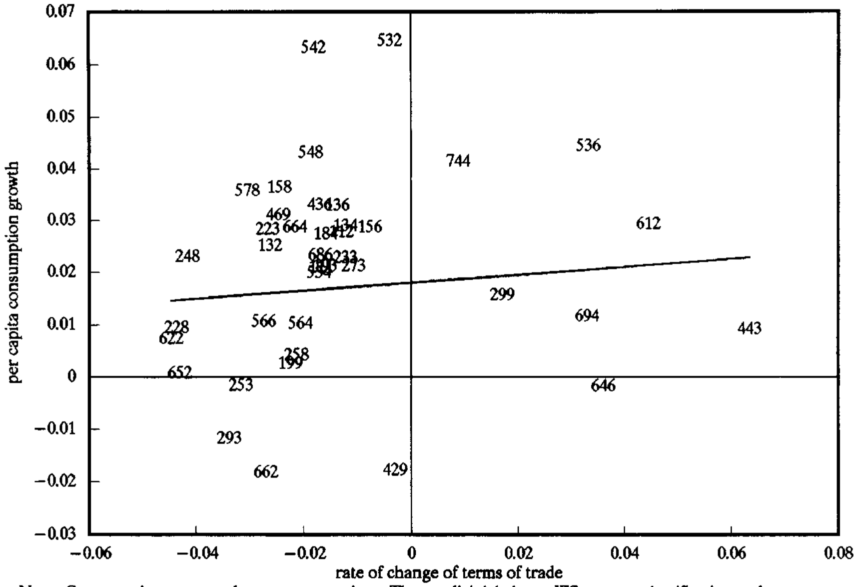
This section documents empirical regularities that characterize terms of trade and real per-capita consumption growth in 40 industrial and developing countries over the last two decades. The data are then used to conduct econometric tests based on the model's closed-form solutions. The emphasis on consumption growth, rather than GDP growth, follows from the structure of the model. The countries studied include nine industrial countries (the group of seven largest industrialized countries, G-7, plus Australia and Spain) and 31 developing countries from different regions of the world. [15]

The classification in terms of developing and industrial countries follows the *World Economic Outlook* (International Monetary Fund, 1994). The 40 countries are also classified as commodity exporters or diversified exporters using the classification system of the *WEO*, with the aim of exploring whether the export

[13] If $\sigma = 0.013$, $\gamma = 1.00001$ and μ_r is adjusted to produce growth of 3%, the model produces a welfare cost of uncertainty of 0.0016, compared with 0.00008 obtained under similar assumptions by Lucas (1987).

[14] Another important difference with Lucas (1987) is that in the savings-under-uncertainty model welfare costs of uncertainty are zero if $\gamma = 1$ because in this case λ is independent of σ. Values of γ near 1 produce small welfare losses, but these reflect the fact that the exponent in Eq. (11), which goes to infinity as γ approaches 1, amplifies even minuscule differences in propensities to consume.

[15] The full list of countries is provided later in Table 3.



Note: Consumption measured at consumer prices. Three-digit labels are IFS country classification codes.

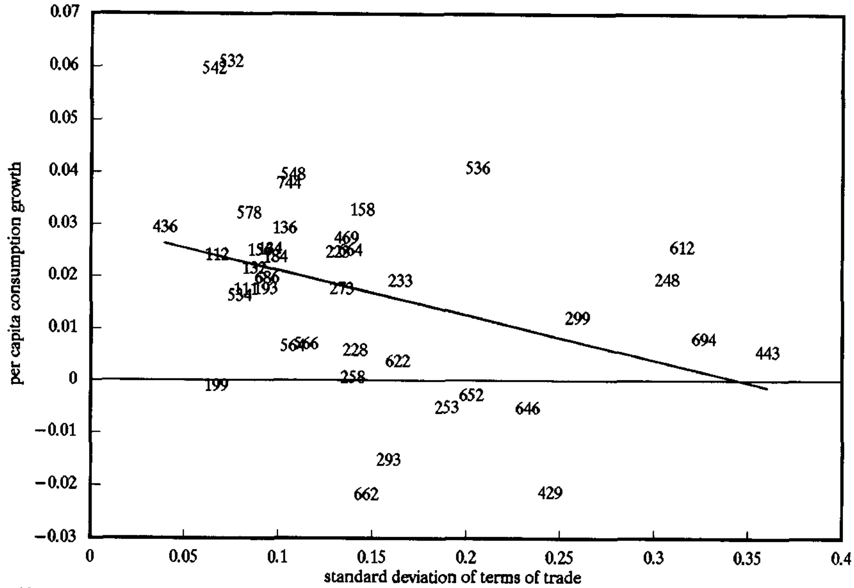
Fig. 3. Consumption growth and growth of terms of trade (country averages, 1971–1991).

base affects the growth effects of terms-of-trade uncertainty. [16] The data were obtained from the World Bank (1994) World Tables using the *Socio-economic Time-series Access and Retrieval System* (STARS) as of June 1994. The data are annual time series of private consumption at constant and current prices from national accounts; average US dollar exchange rates; US dollar import and export unit values; and total population. The sample covers the period 1970–1991.

To be consistent with the model's structure, imports are chosen as the 'numeraire,' and the terms of trade are the ratio of export to import unit values. [17] Consumption is expressed in per-capita terms and deflated using two price indices: import unit values and the consumption deflator. The econometric tests are conducted with both measures. Consumption at import prices is the measure consistent with the model, and consumption at consumer prices is used to examine valuation effects. Growth rates are measured as logarithmic first differences.

[16] Diversified exporters include countries in the sample that *WEO* classifies as exporters of manufactures or diversified exporters. Commodity exporters include developing countries in the sample that *WEO* classifies as exporters of fuel, non-fuel primary products, and exporters of services or recipients of transfers. *WEO* classifications are based on average export shares for 1984–1986 (e.g., a country is a fuel exporter if fuel exports accounted for more than 50% of total exports on average in 1984–1986).

[17] See Mendoza (1995) for a discussion of alternative measures of the terms of trade.



Note: Data measured at consumer prices. Three-digit figures are IFS country codes.

Fig. 4. Consumption growth and the variability of terms of trade (country averages, 1971–1991).

The empirical analysis begins with an informal illustration of the connection between terms of trade and growth. Figs. 3 and 4 plot scatter diagrams of country averages of consumption growth rates and the means and standard deviations of the rates of change of terms of trade, including simple regression lines. These charts provide visual evidence in favor of the model's key predictions, which are tested formally below. Fig. 3 shows that consumption growth tends to be faster in countries where the rate of change of terms of trade is higher—a well-known fact highlighted in the recent empirical growth literature. Fig. 4 shows the less well-known regularity that growth tends to be slower as the variability of terms of trade increases.

As noted in Section 2, the closed-form solutions Eqs. (9) and (10) provide a framework for testing two hypotheses: (a) the time-series hypothesis that consumption growth rates and rates of change of terms of trade are positively related over time within each country, and (b) the cross-section hypothesis that the variability of the rate of change of terms of trade, as a measure of risk, provides relevant information for explaining growth not captured by the mean. These hypotheses are tested using the multi-country data base described above and the panel estimation techniques typical of the modern empirical growth literature.

The goal of the econometric analysis is limited to establishing whether there is evidence to support or reject the two hypotheses derived from the model, rather than providing a comprehensive analysis of the determinants of growth. In particular, the first battery of tests does not consider variables other than the terms of trade for explaining growth, so panel regressions are not expected to produce high R^2 or adjusted R^2 statistics. Note, however, that Easterly et al. (1993), Fischer (1993), and Barro and Sala-i-Martin (1995) find that the terms of trade are robust determinants of economic growth even in the presence of variables that measure country characteristics and economic policies. Moreover, when these authors address simultaneity problems and apply instrumental variable methods, they find that treating the terms of trade as a truly exogenous variable does not alter the outcome of panel regressions significantly. Thus, there is strong evidence to support the view that the contribution of the terms of trade to explain growth can be examined in a simple bivariate framework.

The model's hypotheses are tested jointly with other important assumptions of the model (i.e., that consumption growth rates and rates of change in terms of trade are log-normal stationary processes, that the marginal propensity to save with respect to wealth is a time-invariant positive fraction, and that this marginal propensity to save varies with structural parameters and with the mean and variance of the terms of trade). The empirical tests also assume that the gross domestic real rate of return (i.e., R_t) is a stationary process with mean R (where $R > 1$), and independent of the process governing terms-of-trade shocks.

The validity of the assumptions that consumption growth and the rates of change in terms of trade in each of the 40 countries are white-noise processes is assessed using Box–Jenkins methods and augmented Dickey–Fuller tests. Since the data are expressed as logarithmic first differences, they are expected to be stationary and display only weak serial autocorrelation. The tests generally support the view that the data are stationary, but for some countries the Box–Jenkins method could not reject the hypothesis that there are significant, albeit low, first-order autocorrelations. To explore the issue of stationarity further, the panel tests discussed below were also performed adding linear trends. These were not statistically significant and did not affect the results markedly. Thus, the results presented in the tables exclude the linear trend.

3.1. Hypothesis 1: does average terms-of-trade inflation explain growth?

Tables 2–9 report the results of several panel regressions that test the hypothesis that the rate of change of terms of trade is a determinant of consumption growth. The tables list results for panel models based on total or pooled, between-means, fixed-effects, random-effects, and independent regression models with the aim of exploring the validity of the hypothesis for each country's time series as well as across countries. This also provides evidence as to whether the structure of the link between terms of trade and growth is homogeneous across countries, and helps establish a close link with the recent literature on empirical growth studies. Results are reported for tests that use consumption growth

Table 2
Consumption growth and the rate of change of the terms of trade: panel test results[a]

	Intercept	Slope	F-test Against Total	F-test Against Independent	$\bar{R}^2$
Data at import prices					
Total	0.006 (1.128)	0.201 (6.498)[b]	–	2.844[b]	0.048
Between means	0.006 (1.143)	0.162 (0.916)	–	–	–
Fixed effects	–	0.202 (6.404)[b]	0.699	4.855[b]	0.001
Random effects	0.006 (0.865)	0.202 (6.367)[b]	–	–	0.001
Data at consumer prices					
Total	0.018 (7.535)[b]	0.048 (3.211)[b]	–	1.810[b]	0.011
Between means	0.018 (5.435)[b]	0.076 (0.624)	–	–	–
Fixed effects	–	0.047 (3.147)[b]	1.470[b]	2.073[b]	–
Random effects	0.0178 (6.098)[b]	0.048 (3.133)[b]	–	–	–

[a]Intercept and slope coefficients for regressions of consumption growth on the rate of change of the terms of trade and a constant. Numbers in brackets are t-statistics.
[b]Denotes statistical significance at the 5% level.
Results for independent time-series regressions are reported in Table 3 and Table 4.

measured at import and consumption prices, and they are also applied to sub-samples arranged according to geographical groups and the composition of exports.

The results show that, in general, the simple endogenous growth model of Section 2 performs well when confronted with the data. Consider first the general panel results (Table 2). Regardless of whether data at import or consumer prices are used, total, fixed-effects, and random-effects models show that terms of trade are a statistically significant explanatory variable of consumption growth, although short of the unitary coefficient that the model predicts. An increase of 1 percentage point in the rate of change of terms of trade increases the growth rate of consumption at import prices by 0.2% and that of consumption at domestic prices by about 0.05%. The weaker effect on the latter is suggestive of the implications of relative price changes on the consumer goods basket. Adjusted R^2 statistics are generally low, as expected, because the role of growth determinants other than the terms of trade is not considered.

The between-means models are the only ones in Table 2 that reject the hypothesis that the terms of trade explain growth. This is consistent with the finding of Fischer (1993) that the role of terms of trade is more significant in pooled regressions than in-between-means regressions, and with the conjecture of de Gregorio (1992) that the terms of trade are irrelevant for explaining growth, which he derived from between-means estimation. Moreover, as shown later, the poor results of the between-means models reflect the fact that relevant information contained on the variability of the terms of trade is missing. Panel models that include a time-series dimension capture some of this information, and this explains in part their good performance.

Table 3
Real per-capita consumption growth at import prices and rate of change of the terms of trade: country time series regressions (Data for 1971–1991)

Country	Intercept	Slope	F-statistic	$\bar{R}^2$	D.W.
Industrial countries					
United States	0.016	1.183[b]	51.010[b]	0.714	1.770
United Kingdom	0.033[b]	1.549[b]	54.234[b]	0.727	1.929
Japan	0.075[b]	1.429[b]	334.535[b]	0.943	1.156
France	0.045[b]	1.361[b]	82.947[b]	0.804	1.729
Italy	0.039[b]	1.241[b]	130.020[b]	0.866	2.013
Canada	0.019	0.421[b]	4.722[b]	0.166	1.192
Germany	0.032[b]	1.189[b]	163.584[b]	0.890	1.798
Australia	0.020	0.506[b]	10.794[b]	0.329	2.274
Spain	0.052[b]	1.204[b]	74.048[b]	0.785	1.089
Developing countries					
Asia					
Hong Kong	0.055[b]	0.216	0.892	–	2.188
India	−0.017	1.289[b]	25.541[b]	0.551	2.331
Indonesia[a]	0.017	0.073	0.254	–	1.688
South Korea	0.086[b]	1.481[b]	59.421[b]	0.745	1.338
Malaysia	0.006	0.374[b]	4.096[b]	0.134	1.073
Pakistan[a]	−0.058	−1.005[b]	8.912[b]	0.284	1.238
Philippines	−0.018	−0.096	0.387	–	1.395
Thailand	0.015	0.255	1.233	0.012	1.540
Africa					
Algeria[a]	0.001	0.037	0.121	–	0.866
Cameroon[a]	0.001	−0.158	1.125	0.006	1.743
Côte d'Ivoire[a]	−0.004	0.159	0.551	–	1.895
Gabon[a]	0.042	−0.154	0.629	–	2.508
Ghana[a]	−0.048	−0.172	0.210	–	1.200
Kenya[a]	−0.035	0.156	0.683	–	2.233
Nigeria[a]	−0.072	0.331[b]	5.484[b]	0.183	1.684
South Africa	−0.005	0.696	3.123	0.096	1.234
Tunisia	0.014	−0.349	2.697	0.078	2.278
Morocco	−0.004	0.168	0.266	−0.038	2.046
Middle East					
Egypt[a]	−0.023	0.091	0.108	–	1.459
Iran[a]	0.058	0.006	0.002	–	1.241
Israel	0.028	0.714	1.542	0.026	2.088
Kuwait[a]	−0.010	0.102	1.299	0.016	1.721
Western Hemisphere					
Brazil	0.034	0.845[b]	15.452[b]	0.419	2.223
Colombia[a]	0.003	0.152	2.247	0.059	0.701
Ecuador[a]	−0.013	0.175[b]	5.714[b]	0.191	1.642
El Salvador[a]	0.007	0.135	1.572	0.028	1.650
Guatemala[a]	−0.009	0.239	1.357	0.018	1.878
Mexico[a]	0.006	0.808[b]	9.248[b]	0.292	2.135
Peru[a]	−0.003	−0.038	0.014	–	2.039
Venezuela[a]	0.002	−0.046	0.178	−0.043	1.969
Chile[a]	−0.019	0.293	0.553	–	1.144

[a]Countries classified as commodity exporters, as explained in the text.
[b]Denotes that intercept, slope, or F − statistic estimates are significantly different from zero at the 5% confidence level.

Table 4
Real per-capita consumption growth at consumer prices and the rate of change of the terms of trade: country time series regressions (Data for 1971–1991)

Country	Intercept	Slope	F-statistic	$\bar{R}^2$	D.W.
Industrial countries					
United States	0.019[b]	0.114[b]	5.570[b]	0.186	0.986
United Kingdom	0.024[b]	0.017	0.024	–	1.131
Japan	0.033[b]	0.031	1.115	0.001	1.898
France	0.023[b]	0.050	3.268	0.102	0.931
Italy	0.028[b]	−0.059	2.439	0.067	1.490
Canada	0.025[b]	−0.008	0.010	–	1.077
Germany	0.026[b]	0.079[b]	4.724[b]	0.157	0.806
Australia	0.020[b]	0.143[b]	26.000[b]	0.556	1.982
Spain	0.024[b]	−0.001	–	–	0.602
Developing countries					
Asia					
Hong Kong	0.061[b]	−0.070	0.269	–	1.841
India	0.019	0.150	1.154	0.008	2.806
Indonesia[a]	0.037[b]	0.095[b]	6.803[b]	0.225	1.569
South Korea	0.063[b]	0.193	3.034	0.092	2.095
Malaysia	0.045[b]	0.287[b]	7.310[b]	0.240	1.122
Pakistan[a]	0.003	−0.123	1.162	0.008	1.632
Philippines	0.008	0.028	0.231	–	1.133
Thailand	0.039[b]	0.245[b]	5.427[b]	0.181	1.927
Africa					
Algeria[a]	0.020	0.125[b]	10.217[b]	0.315	2.139
Cameroon[a]	0.005	0.023	0.056	–	1.404
Côte d'Ivoire[a]	−0.016	0.216	3.715[b]	0.120	1.917
Gabon[a]	0.005	−0.293[b]	7.164[b]	0.236	2.171
Ghana[a]	−0.007	−0.081	0.334	–	2.101
Kenya[a]	0.025	−0.019	0.011	–	1.546
Nigeria[a]	0.001	0.216[b]	6.057[b]	0.202	2.396
South Africa	0.005	0.263	1.478	0.023	2.582
Tunisia	0.037[b]	0.019	0.043	–	1.655
Morocco	0.020[b]	0.020	0.048	–	2.712
Middle East					
Egypt[a]	0.030[b]	0.104	1.389	0.023	1.292
Iran[a]	−0.022	−0.118	0.740	–	2.005
Israel	0.021	−0.508	2.858	0.085	2.586
Kuwait[a]	−0.002	0.110	1.646	0.037	1.785
Western Hemisphere					
Brazil	0.025[b]	0.020	0.054	–	1.564
Colombia[a]	0.019[b]	−0.037	1.173	0.010	1.225
Ecuador[a]	0.020[b]	0.019	0.753	–	0.533
El Salvador[a]	−0.002	0.105	3.156	0.097	0.933
Guatemala[a]	0.001	0.032	0.309	–	0.917
Mexico[a]	0.019	0.123	2.605	0.074	1.251
Peru[a]	−0.017	−0.038	0.176	–	1.304
Venezuela[a]	0.013	−0.082	2.941	0.108	1.065
Chile[a]	0.017	0.262	0.951	–	2.236

[a]Countries classified as commodity exporters, as explained in the text.
[b]Denotes that intercept, slope, or F-statistic estimates are significantly different from zero at the 5% confidence level.

Table 5
Consumption growth and the rate of change of the terms of trade: results for regional panels[a] (Data at import prices for 1971–1991)

	Intercept	Slope	F-test Total	Against Independent	$\bar{R}^2$
Industrial countries					
Total	0.036[b]	1.142[b]	–	4.543[b]	0.722
Between means	0.011	−0.360	–	–	–
Fixed effects	–	1.147[b]	1.122	7.631	0.721[b]
Random effects	0.036[b]	1.145[b]	–	–	0.716
Asia					
Total	0.008	0.131	–	5.898[b]	0.009
Between means	0.013	0.569	–	–	–
Fixed effects	–	0.118	2.009	9.0725[b]	–
Random effects	0.007	0.126	–	–	–
Africa					
Total	−0.011	0.070	–	0.923	0.001
Between means	−0.011	0.191	–	–	–
Fixed effects	–	0.067	0.550	1.289	–
Random effects	−0.011	0.069	–	–	–
Middle East					
Total	0.024[b]	0.088[b]	–	0.783	0.013
Between means	0.026	−0.328	–	–	–
Fixed effects	–	0.099	0.874	0.704	–
Random effects	0.024	0.093	–	–	–
Western Hemisphere					
Total	−0.001	0.200[b]	–	1.003	0.046
Between means	0.005	0.445	–	–	0.260
Fixed effects	–	0.197[b]	0.119	1.882	0.001
Random effects	−0.001	0.199[b]	–	–	0.002

[a] Intercept and slope coefficients for regressions of consumption growth on the rate of change of the terms of trade and a constant. Numbers in brackets are t-statistics.
[b] Denotes statistical significance at the 5% level.
Results for independent time-series regressions are reported in Table 3 and Table 4.

The results in Table 2 also show that F-tests reject the hypothesis that intercept and slope coefficients of the pooled and fixed effects models are equal to those of independent regressions. The failure of these homogeneity tests is not surprising given large cross-country differences between slope and intercept estimates in the independent regressions of Tables 3 and 4. This in turn reflects the model's prediction that the propensity to save varies with the degree of risk aversion, the subjective rate of time preference, and the mean and variance of the terms of trade,

Table 6
Consumption growth and the rate of change of the terms of trade: results for regional panels[a] (Data at consumer prices for 1971–1991)

	Intercept	Slope	F-test Total	Against Independent	$\bar{R}^2$
Industrial countries					
Total	0.025[b]	0.036[b]	–	1.308	0.019
Between means	0.023[b]	−0.067	–	–	–
Fixed effects	–	0.037[b]	1.094	1.499	–
Random effects	0.025[b]	0.037[b]	–	–	–
Asia					
Total	0.034[b]	0.101[b]	–	3.579[b]	0.048
Between means	0.038[b]	0.385	–	–	–
Fixed effects	–	0.092[b]	4.739[b]	2.175	0.005
Random effects	0.034[b]	0.094[b]	–	–	0.005
Africa					
Total	0.009	0.061[b]	–	1.990[b]	0.011
Between means	0.001	0.171	–	–	–
Fixed effects	–	0.058[b]	0.726	3.178[b]	–
Random effects	0.009	0.059[b]	–	–	–
Middle East					
Total	0.011	0.036	–	1.349	–
Between means	0.011	−0.197	–	–	–
Fixed effects	–	0.042	1.084	1.585	–
Random effects	0.011	0.038	–	–	–
Western Hemisphere					
Total	0.001	0.026	–	0.934	–
Between means	0.013	0.184	–	–	–
Fixed effects	–	0.024	0.714	1.150	–
Random effects	0.009	0.025	–	–	–

[a] Intercept and slope coefficients for regressions of consumption growth on the rate of change of the terms of trade and a constant. Numbers in brackets are t-statistics.
[b] Denotes statistical significance at the 5% level.
Results for independent time-series regressions are reported in Table 3 and Table 4.

all of which differ widely across countries. [18] In general, however, the negative results of many of the independent regressions cast doubt on the relevance of Eq. (9) as a model of *annual* changes in growth rates. Even when these regressions produce significant coefficients for the terms of trade, the coefficients tend to be in

[18] Ostry and Reinhart (1992) report evidence on differences in risk aversion and time preference across developing countries, and Mendoza (1995) provides evidence of large cross-country differences in terms-of-trade variability.

Table 7
Consumption growth and the rate of change of the terms of trade: results for export category panels[a] (Data for the period 1971–1991)

	Intercept	Slope	F-test	Against	$\bar{R}^2$
			Total	Independent	
Data at import prices					
Commodity-based exporters					
Total	−0.005	0.093[b]	–	1.023	0.011
Between means	−0.003	0.205	–	–	0.005
Fixed effects	–	0.090[b]	0.516	1.518	–
Random effects	−0.005	0.092[b]	–	–	–
Non-commodity-based exporters					
Total	0.025[b]	0.765[b]	–	5.279[b]	0.339
Between means	0.022[b]	0.638	–	–	0.007
Fixed effects	–	0.766[b]	1.153	8.967[b]	0.319
Random effects	0.025[b]	0.765[b]	–	–	0.309
Data at consumer prices					
Commodity-based exporters					
Total	0.007[b]	0.046[b]	–	1.432	0.009
Between means	0.008[b]	0.126	–	–	0.012
Fixed effects	–	0.044[b]	0.651	2.175[b]	–
Random effects	0.007[b]	0.045[b]	–	–	–
Non-commodity-based exporters					
Total	0.028[b]	0.067[b]	–	2.431[b]	0.022
Between means	0.037[b]	0.616	–	–	0.093
Fixed effects	–	0.063[b]	3.145[b]	1.623	–
Random effects	0.028[b]	0.064[b]	–	–	–

[a] Intercept and slope coefficients for regressions of consumption growth on the rate of change of the terms of trade and a constant. Numbers in brackets are t-statistics.
[b] Denotes statistical significance at the 5% level.
Results for independent time-series regressions are reported in Table 3 and Table 4.

excess of 1, which is the value predicted by the model. Thus, as noted earlier, the model of Section 2 seems inadequate for studying high-frequency features of the data. Still, it is surprising to note that even at the annual frequency the model performs significantly better for industrial countries than for developing economies, for which terms of trade are widely believed to be a key factor in explaining growth and business cycles. For some G-7 countries, the model explains more than 75% of the annual changes in per capita consumption growth at either consumer or import prices. It is also interesting to note that, if the mean rate return on risky assets is set at 7%, as estimated by Mehra and Prescott (1985), the value of λ implied by intercept estimates of the independent regressions of industrial coun-

Table 8
Consumption growth and the rate of change of the terms of trade: panels for decade sub-samples according to export category[a] (Data at import prices for 1971–1991)

	1971–1980		1981–1991	
	Intercept	Slope	Intercept	Slope
Data at import prices				
Commodity-based exporters				
Total	0.005	0.069	−0.012	0.113
Between means	−0.001	0.322[b]	0.014	0.667
Fixed effects	–	0.043	–	0.102
Random effects	0.006	0.059	−0.012	0.109
Non-commodity-based exporters				
Total	0.001	0.659[b]	0.043[b]	0.909[b]
Between means	−0.008	0.381[b]	0.042[b]	1.137[b]
Fixed effects	–	0.676[b]	–	0.892[b]
Random effects	0.002	0.665[b]	0.043[b]	0.900[b]
Data at consumer prices				
Commodity-based exporters				
Total	0.023[b]	0.033	−0.006	0.032
Between means	0.020[b]	0.069	0.008	0.338[b]
Fixed effects	–	0.026	–	0.027
Random effects	0.022[b]	0.029	−0.006	0.030
Non-commodity-based exporters				
Total	0.034[b]	0.059[b]	0.022[b]	0.134[b]
Between means	0.039[b]	0.198	0.022[b]	0.192
Fixed effects	–	0.051[b]	–	0.130[b]
Random effects	0.034[b]	0.055[b]	0.022[b]	0.132[b]

[a] Intercept and slope coefficients for regressions of consumption growth on the rate of change of the terms of trade and a constant. Numbers in brackets are t-statistics.
[b] Denotes statistical significance at the 5% level.
Results for independent time-series regressions are reported in Table 3 and Table 4.

tries is $\lambda = 0.034$, which is close to what the calibrated version of the model implies.[19]

If the results of independent regressions are compared across geographical regions or between commodity and diversified exporters, a certain degree of homogeneity emerges in coefficient estimates. For instance, many industrial country intercepts in Table 3 are clustered in the range 0.016–0.045, whereas developing countries have smaller and more variable intercepts—some even negative values—which are suggestive of lower savings rates. The results for panel models based on data organized by regions (Tables 5 and 6) or export base (Table 7) also show much less heterogeneity on regression coefficients than the

[19] λ is solved for from the equation $0.033 = (1 - \lambda)R$, where $R = 1.07$.

Table 9
Cross-sectional regressions of consumption growth on the mean and variance of the terms of trade[a]

Sample	Number of observations	Regression coefficient for			Coefficent restrictions F-tests		White tests[b]		R^2
		Intercept (α_0)	μ_y (α_1)	σ^2 (α_2)	$\alpha_0=\alpha_1 \ln(\beta)$	$\alpha 2=(2\ \alpha_1)^{-1}-1$	F	X^2	
Data at import prices									
(a) Full sample (1971–1991)	40	0.020 (3.062)[d]	0.474 (3.069)[d]	−0.634 (−2.843)[d]	10.65[d]	9.015[d]	1.927[f]	8.834[f]	0.103
(b) Two sub-samples (1971–1980, 1981–1991)	80	0.023 (3.590)[d]	0.593 (8.381)[d]	−0.807 (−3.418)[d]	18.11[d]	14.88[d]	3.212[d]	14.266[d]	0.478
(c) First sub-sample (1971–1980)	40	−0.003 (0.453)	0.386 (4.922)[d]	−0.267 (−1.496)[f]	0.006	11.10[d]	0.672	3.596	0.373
(d) Second sub-sample	40	0.040 (6.251)[d]	0.711 (7.289)[d]	−1.158 (−4.057)[d]	50.61[d]	11.40[d]	1.193	5.971	0.641
Data at consumer prices									
(a) Full sample (1971–1991)	40	0.031 (4.627)[d]	0.344 (1.772)[e]	−0.513 (−2.063)[d]	30.73[d]	4.416[d]	1.298	6.410	0.171
(b) Two sub-samples[c] (1971–1980, 1981–1991)	78	0.024 (6.245)[d]	0.231 (4.037)[d]	−0.222 (−1.598)[f]	40.01[d]	10.23[d]	2.278[e]	10.656[e]	0.196
(c) First sub-sample[c] (1971–1980)	39	0.031 (6.717)[d]	0.109 (1.947)[e]	−0.071 (−0.575)	42.19[d]	2.647[f]	1.224	6.103	0.066
(d) Second sub-sample[c] (1981–1991)	39	0.019 (4.817)[d]	0.287 (2.631)[d]	−0.527 (−4.993)[d]	27.02[d]	3.173[d]	0.564	3.071	0.426

[a] Heteroskedastic-consistent t-statistics are shown in brackets.
[b] White Heteroskedasticity tests with cross terms.
[c] Excluding Iran due to very large negative average consumption growth in 1971–1980 and very large negative growth in 1981–1991.
[d] Denotes statistical significance at the 5% level.
[e] Denotes significance at the 10% level.
[f] Denotes significance at the 15% level.

general panel regressions. Between-means models continue to perform poorly regardless of the manner in which the data are organized, and of whether consumption is measured at consumer or import prices. Pooled, fixed-effects, and random-effects models using data at import prices detect strong positive growth effects of terms of trade in the Western Hemisphere but not in Asia, Africa and the Middle East, while the data at consumer prices suggest the opposite. In both cases, the results for industrial countries show a significant effect of terms of trade on consumption growth—in Table 6, it is even the case that the hypotheses that the savings rate and the growth effects of terms of trade are similar across industrial countries cannot be rejected.

The results for data organized according to export base in Table 7 indicate that regardless of the deflator used to measure consumption, the terms of trade are a significant determinant of growth for both commodity-based and non-commodity-based exporters. Thus, the model is more robust to the ordering according to export base than to the ordering according to geography. As before, the model performs generally better when applied to industrial countries, or to diversified exporters, than when applied to developing countries or commodity exporters.

Table 8 provides results of the panel models for two sub-samples, 1971–1980 and 1981–1991, using data organized according to export-based and both consumer and import price deflators. The results indicate that the relationship between terms of trade and growth has not been stable over time, and that growth-effects of terms of trade were generally stronger in the 1980s than in the 1970s, as previously argued by Easterly et al. (1993). Surprisingly, when the data are broken down into two sub-samples, the models for commodity-based exporters fail to detect a statistically significant link between the terms of trade and growth. The slope coefficients are slightly different from those of the full sample panels in Table 7, so the low *t*-statistics are mostly due to larger standard errors in the smaller samples used in Table 8.

3.2. Hypothesis 2: does terms-of-trade variability contribute to explain growth?

Eq. (10) establishes two cross-sectional predictions with regard to growth effects of terms of trade, given preference and technology parameters. First, a higher average gross rate of change of terms of trade increases growth. Second, increased risk, measured as a mean-preserving spread of the stochastic process driving the rate of change of terms of trade, reduces (increases) growth if $\gamma < 2$ ($\gamma > 2$). The assumption that the rate of change of terms of trade follows a log-normal process is critical for this decomposition of growth determinants in terms of mean returns and risk. The two predictions and the log-normality assumption are tested here by running cross-sectional regressions using time-series averages and variances from the multi-country database.

As in the previous tests, the tests conducted here have implicit the model's assumptions regarding linear technology, isoelastic preferences, and stationary, log-normal random shocks. However, the cross-sectional approach provides an alternative means for examining the robustness of the assumption of log-normal disturbances explicitly. If $\ln(r_i)$ is white-noise, with mean μ_i and variance σ_i^2 for country i, r_i should be white-noise with mean $\mu_{ri} = \exp(\mu_i + \sigma_i^2/2)$ and variance $\sigma_{ri}^2 = \mu_{ri}^2\ (\exp(\sigma_i^2 - 1))$. Estimates of μ_{ri} can be constructed for each country by computing time-series averages of the ratio p_{t+1}/p_t, and the log of these averages can be regressed on $(\mu_i + \sigma_i^2/2)$, where μ_i and σ_i^2 are country i's time-series mean and variances of $\ln(p_{t+1}/p_t)$. If the joint hypothesis that the slope coefficient of this regression is equal to 1 and the intercept is equal to zero cannot be rejected, the log-normality assumption is supported by the data. The results of this regression produce a slope coefficient of 1.13 with a heteroskedastic-consistent t-statistic of 10.5 and an 85% adjusted R^2. The hypothesis that this coefficient is not different from 1 and the intercept is not different from zero cannot be rejected at the 5% level—the Wald F-statistic is 1.538 with a probability level of 0.22. [20] Thus, the data cannot reject the hypothesis that the rate of change of terms of trade follows a log-normal distribution.

The results of the cross-sectional regressions testing Eq. (10) are reported in Table 9. The regressions use country averages and variances of the log first difference of the terms of trade (μ_i and σ_i^2 for country i) and country averages of per capita consumption growth at consumer and import prices, calculated over the full sample 1971–1991, a panel of two sub-samples, 1971–1980 and 1981–1991, and the two-sub-samples separately. The dependent variable in each regression is average per capita consumption growth. To be consistent with the mean-preserving spread decomposition expressed in Eq. (10), the independent variables are the log of the average gross rate of change of terms of trade $\ln(\mu_{ri})$—proxied as $(\mu_i + \sigma_i^2/2)$, given the results of the log-normality test—and the variance of the log first-difference of the terms of trade σ_i^2. Thus, the coefficient on σ_i^2 measures the growth effect of a change in the variance of the rate of change of terms of trade keeping the mean constant (i.e., it measures the growth effects of increased risk).

Table 9 reports, in addition to regression coefficients, the results of Wald F-tests on cross-coefficient restrictions implied by the closed-form solution Eq. (10), results of White Heteroskedasticity Tests using cross-products of explanatory variables, and adjusted R^2 statistics. Results are based on Heteroskedastic-consistent standard errors, although these were not substantially different from ordinary least squares standard errors.

The regression results of Table 9 provide strong support for the cross-sectional predictions of the model. The coefficient estimates are generally statistically

[20] Excluding Iran and Venezuela, where violation of log-normality is clear from visual inspection of the data, the slope coefficient is 1.01 with a heteroskedastic-consistent t-statistic of 71.4 and a R^2 of 0.999.

significant and with the correct sign. The coefficients on the mean gross rate of change of terms of trade, which correspond to the intertemporal elasticity of substitution $1/\gamma$, are 0.47 and 0.34 for the full-sample models based on data at import and consumer prices, respectively. These coefficients are highly significant, although, as discussed below, the tests of cross-coefficient restrictions that link them to the intercepts and the coefficients on the variance of the terms of trade produce poor results. The basic implication that the risk reflected in the variability of the terms of trade is a key determinant of growth is also strongly supported by the data. According to full sample regressions, a mean-preserving increase of 1 percentage point in the variability of the terms of trade reduces growth by slightly more than 1/2 of a percentage point for both consumption at import prices and at consumer prices. The two full-sample models explain 10 and 17% of the cross-country differences in average growth rates, respectively.

The results of the sub-sample models suggest that there have been structural changes in the relationship between terms of trade and growth, as indicated earlier by the results of the panel regressions in 3.1. The growth effects of the mean and variance of terms of trade are significantly smaller during the 1970s than during the 1980s. In the case of consumption at import (consumer) prices, an increase of 1% in the measure of risk associated with terms-of-trade volatility reduces growth by more than 1.5 (0.5) percentage points in the 1981–1991 regression, compared to only about 0.25 (0.07) percentage points in the 1971–1980 regression. Compared to the full-sample model with data at import prices, allowing for this structural change by running a cross-sectional regression combining the two sub-samples increases the growth effect of risk from -0.6 to -0.8 and allows the model to account for nearly 50% of the differences in cross-country growth rates, instead of only 10%.

The cross-coefficient restriction implied by Eq. (10) stating that the intercept in the regressions of Table 9 should be equal to the coefficient on the mean rate of change of terms of trade times the log of the subjective discount factor (assumed at 0.99) is rejected by the data. This, however, should not be viewed as a significant drawback because the definition of asset returns implicit in the tests ignores relevant information pertaining to real domestic asset returns (R_t in the model of Section 2), and hence it is likely to be a poor proxy for the corresponding variable defined in the closed-form solution Eq. (10). Suppose, for instance, that domestic assets yield a risk-free real return of 4% in units of exportable goods. The intercept should then be equal to the coefficient on the mean of terms-of-trade growth, times the logarithmic sum of the discount factor and the risk-free domestic rate of return. The Wald test for this restriction, using the full sample regression with import price data, cannot reject the hypothesis that the restriction holds (the F-statistic is 1.521 with a probability value of 0.225). Thus, the rejection of the intercept restriction is in fact a positive aspect of the results indicating that terms-of-trade risk and return factors are not the only determinants of consumption-based, domestic asset returns.

3.3. Limitations of the analysis

The observation that there is a negative relationship between terms-of-trade variability and economic growth is an important finding of this paper that relates it to other recent empirical research on volatility and growth (see, e.g., Ramey and Ramey, 1994). However, the theoretical framework proposed in Section 2 has several important limitations in explaining this relationship that are worth noting.

3.3.1. Parameter uncertainty

Tests of the restriction that links the coefficients on the mean and variance of terms of trade in estimates of Eq. (10) examine whether the value of γ or $1/\gamma$ implied by the coefficient on the mean is consistent with that implicit in the risk effect captured by the variance. This restriction clearly does not hold, as the coefficients on the variance of terms-of-trade growth in the full-sample regressions of Table 9 imply estimates of $1/\gamma$ around 1 and 1.25, while those implied by the coefficients on the mean are 0.47 and 0.34. Thus, while the simple stochastic endogenous growth model proposed in this paper is a good first approximation, it cannot fully account for the link between growth and uncertainty. One important missing element is preference and technology parameter uncertainty. The cross-coefficient restriction on the mean and variance coefficients assumes that tastes and technology are identical across countries, which is unlikely to be true. For instance, the estimates of the intertemporal elasticity of substitution of the full-sample and joint sub-sample models, ranging between 0.23 and 0.59, are very close to the developing country pooled estimates of Ostry and Reinhart (1992) at 0.38 or 0.5, but these authors' also found significant differences in the value of this elasticity in Asia relative to Africa and Latin America. [21] Similar arguments could be made for technology parameters if the model included production and investment.

As a way of illustrating the potential implications of differences in economic structure, the full-sample models of Table 9 were reestimated introducing a dummy variable set at 0 for industrial countries, 1 for Asian countries, and 2 for all other countries. The F-statistics for the cross-coefficient restriction tests linking coefficients on the mean and variance of terms of trade are now 4.70 and 2.47 for data at import and consumer prices, respectively, compared to 9 and 4.4 in Table 9. For data at import (consumer) prices, the probability value associated to the F-statistic is 0.04 (0.22). In both cases, the restriction cannot be rejected at the 1% confidence level. Thus, there is rough evidence suggesting that cross-country differences on preference and technology parameters may account in part for the failure of cross-coefficient restrictions.

[21] These authors also showed that the decomposition of consumption between traded and non-traded goods affects estimates of the intertemporal elasticity of substitution. The model of Section 2 abstracts from this sectoral issue.

3.3.2. Time-separable preferences

Another restrictive aspect of the model that could help explain the failure of the coefficient restriction between μ_r and σ^2 is the assumption of the isoelastic, time-separable utility function Eq. (1) that forces risk aversion and intertemporal substitution to be governed by the same parameter. The failure of the restriction could be interpreted as evidence suggesting that this assumption is rejected by the data, thus favoring an alternative specification of preferences such as the non-expected utility function proposed by Epstein and Zin (1991). Obstfeld (1994) shows that in a linear technology setting with log-linear uncertainty similar to the one examined here, the use of non-expected utility results in similar closed-form solutions, except that the coefficient on σ^2 in the growth equation is a function of the product of risk aversion and (1 minus) intertemporal substitution. The result of the conventional expected utility function applies only in the case that the latter is truly the reciprocal of the former. Otherwise, increased risk reduces growth if intertemporal substitution exceeds unity, *regardless of the degree of risk aversion.* [22] Thus, under the Epstein–Zin preferences a negative effect of terms-of-trade variability on growth is consistent with a high degree of risk aversion, as may be the case in developing economies. The larger the intertemporal elasticity of substitution relative to the degree of risk aversion, the stronger the negative effect of risk on growth. Moreover, if non-expected utility holds with an intertemporal elasticity of substitution higher than 1, and in addition intertemporal substitution exceeds the reciprocal of the coefficient of risk aversion, one could explain negative coefficients on σ^2 larger in absolute value than the positive coefficients on μ_r, as reported in Table 9.

3.3.3. Robustness

The argument given for excluding from the analysis other growth determinants is the evidence provided in the empirical literature based on Barro (1991) growth–regression framework, showing that the terms of trade are exogenous with respect to those other determinants. For instance, Barro and Sala-i-Martin (1995) argue that the terms of trade should enter as their own instrument in the instrumental-variables (IV) estimation of cross-country panel growth regressions. Nevertheless, it is valuable to strengthen this argument by providing evidence that

[22] According to Obstfeld (1994), higher risk reduces growth if $\gamma > 1$, not 2, because growth is defined as $\ln E[C_{t+1}/C_t]$, instead of $E[\ln(C_{t+1}/C_t)]$. As noted in [10], these two measures differ under log-normal uncertainty. The second measure is consistent with the empirical growth literature, which uses log-first differences or 'net' growth rates.

the adverse growth effect of terms-of-trade variability is robust to the addition of other growth determinants. This evidence is reported in Table 10, which repeats some of the exercises of Table 9 adding measures of educational attainment, conditional growth convergence, the size of the physical capital stock, and government distortions on market forces. The corresponding proxies for these variables are enrollment in secondary education, initial real GDP per capita, the share of investment in GDP, and the black-market premium on foreign exchange, obtained from the database prepared by Barro and Lee (1993). Table 10 reports results for IV regressions, executed as in the method of Barro and Sala-i-Martin (1995), and for reduced-form equations of the simultaneous-equation system postulated implicitly in the IV setting. The coefficient on terms-of-trade variability in IV regressions is a 'partial derivative' that measures the direct effect of terms-of-trade variability on growth, keeping constant the other right-hand-side variables, which are functions of terms-of-trade variability in the 'first stage' regressions. The reduced-form coefficient measures the total effect of terms-of-trade variability on growth operating through different channels, and thus it is more appropriate for comparison with Table 9 in assessing the robustness of the growth effect of terms-of-trade variability. The analysis also includes a growth regression that uses GDP growth, not consumption growth, as the dependent variable.

As Table 10 shows, the negative growth effect of terms-of-trade variability is robust to the addition of the other major growth determinants. In particular, the reduced-form models show that a mean-preserving rise in the variance of the terms of trade reduces GDP growth and consumption growth at consumer prices by 2/5 of a percentage point, while the effect on consumption growth measured at import prices is stronger at 1/2 of a percentage point. These growth effects are slightly smaller than those identified in Table 9, but they remain statistically significant. Also, the choice of GDP or consumption growth as dependent variables has no significant implications for the results.

3.3.4. *Alternative transmission mechanisms*

The model proposed in Section 2 is a useful first approximation to explain why terms-of-trade uncertainty may affect growth through a basic precautionary-savings channel. It has the advantages that it provides closed-form solutions that are easily interpreted to establish analytically the factors that determine the effect of terms-of-trade variability on growth, and to show how welfare costs of macroeconomic uncertainty can be much larger than first thought. However, the model fails to capture other relevant transmission mechanisms by which terms-of-trade uncertainty could affect growth. Thus, the strong evidence of a negative effect of terms-of-trade variability on growth does not need to be tied to the one-sector, savings-under-uncertainty growth model.

One alternative is to consider the existence of a non-traded goods sector. If, for instance, preferences are represented by a Cobb–Douglas utility function that

Table 10
Cross-sectional regressions of consumption and GDP growth on the mean and variance of the terms of trade and other growth determinants[a]

Sample	Number of observations[b]	Regression coefficients							R^2	Estimation method
		Intercept	μ_{TOT}	σ^2_{TOT}	I/Y	SYR	LIGDP	LBMP		
Consumption growth at import prices										
(1) Full sample	39	–	0.375	−0.548	0.090	–	–	–	0.251	IV
		–	(2.615)^{c}	(−3.153)^{c}	(4.421)^{c}					
(2) Full sample	39	–	0.250	−0.403	0.088	–	–	−0.023	0.293	IV
		–	(1.897)^{d}	(−2.386)^{c}	(4.458)^{c}			(−2.140)^{c}		
(3) Two sub-samples	78	–	0.618	−0.791	0.108	–	–	–	0.505	IV
		–	(9.000)^{c}	(−3.638)^{c}	(4.364)^{c}					
(4) Two sub-samples	67	–	0.623	−0.488	0.104	–	–	−0.018	0.454	IV
		–	(7.264)^{c}	(−3.251)^{c}	(4.373)^{c}			(−2.085)^{c}		
(5) Full sample	39	–	0.435	−0.505	0.089	–	–	–	0.168	RF
		–	(2.545)^{c}	(−2.761)^{c}	(4.149)^{c}					
Consumption growth at consumer prices										
(6) Full sample	39	–	0.142	−0.279	0.117	–	–	–	0.244	IV
		–	(1.166)	(−2.288)^{c}	(6.729)^{c}					
(7) Full sample	33	0.077	0.258	−0.346	0.152	0.003	−0.010	−0.002	0.475	IV
		(3.543)^{c}	(2.277)^{c}	(−1.656)^{d}	(2.911)^{c}	(1.102)	(−3.329)^{c}	(−2.140)^{c}		
(8) Two sub-samples	76	–	0.202	−0.133	0.110	–	–	–	0.277	IV
		–	(4.074)^{c}	(−1.417)^{d}	(8.678)^{c}					
(9) Full sample	34	0.128	0.498	−0.413	0.128	0.008	−0.016	−0.019	0.387	RF
		(3.784)^{c}	(4.177)^{c}	(−2.144)^{c}	(3.681)^{c}	(1.997)^{c}	(−3.413)^{c}	(−1.315)^{e}		
Real per-capita GDP growth—Heston–Summers PPP measure										
(10) Full sample	34	0.184	0.421	−0.423	0.120	0.012	−0.023	−0.029	0.444	RF
		(4.568)^{c}	(3.123)^{c}	(−1.971)^{c}	(2.948)^{c}	(2.993)^{c}	(−4.060)^{c}	(−2.022)^{c}		

[a] I/Y is the real investment–GDP ratio based on Heston–Summers data, SYR are average years of secondary education in population older than 25, LIGDP is the logarithm of initial real per-capita GDP from Heston–Summers data, LBMP is the logarithm of 1 + the currency black market premium. I/Y, SYR, LIGDP, LBMP, and GDP growth were obtained from the Barro–Lee database. IV denotes estimation by instrumental variables as in Chapter 12 of Barro and Sala-i-Martin (1995). RF corresponds to reduced-form models in which the instruments are used as explanatory variables in growth regressions. Heteroskedastic-consistent t-statistics are shown in brackets.

[b] Sample size adjusted according to data gaps in Barro–Lee database.

[c] Denotes statistical significance at the 5% level.

[d] Denotes significance at the 10% levels.

[e] Denotes significance at the 15% levels.

depends on consumption of an imported good and a home, non-traded good, it is straightforward to show that, under the same assumptions maintained in Section 2, the consumption of *traded* goods will grow according to an optimal rule identical to Eq. (9). Aggregate consumption growth will then follow the rule:

$$\frac{C_{t+1}}{C_t} = \left[(1-\lambda)r_t\right]^{\alpha} n_{t+1}^{1-\alpha}, \tag{12}$$

where α (for $0<\alpha<1$) is the share of traded goods consumption in total consumption and n_{t+1} is the growth rate of the non-traded goods sector. This result suggests the addition of right-hand side variables to the consumption growth regression to capture the effects of the non-traded sector growth, and also indicates that the growth effect of terms-of-trade variability is weaker than predicted by Eq. (10), since the coefficient on σ^2 would now be multiplied by α. Moreover, this result provides another justification for the failure of cross-coefficient restrictions in Table 9.

Another important extension is to incorporate investment decisions under uncertainty, particularly taking into account irreversible-investment or cost-of-adjustment models as studied in Dixit and Pindyck (1994). Large oil-producing countries, plan irreversible and costly exploration and extraction projects on the basis of the expected performance of the world relative price of oil. It seems therefore natural that increased variability in terms of trade, under incomplete contingent-claims markets, would result in decreased investment and slower long-run growth. The analysis of Dixit and Pindyck shows, however, that this prediction may not necessarily follow as a general result. A one-time rise in volatility reduces investment in the *short-run*, as net returns on projects that were initially profitable but close to the 'threshold' level fall below this level. But unambiguous predictions regarding *long-run* effects of increased volatility on the investment rate, and hence growth, are difficult to obtain because increased volatility reduces the relative price of capital goods (Pk/P) but rises the capital–output ratio (K/Y). Since the long-run investment rate is equal to $(Pk/P)(dK/Y)$, where d is the depreciation rate, increased volatility has two effects on the investment rate that operate in opposite directions. This occurs because higher volatility rises the net return required to make an investment project profitable, but it also makes it more likely that firms will find themselves with excess capacity, so that the expected value of the productivity of installed capital rises. Due to this ambiguity in predicting the effect of increased uncertainty on the long-run investment rate, and because of additional complications resulting from nonlinearities in optimal investment rules and problems in finding adequate risk indicators, Dixit and Pindyck warn that incorporating irreversibility into econometric studies of the effects of risk on investment rates is a difficult task. Nevertheless, their analysis provides an insightful alternative framework for explaining the need to add risk indicators to cross-country growth regressions.

4. Conclusion

This paper conducts a theoretical and empirical analysis of the growth effects of terms-of-trade uncertainty. The theoretical analysis is based on a one-sector, stochastic model of endogenous growth for a small open economy. This model predicts that (a) within a country, consumption growth is linearly related to the rate of change of terms of trade over time, and (b) across countries, average consumption growth is positively related to the average rate of change of terms of trade, and positively or negatively related to the variance depending on the degree of risk aversion. If the coefficient of relative risk aversion is smaller (greater) than 2, a mean-preserving increase in terms of trade variability reduces (increases) growth. In both cases, increased uncertainty reduces social welfare. Numerical simulations show that the welfare costs of uncertainty predicted by this model are much larger than those produced by conventional real-business-cycle models because the benefits of reducing consumption instability affect not only consumption fluctuations around trend, but the trend level of consumption as well.

The model's predictions follow from closed-form solutions that highlight the role of terms of trade as a determinant of the risk and return properties of domestic assets, and hence as a determinant of savings and growth. Thus, the link between terms of trade and growth is derived here as a feature of a neoclassical savings-under-uncertainty framework, without recourse to the market rigidities emphasized in classic trade models.

The major finding of the empirical analysis, based on data for 40 industrial and developing countries, is a large adverse effect of terms-of-trade variability on economic growth, which is robust to the addition of the other key growth determinants emphasized in recent empirical studies based on Barro's growth–regression framework. The theoretical model interprets this effect as resulting from the role of the variance of the terms of trade as an indicator of risk, although alternative interpretations based on non-separable utility or irreversible investment are equally plausible. Between-means regressions of country averages of consumption growth rates on averages and variances of the rate of growth of terms of trade produce statistically significant coefficients and high levels of explanatory power. The variance of terms of trade has a significant negative effect on growth, and includes information relevant for explaining growth not included in the mean. However, cross-coefficient restrictions implied by the model generally fail, reflecting some of its weaknesses. In particular, the model and the econometric application assume identical preference and technology parameters across countries. Tests aimed at controlling for parameter uncertainty using dummy variables produce results in which cross-coefficient restrictions do hold.

Econometric tests also provide strong support for the model's first prediction that economic growth is linearly related to the rate of change of terms of trade. The rate of change of terms of trade is found to be significant in panel regressions of per-capita consumption, reflecting the findings of some recent empirical growth studies. Between-means models generally fail, but pooled, fixed, and random effects models produce very favorable results. Also, the link between terms of trade and growth is stronger during the 1980s than during the 1970s, for non-commodity-based exporters than for commodity-based exporters, and for industrial countries than for developing countries.

The proposition that indicators of risk are relevant for growth can be extended to the other explanatory variables typically emphasized in empirical growth analysis. Thus, further research could elaborate on the role of risk and uncertainty using a more comprehensive stochastic growth model. Some progress in this area has been made recently in an ongoing study of the effects of macroeconomic volatility by the Inter-American Development Bank ('Macroeconomic Volatility in Latin America: Causes, Consequences and Policies to Assure Stability,' Inter-American Development Bank, Washington, DC, 1995). Further research could also explore the implications of allowing for heterogenous preference and technology parameters vis-a-vis alternative specifications of utility and production functions.

Acknowledgements

Helpful suggestions by Guillermo Calvo, Ravi Kumar, Jong-Wha Lee, Sergio Rebelo, Carmen Reinhart, and Jorge Roldos, as well as the comments of an anonymous referee, are gratefully acknowledged.

References

Backus, D.K., Kehoe, P.J., Kydland, F.E., 1992. International real business cycles. J. Pol. Econ. 100, 745–775.

Barro, R.J., 1991. Economic growth in a cross-section of countries. Q. J. Econ. 106, 407–443.

Barro, R.J., Lee, J., 1993. International comparisons of educational attainment. J. Monetary Econ. 32, 363–394.

Barro, R.J., Sala-i-Martin, X., 1995. Economic Growth. McGraw-Hill, New York.

Calvo, G.A., Mendoza, E.G., 1994. Trade reforms of uncertain duration and real uncertainty: a first approximation. IMF Staff Pap. 41, 555–586.

Cole, H.L., Obstfeld, M., 1991. Commodity trade and international risk sharing: how much do financial markets matter?. J. Monetary Econ. 28, 3–24.

de Gregorio, J., 1992. Economic growth in Latin America. J. Dev. Econ. 39, 59–84.

Dixit, A.K., Pindyck, R.S., 1994. Investment under uncertainty. Princeton University Press, Princeton.

Easterly, W.K., Rebelo, S.T., 1993. Fiscal policy and economic growth: an empirical investigation. J. Monetary Econ. 32, 417–458.

Easterly, W.K., Pritchett, L.M., Summers, L., 1993. Good policy or good luck? Country growth performance and temporary shocks. J. Monetary Econ. 32, 459–484.

Epstein, L.G., Zin, S., 1991. Substitution, risk aversion, and the temporal behavior of consumption and asset returns: an empirical analysis. J. Pol. Econ. 99, 263–286.

Findlay, R., 1980. The terms of trade and equilibrium growth in the world economy. Am. Econ. Rev. 70, 291–299.

Findlay, R., 1984. Growth and development in trade models. In: Jones, R., Kenen, P. (Eds.), Handbook of International Economics. North-Holland, Amsterdam.

Fischer, S., 1993. The role of macroeconomic factors in growth. J. Monetary Econ. 32, 485–512.

Hopenhayn, H.A., Muniagurria, M.E., 1993. Policy variability and economic growth. Economics Working Paper 30, Universitat Pompeu Fabra, Barcelona, February.

International Monetary Fund, 1994. World Economic Outlook. International Monetary Fund, Washington, May.

Levhari, D., Srinivasan, T.N., 1969. Optimal savings under uncertainty. Rev. Econ. Stud., 153-163.

Lewis, W.A., 1954. Economic development with unlimited supplies of labor. Manchester School 21, 139–191.

Lucas Jr., R.E., 1987. Models of Business Cycles. Blackwell, New York.

Mehra, R., Prescott, E.C., 1985. The equity premium: a puzzle. J. Monetary Econ. 15, 145–161.

Mendoza, E.G., 1991. Capital controls and the gains from trade in a business cycle model of a small open economy. IMF Staff Pap. 38, 480–505.

Mendoza, E.G., 1995. The terms of trade, the real exchange rate and economic fluctuations. Int. Econ. Rev. 36, 101–137.

Obstfeld, M., 1994. Risk taking, global diversification, and growth. Am. Econ. Rev. 84, 1310–1329.

Ostry, J.D., Reinhart, C.M., 1992. Private saving and terms of trade shocks. IMF Staff Pap. 39, 495–517.

Phelps, E.S., 1962. The accumulation of risky capital: a sequential utility analysis. Econometrica 30, 729–743.

Prebisch, R., 1950. The Economic Development of Latin America and its Principal Problems. United Nations, New York.

Ramey, G., Ramey, V.A., 1994. Cross-country evidence on the link between volatility and growth, NBER Working Paper No. 4959, National Bureau of Economic Research.

Razin, A., Yuen, C., 1994. Convergence in growth rates: the role of capital mobility and international taxation. In: Leiderman, A., Razin, L. (Eds.), Capital Mobility: the Impact on Consumption, Investment and Growth. Cambridge Univ. Press, Cambridge.

Rebelo, S.T., 1991. Long-run policy analysis and long-run growth. J. Pol. Econ. 99, 500–521.

Reinhart, C.M., Wickham, P., 1994. Commodity prices: cyclical weakness or cecular decline?. IMF Staff Pap. 41, 175–213.

Singer, H.W., 1950. The distribution of gains between borrowing and investing countries. Am. Econ. Rev. 40, 473–485.

Tesar, L.L., 1995. Evaluating the gains from international risksharing, Carnegie-Rochester Conference Series on Public Policy, forthcoming.

World Bank, 1994. World Tables: Socio-Economic Time-Series Access and Retrieval System, Version 1.0, The World Bank/IFC, Washington, DC, June.