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Country Terms of Trade 1960-2012: Trends, unit roots, over-differencing, endogeneity, time dummies, and heterogeneity

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Abstract. *The debate about the Prebisch-Singer thesis has focused on primary commodities with some extensions to manufactured goods. We analyse trends in country terms-of-trade for goods and services rather than those for commodities according to the World Bank income classification. We find that the natural logarithm of the terms of trade for all groups except for the poorest has common unit roots, but none has individual unit roots. As low-income countries have no unit roots over-differencing is inefficient and biases significance levels in first differences against the fall in the terms of trade. For the low-income countries the terms of trade of goods and services are falling at a rate that is significantly negative without and with endogeneity treatment by system GMM. A comprehensive analysis of the effects of time dummies supports the result of falling terms of trade for low-income countries. When all coefficients are country-specific 50 per cent of all low-income countries have falling terms of trade in a simultaneous equation estimation using the SUR method. Food crisis and financial crisis have no effect on the number of countries with falling terms of trade, but improve or dis-improve the terms of trade for a very small number of countries.*

Key words: country terms of trade; Prebisch-Singer thesis; long-run development; World Bank income classification.
JEL-code: F43, O19.

1. Introduction

Prebisch and Singer found a fall in the prices of developing countries' primary commodities relative to those of British manufactured goods. From their work three branches of literature emerged. First, a statistical debate did arise in regard to the question whether or not developing country terms of trade or indirect indicators for them are really falling. Second, a series of theoretical models were developed in which terms of trade changes over time could be explained. Third, the policy consequences of falling terms of trade were discussed, mainly the question whether a fall in the terms of trade should lead to industrialization policies. Our paper tries to contribute to the first branch of literature, because the prominent use of closed economy models for developing countries is misleading when terms of trade are not constant.

There are two widespread versions of the Prebisch-Singer thesis (Singer 1999). The narrow one is a statistical view on the hypothesis of a time trend in the relation between primary

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commodities and manufactured goods, also called Prebisch-Singer hypothesis (PSH). The broader one is called Prebisch-Singer thesis (PST). It is interested in developing countries' terms of trade because they are related to exports and exports are related to growth and welfare as well as questions like convergence versus divergence. The special aspect here is that trade and growth are linked through developing countries' imports of capital goods (Prebisch 1950; 1962, p.2). In this broader perspective, the commodity terms of trade were the most relevant indicator around 1950 when commodities had a larger share in exports than they had later.[2] Moreover, other data were not available for a long time. The crucial question then is whether or not the country rather than the commodity terms of trade fall in the long-run average, but not necessarily in the form of a time trend doing better than other forms.

The empirical literature on the long-run development in the terms of trade, once put into this broader perspective, indicates that what is needed are not only *commodity* terms of trade or those of *manufactures*, but also terms of trade analyses on the *country* level for all goods and services.[3] From a theoretical point of view, what matters for growth is investment; and capital goods of developing countries are mainly imported. Exports are required to pay for imported capital goods. But export growth depends on the terms of trade (see the model by Bardhan and Lewis 1970). Solow type models with imported inputs paid for by exports and endogenous terms of trade generate Solow-type results under those special conditions where they generate constant terms of trade with respect to time (see Mutz and Ziesemer 2008; Habiyaremye and Ziesemer 2012). Therefore it is important to get to know whether terms of trade are falling or not. When terms of trade change, open economy growth models will not be observationally equivalent to closed economy Solow models.

[2] Part of the mirror image is the share of manufactures in exports. It grew from about to 10 per cent to about 65 per cent in the period 1960-1995. See Chakraborty (2012), Fig.1a.

[3] Textual emphasis on total trade of a country goes back to writers such as Imlah, Baldwin, Haberler, Emmanuel, Singer and Streeten (see Sarkar 2001, Sarkar and Sarkar 2008, and Ocampo and Parra (2007) for broad surveys including this point).

Especially if the empirical problem once was in the commodity terms of trade, the more or less strong diversification of the economies then may have mitigated the problem for some countries (see Athukorala 2000 on Sri Lanka) unless developing countries specialize also on industrial goods and services with low income and price elasticities (see Sarkar (2004) for a panel of 24 countries) and with strong productivity growth (see Sarkar 2005 on Korea).[4] Keesing (1979) and Sarkar and Singer (1991) broadened the literature to include the analysis of manufactures. Kaplinsky (2006, section 2) and Chakraborty (2012) review and extend this literature. Falling terms of trade in commodities or manufactures do not necessarily imply falling country terms of trade for a panel of countries, because they may happen to occur in different sub-samples of a panel. If one of two groups of countries has falling commodity terms of trade and increasing ones for manufactures and the opposite for the other group you may find falling terms of trade for commodities and for manufactures separately with or without finding falling country terms of trade because country terms of trade include also the weights of the two groups of goods which the separate series do not take into account. UN (2008) provides a detailed analysis of terms of trade from the perspectives of groups of goods and regions. Prices of manufactures have fallen relative to those of oil, mining products and food at least since 2003. As a consequence, countries which have shifted their specialization to manufactures often have falling terms of trade, especially if they are oil importers, and some commodity exporters have increasing terms of trade. However, falling prices of manufactures are also mitigated in regard to the terms of trade when manufactured inputs with falling prices are imported.

[4] A slightly different topic is the impact of export diversification on growth. Ocampo and Parra (2007) provide data plots suggesting a positive relation. The poorest countries are least diversified. However, the very successful countries, NICs and China, as well as Middle Income Latin America and Central America and Caribbean are pretty far away from the regression line. Moreover, the direction of causality is unclear. One may speculate that third factors are driving both arguments of a spurious correlation, which we discuss briefly in the last section. One such factor mentioned by the authors themselves is integration of exporting sectors into the domestic economy. The strong association of growth with high-tech manufactures (their Figure 8.3) in connection with a shift from secondary to tertiary education in growth regressions (see Ziesemer 2011) suggests that upgrading education is another important factor.

As the results for all of these separate questions will depend not only on the groups of goods but also the samples of countries, goods and years under consideration, we look at the country terms of trade in this paper, for developed and developing countries. We are therefore not mainly interested in primary commodities (the traditional approach) or in manufactures or their cointegration in this paper.[5] Bleaney and Greenaway (1993) have shown that commodity price changes of 1 per cent induce a change in net barter term of trade of 0.3 per cent. Powell (1991) and Lutz (1999a) find a value about 0.5 per cent. But even this aspect of the terms of trade debate is not uncontroversial. Aggregate commodity indices and country-level terms of trade are found to be unrelated by Cashin and Pattillo (2006) for Sub-Saharan Africa. These papers do not provide results for trends in country terms of trade though. Bidarkota and Crucini (2000) report trends in country terms of trade, which are negative throughout but insignificantly so. They group countries according to volatility in terms of trade, not income or poverty. Ram (2004) looked at net barter terms of trade at the country level and found that 16 of 26 countries investigated had significantly negative trends (5 others had insignificantly negative trends). We will look at a larger set of countries classified according to their per capita income. Whether the changes come in the form of trends shifting up and down, in a few steps, swings, cycles or other forms, what matters for long-run development is the long-run average trend. Supply (factor accumulation and technical progress) and demand forces (and the implied income and price elasticities of export demand) are assumed to determine these developments.[6] Many of these developments (including speculation and buffer stocks) behind

[5] For papers, which are interested in resource scarcity, it is of course meaningful to look at primary commodity prices. For papers interested in industrialization strategies it is of course meaningful to look at prices of manufactures and their terms of trade (see Chakraborty 2012).

[6] See Ziesemer (1995), Bloch and Sapsford (2000), Sarkar (2001), Ocampo and Parrà (2007), Mutz and Ziesemer (2008) and Habiyaremye and Ziesemer (2012) for (references to) formal models. Each newly emerging market like first Japan, then Korea, Taiwan, Hong Kong and Singapor and most recently China and many less prominent ones bring the chance of higher export demand driving up the terms of trade of other countries and the risk of new competition undermining the terms of trade if exports are hit. The four tigers did mainly hit African and Latin American countries, but China seems to hit everyone. Intense competition from China in manufactures seems to dominate for low-income countries in regard to prices of manufactured goods (Kaplinsky 2006, p.988). As China is demanding much iron ore, steel, and soybeans (Kaplinsky (2006, p. 986)) exporters of these goods

the terms of trade may take forms other than smooth trends of course. Cuddington and Urzua (1989) correctly argue that one should not talk of a secular deterioration if statistical analysis can replace it by a one-time jump without a trend being left over. However, their sample ended in 1983 and they discuss only a one-time jump in 1921. Powell (1991) suggests drops also for 1938 and 1975, Bleaney and Greenaway (1993) find another one for the early 1980s, and Ocampo and Parra-Lancourt (2010) find downward jumps in commodity price indices around 1897, 1920 and 1979, where the latter is stretched out somewhat from 1974-1986 though, and perhaps an upward jump for 2003 with so far unclear persistence. These drops may be a reaction to a postponed smooth adjustment due to preceding extraordinary events (Powell 1991). WWI demand may have kept prices high before 1921, and resource booms before 1938 and 1975 and the Latin American debt crisis, mainly caused by a world recession in 1981-82 together with the 1982 drop. With several jumps the difference with a trend is not so big anymore. What matters, is not mainly the form but how countries are affected. Other than smooth developments may be harder to anticipate and probably cause more severe adjustment costs. From a welfare point of view this is worse than a negative time trend, as was already pointed out by Powell (1991), and therefore should not be interpreted as argument against Prebisch and Singer. Refinements are interesting but not the issue of this paper. We will limit ourselves to the analysis of time trends and the consequences of the introduction of time dummies. A major result for poor countries is that the coefficients of time dummies reflect falling country terms of trade as well as time trends do. Other results will become clearer once the model is explained.[7]

should benefit if they do not lose much on other goods. Those countries also importing these goods lose through increased prices (Kaplinsky (2006), section 3(ii)).

[7] The literature discussed here includes time period after WWII. Papers focusing on periods before WWII are Hadass and Williamson (2003) and Sarkar and Sarkar (2008).

2. The Model

The long-run trend is obtained from a regression of the natural logarithm of the terms of trade, *p*, on a time trend. Straightforward additional regressors from the time-series literature are one or more lagged dependent variables. We write this basic model per observation for country *i* at time *t* as follows.

$$log\ p_{it} = c_i + \gamma_i log\ p_{i,\ t-1} + \beta_i\, t + u_{it} \qquad (1)$$

Taking first differences (making the lagged version of this equation and subtracting it from the equation above) it yields:

$$d(log(p_{it})) = \gamma_i d(log(p_{i,\ t-1})) + \beta_i + u_{it} - u_{it-1} \qquad (2)$$

If we take expected values error terms drop out, and if $\gamma_i < 1$ this equation is stable in growth rates. The expected long-run growth rate then is:[8]

$$d(log(p_i)) = \beta_i/(1-\gamma_i) \qquad (3)$$

Ram (2004) presents a special case of this model where $\gamma_i = 0$, but adds an autoregressive process. In general, without lagged dependent variable one might run into an omitted variable bias, because lagged dependent variables tend to be highly significant. Moreover, the use of lagged dependent variables reduces serial correlation and the bias possibly caused by it. This can be seen as follows. Suppose the basic idea is captured by

[8] Bleaney and Greenaway (1993) and Erten (2011) discuss this model at greater length with all its possible outcomes.

$$log\, p_{it} = c_i + \beta_i t + u_{it} \qquad (1')$$

Next, assume that there is second order serial correlation

$$u_i = \rho_{1i} u_{it-1} + \rho_{2i} u_{it-2} + \varepsilon_{it} \qquad (4)$$

From rewriting (1') in lagged form we can find expressions for u_{it-1} and u_{it-2}. Insertion of these expressions into (4) and the result into (1') yields

$$log\, p_{it} = \rho_{i1} logp_{it-1} + \rho_{2i} logp_{it-2} + \beta_i(1-\rho_{1i}-\rho_{2i})t + c_i(1-\rho_{1i}-\rho_{2i}) + \rho_{1i}\beta + 2\rho_{2i}\beta_i + \varepsilon_{it} \qquad (1'')$$

With an adequate redefinition of coefficients this equation is identical to (1) in case of first-order serial correlation ($\rho_2 = 0$). By implication our equations and those used by Ram (2004) are equivalent if first-order autocorrelation is assumed.

Other regressors should not be included if one is only interested in getting to know whether there is a significant trend in the terms of trade. This is different of course if one is interested in explaining the terms of trade development in the sense of economic theory; then more regressors related to supply and demand are needed and the role of the time trend is to mimic technical progress.[9] But this is not the interest of this paper. We only want to know (i) to which extent there is still a negative time trend; (ii) what the analysis looks like if we add time dummies as Cuddington and Urzua (1989), and Bleaney and Greenaway (1993) did; and (iii) how country-specific time trends differ, when other coefficients are homogenous or heterogeneous.

[9] See Bloch and Sapsford (2000) and Mutz and Ziesemer (2008).

3. Data and Econometric Method: From fixed effects to more heterogeneity

We follow the 2009 World Bank classification for countries: low income (per capita income (GNI) of $975 or less in 2008), lower-middle income ($976-3855), upper-middle income ($3856-11905), high-income-non-OECD and high-income OECD (above $11906). As the issue of the trend in the terms of trade is typically discussed in regard to poor and hardly diversified countries other groupings of countries then in regard to income levels are more likely to hide the trends rather than to reveal them. The data are taken from the World Development Indicators (World Bank 2009) for the first data set and World Bank (2013) for the second. We found very similar results using the classification of 2008, which differs quite a bit from that of 2009. The similarity of the results indicates that they are robust in regard to the classification of countries.

We define the terms of trade as exports as capacity to import (ecm) divided by exports (ex), both for trade in goods and services and measured in constant local currency units. The data for the first data set were available from 1960 to 2008, with some non-available observations of course. For 2008, the year of the food crisis, the first data set has only data points for half of the countries. But in principle we have 49 observations per country.

In order to check the sensitivity with respect to the food crisis 2008 (Headey 2011)[10] and the financial crisis 2009-2012 we will update the data set including data until 2012.

In WDI 2013 there are no data for exports as capacity to import and exports anymore for the low-income countries Afghanistan, Guinea-Bissau, Niger, and Somalia; we take them over from the older data set thereby missing any update. Only all old data are used for Ghana as well because in the new data five of the seven observations lead to terms of trade of unity,

[10] Headey (2011) explains in detail how the 2008 food crisis has been triggered by a series of protectionist measures after a mild price increase since 2003. Mansfield and Reinhardt (2008) argue that such protectionist measures are often excluded by WTO or PTAs (Preferential Trade Agreements). By implication, WTO and PTA contribute to the reduction in fluctuations and improve *global* welfare. In contrast, the national protectionist measures maximize *national* welfare, which can be undermined by free trade in cases of negatively correlated stochastic events where free trade undermines the insurance function of the autarkic markets according to the Newberry-Stiglitz theorem (Bhagwati et al 1998, chap. 39).

implying constant prices for all exports and imports for five years, which is simply false. There are less data in the new compared to the old WDI for Burundi, Comoros, Haiti, and Chad; we add old data to the new ones for Burundi because they fit together without risk of using two different base years. This does not hold for Chad where we see revised data in the end of period and the missing ones are the most recent. We use only the new data for Chad. We use only new data also for Haiti, because overlap in data availability for the years 1991-3 indicates data revision of a large order of magnitude. We also use only new data for Comoros because other years yield terms of trade unity for all but one year, again this is empirically implausible. For Malawi the terms of trade are falling over the whole period but have value unity exactly for 2006-2011, casting some doubt on the new data – we delete observations 2006-2011. Terms of trade data are all unity for Uzbekistan, Laos and Yemen in the new data set. YEM is dropped together with Afghanistan, for which we have only four observations. For LAO and UZB we use the old data. For several countries there is no year in the recent data set in which the terms of trade are exactly unity, pointing to asymmetries in the rounding of the underlying two data series. Dropping AFG and YEM leaves us with 38 countries in the second updated data set.

We proceed in six steps. First, we run a fixed effects estimate. For our model as expressed in equation (1) this means that we impose a constraint, that the coefficients are identical for all countries in a sample except for the intercept. The constraint imposed on the model therefore is $\beta = \beta_i$, $\gamma = \gamma_i$. With lagged dependent variables as in our model, fixed effects estimates of the coefficient of the lagged dependent variable are biased. The bias has an order of magnitude of $1/T$, and therefore the estimate is consistent in regard to the time dimension T, but the bias is smaller when more regressors are used (Asteriou and Hall 2011, chap. 19). As a general rule, with more than thirty observations in the time dimension the bias is low enough

to use the panel fixed effects method (see Judson and Owen, 1999; Baltagi, 2008, ch.8) without instruments as we do in the first instance.

Subtraction of $logp_{t-1}$ on both sides of (1'') and rearrangement yields

$$dlogp_t = (\rho_1 + \rho_2 - 1)\, logp_{t-1} - \rho_2 dlogp_{t-1} + \beta(1-\rho_1-\rho_2)t + c(1-\rho_1-\rho_2) + \rho_1\beta + 2\rho_2\beta + \varepsilon_t \qquad (1''')$$

With adequate redefinition of symbols this equation is the one underlying the augmented Dickey-Fuller test for unit roots in time-series analysis. If it is valid for that test it should be applicable for our purposes as well. The only problem is that under the null hypothesis of a unit root the standard assumptions in regard to the distribution of the coefficient of the trend variable do not hold (Davidson and McKinnon 2004, p.617). This means that we can use it for drawing strong conclusions only in the absence of unit roots. Therefore we will test for panel unit roots. A natural way out in case of unit roots is the additional use of first differences of equations (1) or (1'') such as equation (2) with standard residual though (see McCallum 1993, Cuddington 2010). Therefore we will estimate the equations also in first differences. However, if series have no unit roots but rather are stationary this leads to overdifferencing: Differenced stationary series have moving average residuals (Maddala and Kim 1998) a special but relevant case of which is the residual difference in equation (2), and if these moving averages are not taken into account the estimates are inefficient leading to too many rejections as t-values are too low (McCallum 1993; Harvey et al. 2010). Our estimates presented below reveal that the results are indeed different then.

Second, to deal with endogeneity, we will also use system GMM, which combines in principle the level and the differenced equations above, (1) and (2), and applies instrumental variables. The econometric reasoning leading to the choice of the system GMM estimator is as follows (see Baltagi 2008, ch.8). In the presence of lagged dependent variables ignoring

non-redundant fixed effects may lead to a heterogeneity bias. The use of fixed effects leads to an expected bias – mentioned briefly above - for the coefficient of the lagged dependent variable of the order of magnitude of *1/T*, where *T* is the number of periods for which data are available. With more than forty observations in the time dimension its expected bias of order of magnitude *1/T* is small anyway. Taking first differences can remove this bias and leads to the Anderson-Hsiao estimator, which is inefficient though. The first-differences estimator by Arellano-Bond removes this inefficiency. However, it has a small sample bias. The system GMM estimator by Arellano-Bover turns out to be the best estimator according to Monte-Carlo studies by Blundell and Bond (1998) for very small *T* as well as Soto (2009) for *T = 8* and *T=15*. Baltagi (2008, chap.8) points out that even with thirty observations and an expected value of the bias of *1/T* = 3.3 per cent the actual bias may still be as large as 20 per cent. Therefore we will also apply the system GMM estimator, which estimates equations (1) and (2) simultaneously using lagged first differences of the regressors as instruments for equation (1) and lagged levels for equations (2) in the standard version. However, instead of the first-difference version we will apply the orthogonal deviation version, which uses a Helmert transformation subtracting from each residual the sum of all future residuals (see Arellano and Bover 1995). Bun and Windmeijer (2010) have pointed out that the above mentioned Monte-Carlo studies providing support for system GMM have assumed that the variance of the fixed effects and the residuals are unity. They show – for a model with a lagged dependent variable as the only regressor - that system GMM with no other regressors but the lagged dependent variable may have an upward bias of about 9 per cent for *T = 6* and of about 7 per cent for *T = 15* if the ratio of the fixed effects variance and that of the residuals is four instead of unity, but there is no bias if the variance ratio is below unity. Okui (2009) has provided a Monte Carlo study for the orthogonal deviations version of system GMM for variance ratios of unity and ten. Therefore we will also report these variances and their ratio.

As the use of a time trend is essential in our research we cannot use a full set of time fixed effects for all years but one as usual as it would lead to collinearity. We will therefore investigate how the introduction of selected time dummies changes our fixed effect results.

Third, we run the regression for all countries not only with fixed effects but also with country-specific time trends. The only constraint then is the one for a common coefficient of the lagged dependent variable(s), $\gamma = \gamma_i$. Then we will add time dummies again to the regressions for the low-income countries and analyse the consequences.

Fourth, we will relax the constraint on the lagged dependent variables also, and estimate a system of equations. The contemporaneous residuals of the countries may be correlated. Therefore we will use the SUR method (seemingly unrelated regression).

Fifth, we also look at the net barter terms of trade as found in the World Development Indicators, which is the ratio of the export and the import price indices for goods that go through the customs. As services are excluded from these data Ram (2004) speaks of 'commodity (net barter) term of trade'. Unfortunately, these series have less than 30 observations. Therefore we should use the system GMM method (see Baltagi 2008; Roodman 2009; Soto 2009) if fixed effects are not redundant. When using GMM we could not get rid of second-order serial correlation and got mostly implausibly high or low values of the Sargan statistic. Both point to invalid instruments. Probably this is due to the simplicity of our approach and therefore GMM cannot be used here. We use EGLS (estimated generalized least squares) in order to take into account the cross-section heterogeneity. When fixed effects are not redundant this leads to an expected bias in the order of magnitude *1/T* for the lagged dependent variable, which is 1/23 in our case for low income countries. Finally, also for the net-barter terms of trade we will relax all constraints and estimate a system using the SUR method.

Table 1 OVER HERE

Six, in order to check the sensitivity of the results we will use the updated data set to estimate (i) until 2007 to have results until before the food crisis; (ii) estimate results until the year 2008 in order to see the sensitivity in regard to the food crisis year; (iii) estimate again including the year 2009 to see sensitivity in regard to the financial crisis; (iv) using the complete extended data until 2012 and compare its results with all others including those from the old data set. In this last step we will use two lagged dependent variables for the case of all coefficients being country specific as countries are likely to be affected by both crises in a way that is heterogeneous within income groups. Two lags ensure absence of serial correlation and its corresponding bias; with no serial correlation the lagged dependent variables do not suffer from being endogenous. For the estimation using the new data set we follow the suggestion of a referee to estimate together with more other countries. Here we add the 27 high-income OECD countries, who are the main trading partners.[11] This is not necessarily unproblematic as low-income countries do not have individual unit roots whereas other countries do. However, the final result taking data to 2012 has 16 countries with significantly negative trends with and without adding the OECD countries.

4. Results for Country Terms of Trade before the Financial Crisis

Tables 1 - 3 show results using the data of 'exports as capacity to import divided by exports' covering all goods and non-factor services, not just commodities; they are taken in natural logarithms, and abbreviated as log(ecm/ex). We find two significant lagged dependent variables for most samples in the first instance.

[11] Adding also the high-income non-OECD countries containing many oil exporters leads to a 'Near singular matrix' warning, indicating that the inversion of X'X is not possible because of being close to linear dependence.

4.1. Trends in Log-Levels of the Terms of Trade

Table 1 shows the value of the coefficients and the marginal significance levels (p-values) in panel (a). Only the low-income countries have a significant trend, which is negative. The long-run trend, *β/(1-sum of coefficients of the lagged dependent variables)*, is also shown. For the low-income countries it is -0.42 per cent. This value for country terms of trade is less negative than the value for commodity terms of trade of -0.6 per cent of Ardeni and Wright (1992) and Sapsford and Balasubramanyam (1994) and almost equal to the value of –0.44 per cent found by Lutz (1999b). It is also in the range of the values for commodities obtained by Bleaney and Greenaway (1993) for several periods ending in 1991 and in the range of the literature surveyed by Lutz (1999b). Other goods seemingly have a very similar trend as commodities have relative to manufactures. The negative trend is stronger in the earlier periods than in later ones in our analysis (not shown), as we can see from starting the regression successively ten years later. Starting the regression successively one year later does not make the significantly negative time trend vanish though.[12]

4.2. Unit Roots and Over-Differencing

There are neither common nor individual unit roots in the poor country sample according to panel standard unit root tests.[13] Unit root tests can be found in panel (b) of Table 1. For all other country groups than the low-income countries, the hypothesis of common unit roots cannot be rejected, but for individual unit roots most tests have low p-values.[14] However, if a

[12] Alternatively we can add time dummies for the slope of the trend (see below). The Durbin-Watson statistic in Table 1, panel a, column 5 indicates that the serial correlation that might be generated by structural breaks is very limited. Note that we do not get the spurious appearance of a unit root for low-income countries below when not taking into account these breaks.

[13] Similarly, Erten (2011) finds no unit roots in net-barter terms of trade in time-series analysis for several country aggregates.

[14] We have not tested for spuriousness of the unit root as one knows it from time-series literature in case of structural breaks in the trends (see Harvey et al. 2010). This issue is currently under discussion in econometric research (see Chan and Pauwels 2009) especially in regard to the adequate modeling of the potential breaks. Moreover, we are mainly interested in the poor countries, which appear not to have a unit root and therefore are

model holds in levels it should also hold in first differences from the point of view of deterministic economic modelling. From an econometric point of view though first differences of stationary variables are called over-differenced and estimation with over-differenced variables should be avoided, because they have a bias if the implied moving average residuals are not taken into account explicitly. Our estimates for first differences are shown in panel (c) of Table 1. Low-income countries again have a significantly negative long-run time trend, which is even larger, 1.48 per cent. Also lower-middle-income countries and high-income OECD countries have a negative time trend but both are highly insignificant. Results are clearly different under first differences. For the low-income countries first differenced results should not be used because they are based on over-differencing, which in turn stems from absence of unit roots according to panel (b); the trend results of log-levels should be used then, which are in panel (a) of Table 1. For all other country groups there are unit roots and differencing is in order and therefore the results of panel (c) should be used.

4.3. Dealing with Endogeneity using System GMM

In Table 2 we present the results for the system GMM estimator in the orthogonal deviation form in column 3. The lagged dependent variable should have a coefficient which is above the underestimating one for country fixed effects in column 1 and below the one for OLS in column 2, which overestimates it (see also Durlauf et al. 2005). Indeed we found a coefficient between these two.[15] Moreover, our system GMM estimate is indeed about 2 per cent higher than the fixed effects estimator as it should be for a bias *1/T for T = 45* in the presence of no further regressors. However, we have an unbalanced panel for 45 periods and 1158 observation for 39 countries, which implies effectively *T = 30* and the bias *1/T = 3.3 per cent*

not subject to this question. In the updated data set all tests reject the unit root hypothesis at the one percent significance level for low-income countries.

[15] We use only one lag here for two reasons. First, when not using the EGLS method any more the second lag was insignificant. Second, when using two lags we cannot find any result analogous to having the coefficient of the lagged dependent variable between those of OLS and the within-estimator.

without additional regressors. With a time trend included, which should yield a slightly lower bias, a correction of 2 per cent seems reasonable. For the Sargan statistic, Davidson and MacKinnon (2004) state that it should not be too high because of the standard chi-square test. Roodman (2009b) states that it should also not be too low either because then the instruments do not do their work. We find a p-value of 28.5 per cent which is sufficiently far away from the extremes of zero and unity. Using only two lag as instruments is in line with Okui (2009). As with one instrument less we have as many instruments as we have variables and no overidentifying constraints we get a J-statistic of zero, the Sargan difference test for the last lag in the list of instruments is not different from the Sargan test and also reasonable for the same reasoning.[16] The squared ratio of the variances for fixed effects and the residuals is about 0.156 and therefore lower than assumed by Blundell and Bond (1998) and Soto (2009). There is no indication for an upward bias that exists for values higher than unity of this ratio according to Bun and Windmeijer (2010). The long-run trend which comes from this estimate is almost a negative one percent, -0.9 per cent (= 0.001428/(1-0.841584)). This value is more negative than the relevant one from Table 1, panel (a).

TABLE 2 OVER HERE

4.4. Time Dummies instead of Trends for the relevant Cases

As a modification of the simple time trend in Table 1 panel (a), column 5, we can add time dummies for the slope of the trend (see Figure 1). They reveal a positive trend until 1970, ups and down until 1982, and then a negative trend with some fluctuations though and a slightly more negative trend for the last ten or fifteen years. With the exception of the resource price boom 1976-77 price trends are negative after 1970.

FIGURE 1 OVER HERE

[16] Second-order serial correlation is arbitrarily close to zero with p-value of 0.99. But with a coefficient lower than 0.2 it is not relevant anyway (see Roodman 2009a). A random effects estimate shows no cross-section random effects.

Nice results for time trends may break down if time dummies for intercepts are introduced as Cuddington and Urzua (1989) as well as Bleaney and Greenaway (1993) have shown. Therefore we introduce time dummies for the intercepts of the log level equation of low-income countries. If a dummy may undermine a significant trend result it could in principal also improve an insignificant one. Therefore we add time dummies to the result for low-income countries and for the lower-middle-income countries of Table 1, panel (a). The dummies are defined as going always from 1960 to the years mentioned. Similar to Powell's (1991) outlier analysis we find that there are several significant time dummies and not only one. The regression results for low-income countries can be found in Appendix 1, Regression 1 and 2. For the low-income countries we find the following. The dummies go to 1962, 1970, 1974, 1975, 1977, 1981, 1987, 1992, 1995, 1996, 1997 for the fixed effects pooled least squares estimates (see Appendix 1, regression 1). They go to 1970, 1974, 1975, 1977, 1981, 1987, and 1992 for the case of a feasible or estimated Generalized Least Squares (EGLS) estimate that takes into account the heteroscedasticity (see Appendix 1, regression 2). Figure 2 shows the cumulated values of the dummies of low-income countries for each year. The solid line for the EGLS estimate is clearly on average a downward shift over time with some ups and downs, which resembles the negative time trend of column 5 of Table 1 panel (a). The stippled line for the pooled fixed effect estimate, which does not take into account heteroscedasticity, resembles the falling trend less because there are some positive jumps in the 1990s.[17]

FIGURE 2 OVER HERE

For the lower-middle-income countries we find significant dummies for the period from 1960 until the years 1972, 1974, 1975, 1977, 1992, 2007 in the first difference equation (regression output not shown). Figure 3 shows the cumulated value. Since 1993 the dummies

[17] As the time dummies go always from 1960 to some later year the jumps occur when one of the time dummies ends.

indicate a positive intercept, which would equal the long-run growth rate if it were not for the change through the 2007 dummy. It seems clear that there is no resemblance with a negative trend here, whereas before 1992 clearly the opposite seemed to be the case. Of course this may also happen to the low-income countries in the future, for example if they have natural resources for which the prices may start increasing one day.

FIGURE 3 OVER HERE

Giving priority to the EGLS result we find that the dummies crowd out the trend but the overall impression is the same as for the time trend analysis: lower-middle income countries had no falling time trend in the terms of trade for the period under consideration but go up one big step through a time dummy in 1992; low-income countries have falling terms of trade either in a time trend or go downward with ups and downs, with the most serious downward jump indicated by the 1975 dummy, as suggested by Powell's (1991) outlier analysis.

4.5. Country-specific time Trends

Table 3 summarizes the results if countries have a common coefficient of the lagged dependent variable and fixed effects as before but individual time trends. Column 1 shows the number of countries with a significantly negative time trend in each sample. This is largest for the poorest countries, 15; but in percentages of all countries in the respective groups, column 5, the high-income OECD has a larger share. The number of significantly positive trends in column 2 is lowest in low-income countries as a percentage of the total. Insignificant trends are most frequent in all groups except for high-income OECD countries.

Table 3 OVER HERE

When we estimate in first differences instead, only seven low-income countries have significantly negative time trends; three others are just insignificant and when starting the estimation from 1990 onwards there are eleven significantly negative ones. Differencing leads

to an indication of a low number of cases here; this is likely to be generated by the inefficiency implied by over-differencing resulting in too many rejections.[18] The common-unit-root test of Table 1, panel (b), is relevant here under our assumption of a common coefficient of the lagged dependent variable. The results of the upper part of Table 1 are the relevant ones for low-income countries and therefore level results are more plausible in this sub-section. For lower-middle income countries first–differences are relevant and there are only two countries which show falling terms of trade.

Next, we have again added time dummies with common coefficients for the low-income countries to the regressions with country-specific time trends. The result (see Appendix 1, regression 3) is that almost all time dummies are significant. Only four low-income countries keep having a significantly negative time trend. We calculate the intercept for each year through adding up the coefficients of the valid time dummies and the constant as shown in Figure 4. There is clearly a shift down of the whole equation over time indicating falling terms of trade.

FIGURE 4 OVER HERE

4.6. All Coefficients Country specific

Finally, we relax also the last constraint of a common lagged dependent variable and estimate the system of equations (1), which contains forty equations because there are forty countries, using the SUR method (see Appendix 2). We use only one lagged dependent variable. All coefficients of lagged dependent variables are below 0.96 and most much lower. This adds to the information in panel (b) of Table 1, that there are no individual unit roots in low-income countries. The number of low-income countries with significantly negative trends goes from 15 in Table 3 to 19 of 39 countries (49 per cent). There are five significantly positive ones. If

[18] As the problem is low efficiency, one could accept higher significance levels, without guidance how far to go though. We have to go slightly beyond a 30 per cent level to find 15 negative trends as for levels. This indicates the inefficiency.

we exclude countries with less than 15 observations we have 16 of 34 countries with significantly falling terms of trade. If we require at least 20 observations per country we have 15 of 29 with significantly falling terms of trade. If we require at least 30 observations 9 of 21 countries have significantly falling terms of trade. If we estimate the system without constraints using the SUR method again but now for first differences, there are only six countries with significantly negative terms of trade trends. Again over-differencing leads to a low number of cases of falling terms of trade; differencing should only be applied to variables which are integrated of order one.

5. Results for Net Barter Terms of Trade

For the net barter terms of trade the results from estimation with common coefficients on the lagged dependent variable are summarized in Table 4. The sign and significance for the long-run trend are the same as in Table 1 for the low-income countries and more negative and significant for lower-middle income countries. The numerical values may be biased though as we do have only observations for 23 periods for the poor countries. Another reason why the long-term trend is more negative may be that the net barter terms of trade are based only on commodities but not services, which are included in the data for export-as-capacity-to-import/exports used in Tables 1 and 2. When we start the regression only in 1992, the coefficient of the trend is almost the same, but the sum of the coefficients for the lagged dependent variables is smaller and therefore the long-term trend is smaller. Similarly, Ram (2004) found that the trends are more negative before the 1980s (estimating for 1970-1999).

TABLE 4 OVER HERE

For commodities Bleaney and Greenaway (1993) found that the negative trend stems from the period 1980-1992. The stronger growth of African countries since 1990 or 1995 is often attributed to better prices received. All these results together point to the difficulty of

separating trends, volatility, and structural breaks or may indicate that each period may have its own 'trend'. Ocampo and Parra (2010) separate periods through structural break analysis and find a more (rather than less) negative trend after the 1979 break, because the 2003 break separates this period at the other end and thereby takes out the positive trend of the very short last period. Also the results for services, goods and their primary part may be quite different and there are differences in the analysis for different sub-groups of commodities in Ocampo and Parra (2010).

6. The Impact of Food Crisis and Financial Crisis

For this section we use the updated data set going until 2012. We use them in the model without parameter constraints where coefficients for trends and the two lagged dependent variables and constants are allowed to differ by country. The SUR method is used again. Taking data until 2007, just before the 2008 peak in food prices, we find that 17 countries have significantly falling terms of trade. Including 2008 it is also 17 countries. The effect of the food crisis is that Central African Republic gets into the area of significantly falling terms of trade because they import food. The terms of trade make a downward jump of more than 10 per cent in 2008. In contrast, for Guinea the terms of trade jump upward by about one third as it is a net-exporter of agricultural products. Guinea does not return to significantly falling terms of trade after the foods crisis. When the financial crisis hits, using data until 2009, Ethiopia gets out of significantly falling terms of trade and Madagascar gets into this problem. For both it is only a switch in significance of falling terms of trade leaving the number of countries with significantly falling terms of trade at 17. When the financial crisis is over and we use data until 2012 Burundi and Central African Republic change into insignificantly falling terms of trade and Haiti, after having been close to significantly falling terms of trade

during the food crisis, gets into it because oil prices get high again. The results for the trends of the low-income countries are presented in Appendix 3.

Now we have 16 low income countries with significantly falling terms of trade from the new data set going to 2012, whereas we had 19 (or 18 without AFG) with the old data set. But are they the same? Significantly negative trends were found in both data sets for 12 countries. These are Benin, Chad, The Gambia, Ghana, Guinea-Bissau, Malawi, Mali, Mozambique, Myanmar, Tajikistan, Togo, and Uganda.[19] Significantly negative terms of trade only in the first data set were found for Afghanistan,[20] Burundi, Cambodia, Central African Republic, Ethiopia, Guinea, Zambia (six countries), and only in the second for Comoros, Eritrea, Haiti, Madagascar (four countries). Burkina Faso and Niger are always just insignificantly negative in the four regressions of Appendix 3 using the new data set. In total there are at least 24 at risk in the sense of having significantly falling terms of trade in either the old data set or the new one when taking it into account until 2012. There are twenty countries in Appendix 3 which have significantly falling terms of trade in at least one of the sub-samples. In the results until 2012 for the new data set there are five countries with negative sign of the trend and significance level between 10 per cent and 20 per cent. This is in line with our lessons from the crises: Whenever a country gets rid of significantly falling terms of trade a different one gets into the problem.

As the 27 OECD countries (and all other groups but the poor) have individual unit roots, joint estimation may be a bit dangerous, because they are likely to lead to high covariances if unit root also appear in the residuals. Therefore we reestimated without the OECD countries. The result (not shown) is that there are only small differences. Burundi is never in the significantly negative area; Madagascar is always significantly negative, not only when 2009 is included. Guinea Bissau is only significant when 2009 is included and not before. As the

[19] In the case of Ghana and Guinea Bissau old and new data sets are the same as explained in the data section.
[20] Afghanistan was taken out in the new data set.

SUR method is a GLS estimator we can choose between a one-step estimator used so far or we can use simultaneous or sequential updates until the variance-covariance matrix used is sufficiently similar to that one put in. Using simultaneous updates we find again only marginal changes.[21] For data until 2008, Eritrea goes into and Kyrgyz Republic goes out of the set of countries with significantly falling terms of trade. Including 2009, Ethiopia does not go out of falling terms of trade which it did before. Only when including 2012 Ethiopia does go out of falling terms of trade and Laos does go into it, which it never did before in any of the other regressions. When two lags are used, simultaneous updates do not strongly increase the number of cases with high serial correlation, which it did when only one lag was used. The exceptions here are Eritrea, Laos and Tajikistan. Overall, serial correlation is still stronger under simultaneous updates and therefore estimates are less reliable.

In all regressions with the updated data set there are always 16 or 17 countries with significantly falling terms of trade. Which countries these are depends on the real world situations like the food and financial crises and also on the methods chosen, including or not including OECD countries and using or not using simultaneous updates; finally, the number of observations matter in the sense of more observations lead to more significance – countries with less than 21 observations do not appear on the list of 12 countries with falling terms of trade in both samples.

7. Interpretation and Conclusion

Our interpretation of these results is that 10 of the twenty-seven high-income OECD countries are passing on more of technical change to their customer countries than they get either through technical change of trading partners as suggested by Kravis (1970) or through their high demand growth, whereas the majority has no significantly falling terms of trade before

[21] When two lags are used, simultaneous updates do not strongly increase the number of cases with high serial correlation, which it did when only one lag was used. The exceptions here are Eritrea, Laos and Tajikistan.

the financial crisis. We rely here on the results for first differences as we find panel unit roots for the high-income OECD countries. For low-income countries though there are no indications of unit roots and we should not rely on results in first differences but rather on those for levels. Assuming that the low-income countries have hardly any technical progress, the fall in the terms of trade by about 0.4 per cent according to Table 1 might be due to a lack of growth of export demand, which reduces the growth of imported investment goods as suggested by Prebisch (1950/1962, p.2). Combining first differences and level in the system GMM estimation in order to take care of endogeneity leads to a falling trend too, with values between those from levels and over-differencing. As three low-income countries have significantly positive trends in the old data set and nine in the new one they probably have strong export demand growth relative to the technical change. Applying time dummies to these results makes the time trend variable insignificant, but the calculated intercepts per period also indicate falling terms of trade for the low-income country. In no version of our model do we find falling terms of trade for lower-middle income countries. Using the SUR method and no constraints on the parameters as in equation (1) falling terms of trade are the case for almost 50 per cent of the low-income countries for the older data set and slightly less for the new data set. Estimation in first differences without good reason to do so, leads to inefficiency through over-differencing implying too many rejections and therefore a much lower number of low-income countries with falling terms of trade.

The new data set is suffering from many data problems stemming from having only the old data for several countries. When taking the full data set the results obtained with the old one are nevertheless by and large confirmed: 16 countries have significantly negative terms of trade trends and an additional five come close to significance. Over all the sub periods, 20 countries have significantly falling terms of trade in one of the sub-samples ending alternatively 2007, 2008 (including the food crisis), 2009 (including the worst year of the

financial crisis) and 2012 including all data available. The effect of each of the crises is that price shocks let one or two countries go into or out of the set of countries with significantly negative terms of trade.

There are two common counterarguments in regard to the falling terms of trade results. The first refers to transport costs. Import prices contain cost, insurance and freight (CIF) but export prices are 'free on board' (fob) prices. The stronger the technical change in transport if passed on in transport prices the lower the trend in import price indices.[22] But we have no empirical indication for this for the time under consideration.[23] If anything this biases the trend for country terms of trade upward.

The second common argument is unmeasured trends in quality of goods. It could affect both, price indices of imports and exports, both of which contain raw materials, manufactures and services; here it is also important that many countries have falling shares of raw materials and multinationals are active worldwide and produce quality improvements in all goods, in particular those traded two ways in the vein of global production chains.[24] If manufactures have more quality change LDC import price increases may be overestimated, but the increasing share of manufactures in exports would do the same. It is hard to imagine how relative prices could plausibly become constant through these quality corrections. Moreover, these days many products improve quality without increasing the prices. Indeed, Saadi (2012) finds that an increase in the sophistication - captured by indices that contain implicitly productivity and quality measures - of the developing countries' exports is accompanied by a deterioration of their terms of trade. Then it is hard to have any theory of the value of quality and the appropriate price correction, which normally is based on the idea that increasing

[22] Data on c.i.f./f.o.b. factors are no longer published by the IMF.
[23] Kaplinsky (2006) reports that China has driven up demand and prices for transport services recently. But this is only a short episode to which capacity and supply may adjust.
[24] Wacker (2011) finds a positive impact of multinationals on the terms of trade. This would imply that some measurable impact of multinationals is already captured by the data. But it is hard to link this to the question which quality correction should be applied to the price indices.

quality has caused costs passed on into prices. But price theory is mostly not that simple. In particular, the process improving quality may have technical progress itself, which may ensure that there is no price increase when quality is improved. The forces of asymmetric technical change in processes and income elasticities of export demand would still be in existence if quality were correctly taken into account. Therefore we think that the step from the analysis of trends in commodity terms of trade as initiated by Prebisch and Singer and manufactures as initiated by Keesing (1979) and Sarkar and Singer (1991) to country terms of trade is an important one.

Goods may not only change quality but also variety may be enhanced. In love-of-variety models, resource increases such as the population growth of poor countries increase the number of goods produced. Price indices have positive love-of-variety effects, where in the presence of trade costs the home number of goods has a higher weight. Thus population growth may increase export prices indices through the love-of-variety effect. As poor countries suffer from a lack of diversification this effect seems to be less relevant for them and these models are mostly applied to OECD countries. Moreover, in empirical price indices the number of goods is kept constant and therefore the love-of-variety effect is not included in the data we have analysed (Corsetti et al. 2007; in particular footnote 14). Blanchard and Kiyotaki (1987) eliminate the love-of-variety effect from the model in a way that it does not appear in the rate of inflation because they want to consider the damaging part of inflation. Love-of-variety indicates a welfare effect from a special form of a utility function but note that you can import more under higher terms of trade. This type of effects is fascinating but less relevant for poor countries, but important for emerging markets.

In models with a continuum of goods with different productivities in their production function (Young 1991, Gustafsson and Segerstrom 2010) the broadening of variety leads to the inclusion of less productive goods. This results in a cost-push effect improving the terms

of trade. However, again diversification is not a strength of poor countries and all enhancements of variety may be larger in developed and emerging countries because not only differences in labour endowment growth are important, but also the growth of human and physical capital and productivities, which is larger in the major trading partners, the OECD countries, than in poor developing countries and likely to dominate.

We have given only an intuitive interpretation of the results. More elaborate theorizing is possible but not the intention of this paper. A good model must be able to explain positive and negative trends and should take into account elements that are included by relatively successful closed economy growth models – savings, investment, labour growth and technical change. The preferred elements to be added to a closed economy growth model are exports and imported capital goods as in the model of Bardhan and S.Lewis (1970) a variant of which can be estimated (see Mutz and Ziesemer 2008; Habiyaremye and Ziesemer 2012).[25] This type of model has the property that investment and GDP per capita growth are both positively related to the terms of trade as found in the evidence of Bleaney and Greenaway (2001) and Jawaid and Raza (2012) for India.[26] For the countries with falling (increasing) terms of trade technical change is a stronger (weaker) driving force than world income growth multiplied by the income elasticity of export demand. Only if both forces are equally strong results will equal those of closed economy growth. The results suggest that being richer makes the problem of falling terms of trade less severe because they originate from strong technical change, which increases GDP per capita growth.

[25] These models differ from the application of terms of trade in current standard macroeconomics as in Broda (2004) in three ways: First, price levels are related to terms of trade without any lag, because the underlying goods prices come from a set of goods which are in the GDP and in the exports, whereas the VAR culture not only assumes lags but often also excludes contemporaneous relations, which is taken into account in the more general ARDL models; second, estimated income and price elasticities of export demand from these models differ from country to country, implying a heterogeneity that is absent under the homogeneity assumption of panel cointegration tests used in the macro literature (but also one of our regressions); third, terms of trade are endogenous, because estimated price elasticities are almost never minus infinity, whereas some macroeconomic literature assumes exogenous terms of trade in combination with a VAR where they are endogenous by construct. Broda (2004) explains very well that these assumptions create a hopefully small bias and are simplifiers to avoid econometric complications in his work.

[26] The surveys of Ziesemer (1995) and Sarkar (2001) discuss some other interesting models.

Poor countries may have more favourable terms of trade development if they have a lower share of products with low-income elasticities of demand. This is probably more likely the more countries are diversified. However, diversification policies at each level of growth may or may not mitigate or even avoid falling terms of trade. Athukorala (2000) finds a positive effect for Sri Lanka, but Sarkar (2004) finds no effect of changes in diversification indices on changes in terms of trade for 24 diversifying countries in panel average of the 1980s. Chakraborty (2012) shows that in the past this has at least not happened sufficiently strongly: manufactures of developing countries have falling terms of trade vis-à-vis those of developed countries. Ocampo and Parra (2006) point out though that this is less of a problem because demand for manufactures has high growth. They favour industrialization as a strategy. With the competition from China this has become more difficult though. The best choice of sectors may be one within rather than between the groups of primaries, industrial goods or services. For services again – as with the diversification into manufactures - there will be the question whether the poor countries attract those with increasing or those with decreasing terms of trade. And if they attract those with falling terms of trade the question again will be is this because demand is growing more slowly than supply or is technical change merely passing on cost reductions in falling prices and attracting high growth of export quantities under high price elasticities or low ones under low price elasticities.

Diversification alone is not enough to increase the growth rate of the terms of trade. Only if the income elasticity of export demand is increased more than the rate of change of labour productivity diversification can increase the growth rate of the terms of trade. To achieve this, growth enhancing factors have to join in, because according to all experience protectionist policies do not work (Sarkar and Sarkar 2008) in regard to growth. Moreover, besides limiting imports the induced trade partner’s protectionism reduces the possibility to diversify exports (Mansfield and Reinhardt 2008). Infrastructure and education are likely to be helpful to get

more diversification (see Habiyaremye and Ziesemer 2006) for two reasons: first, larger endowments can carry more sectors with fixed costs as in some models of imperfect competition with endogenous numbers of goods; second, if endowments are more similar trade will be less of the inter-industry type and more of the intra-industry type. Therefore it is tempting to speculate that some of the variables that also support growth and diversification (Cadot et al. 2013) will help stopping the terms of trade from falling through actively changing trade patterns through education and public investment (see Ziesemer 2003 for a theoretical example).[27] Moreover, whatever the results for the past were, diversification should offer some insurance for low-income countries against being hit by both, rising prices of imported goods and falling prices of exported goods. An insurance property will be stronger if there is more diversification than merely going from having one product or sector to still only a very few as in the cases of Sri Lanka, Botswana, Ghana, and Mauritius.[28] But there will clearly be limits to this because policies and globalization cannot do away with at least two important aspects (Mollick et al. 2008): first, goods will remain having different income elasticities of export demand; second, endowments will remain unequal across countries. In particular, natural endowments underlying the trade in mining products will remain distributed unequally across countries. But these aspects imply that terms of trade may remain having bleak prospects for commodities, although the opposite holds since 2003, whereas those for countries can still be mitigated by the policies indicated above by way of reducing the share of goods with low income elasticities of a country's GDP and trade volume as it is the case for rich countries with rich endowments of natural resources like Canada, Australia and the USA.

[27] Changing comparative advantage through investment should carefully be distinguished from policies merely biasing against comparative advantage, which could easily reduce growth and diversification (Parteka and Tamberi 2013).
[28] See Athukorala (2000) on Sri Lanka, Habiyaremye (2013) on Botswana, and Habiyaremye and Ziesemer (2012) on Mauritius.

Although there are many open questions regarding diversification, probably encouraged by falling terms of trade (UN 2008, p.64) and other policies, we hope to have shown though that the problem of falling terms of trade continues to exist for many countries, especially the poor ones, also when taking into account unit roots, endogeneity, dummies, and heterogeneity.

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Appendix 1: Regression output with time dummies

Regression 1: Time dummies instead of common trend for low-income countries

Dependent Variable: (LOG(ECM/EX)). Method: Pooled Least Squares
Sample (adjusted): 1962 2008 Included observations: 47 after adjustments
Cross-sections included: 40 (low income countries)
Total pool (unbalanced) observations: 1238

Period SUR (PCSE) standard errors & covariance (d.f. corrected)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.023	0.007	-3.279	0.001
LOG(ECM(-1)/EX(-1))	0.798	0.030	26.308	0.000
LOG(ECM(-2)/EX(-2))	0.042	0.030	1.432	0.152
DUM6062	-0.038	0.021	-1.825	0.068
DUM6070	0.061	0.026	2.321	0.021
DUM6074	0.123	0.057	2.150	0.032
DUM6075	-0.211	0.067	-3.126	0.002
DUM6077	0.101	0.050	2.018	0.044
DUM6081	-0.037	0.021	-1.752	0.080
DUM6087	0.068	0.019	3.514	0.001
DUM6092	-0.072	0.025	-2.856	0.004
DUM6095	0.075	0.026	2.867	0.004
DUM6096	-0.073	0.028	-2.605	0.009
DUM6097	0.039	0.019	2.035	0.042

Fixed Effects (Cross)

AFG--C	-0.135	MDG--C	-0.009
BGD--C	0.014	MWI--C	0.092
BEN--C	0.002	MLI--C	-0.023
BFA--C	0.014	MRT--C	-0.001
BDI--C	-0.035	MOZ--C	0.051
KHM--C	0.018	MMR--C	0.090
CAF--C	-0.047	NER--C	-0.089
TCD--C	0.022	RWA--C	-0.073
COM--C	0.007	SEN--C	-0.009
ZAR--C	0.011	SLE--C	0.014
ERI--C	-0.017	SOM--C	-0.042
ETH--C	0.013	TJK--C	-0.091
GMB--C	0.004	TZA--C	0.043
GHA--C	-0.051	TGO--C	-0.032
GIN--C	0.013	UGA--C	0.038
GNB--C	-0.091	UZB--C	0.053
HTI--C	-0.005	VNM--C	0.014
KEN--C	-0.012	YEM--C	0.015
KGZ--C	-0.020	ZMB--C	0.076
LAO--C	0.040	ZWE--C	0.004

Effects Specification: Cross-section fixed (dummy variables)

R-squared	0.874	Mean dependent var	0.028
Adjusted R-squared	0.869	S.D. dependent var	0.417
S.E. of regression	0.151	Akaike info criterion	-0.898
Sum squared resid	27.100	Schwarz criterion	-0.679
Log likelihood	609.0	Hannan-Quinn criter.	-0.816
F-statistic	158.2	Durbin-Watson stat	2.044
Prob(F-statistic)	0		

Regression 2: Time dummies instead of common trend for low-income countries (EGLS)

Dependent Variable: (LOG(ECM/EX)) Method: Pooled EGLS (Cross-section weights)
Sample (adjusted): 1962 2008 Included observations: 47 after adjustments
Cross-sections included: 40 (low income countries)
Total pool (unbalanced) observations: 1238 Linear estimation after one-step weighting matrix
Cross-section weights (PCSE) standard errors & covariance (d.f. corrected)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.016	0.005	-3.610	0.0003
LOG(ECM(-1)/EX(-1))	0.788	0.026	30.277	0.0000
LOG(ECM(-2)/EX(-2))	0.046	0.026	1.804	0.0715
DUM6070	0.042	0.016	2.578	0.0101
DUM6074	0.072	0.029	2.438	0.0149
DUM6075	-0.135	0.032	-4.190	0.0000
DUM6077	0.090	0.022	3.999	0.0001
DUM6081	-0.041	0.015	-2.701	0.0070
DUM6087	0.036	0.011	3.187	0.0015
DUM6092	-0.010	0.006	-1.625	0.1044

Fixed Effects (Cross)

AFG--C	-0.135	MDG--C	-0.009
BGD--C	0.016	MWI--C	0.097
BEN--C	0.003	MLI--C	-0.025
BFA--C	0.015	MRT--C	0.000
BDI--C	-0.034	MOZ--C	0.051
KHM--C	0.015	MMR--C	0.096
CAF--C	-0.056	NER--C	-0.090
TCD--C	0.024	RWA--C	-0.075
COM--C	0.005	SEN--C	-0.009
ZAR--C	0.011	SLE--C	0.015
ERI--C	-0.019	SOM--C	-0.041
ETH--C	0.011	TJK--C	-0.093
GMB--C	0.004	TZA--C	0.044
GHA--C	-0.053	TGO--C	-0.032
GIN--C	0.010	UGA--C	0.038
GNB--C	-0.096	UZB--C	0.047
HTI--C	-0.004	VNM--C	0.012
KEN--C	-0.012	YEM--C	0.017
KGZ--C	-0.022	ZMB--C	0.080
LAO--C	0.033	ZWE--C	0.003

Effects Specification: Cross-section fixed (dummy variables)

Statistics			
R-squared	0.878	Mean dependent var	0.045
Adjusted R-squared	0.873	S.D. dependent var	0.421
S.E. of regression	0.150	Sum squared resid	26.85
F-statistic	178.8	Durbin-Watson stat	1.981
Prob(F-statistic)	0		

Unweighted Statistics			
R-squared	0.872	Mean dependent var	0.028
Sum squared resid	27.56	Durbin-Watson stat	2.036

Regression 3: Time dummies with country specific trends and fixed effects for low-income countries

Dependent Variable: (LOG(ECM/EX)) Method: Pooled Least Squares
Sample (adjusted): 1961 2008 Included observations: 48 after adjustments
Cross-sections included: 40 (low income countries). Total pool (unbalanced) observations: 1278
Cross-section weights (PCSE) standard errors & covariance (d.f. corrected)

Variable	Coefficient	Prob.		Coefficient	Prob.	Fixed Effects (Cross)	
C	-0.192	0.061	AFG--@TREND	-0.532	0.000	AFG--C	23.89
LOG(ECM(-1)/EX(-1))	0.724	0.000	BGD--@TREND	-0.004	0.072	BGD--C	0.205
DUM6007	0.037	0.000	BEN--@TREND	0.007	0.001	BEN--C	-0.128
DUM6005	-0.021	0.000	BFA--@TREND	0.006	0.010	BFA--C	-0.062
DUM6003	0.019	0.001	BDI--@TREND	0.005	0.047	BDI--C	-0.129
DUM6002	0.014	0.008	KHM--@TREND	0.004	0.053	KHM--C	-0.020
DUM6001	-0.037	0.000	CAF--@TREND	0.004	0.085	CAF--C	-0.105
DUM6000	0.028	0.000	TCD--@TREND	0.004	0.045	TCD--C	-0.021
DUM6099	0.021	0.000	COM--@TREND	0.006	0.005	COM--C	-0.112
DUM6098	-0.014	0.016	ZAR--@TREND	0.012	0.000	ZAR--C	-0.251
DUM6097	0.043	0.000	ERI--@TREND	0.001	0.630	ERI--C	0.039
DUM6096	-0.070	0.000	ETH--@TREND	0.000	0.846	ETH--C	0.142
DUM6095	0.134	0.000	GMB--@TREND	0.002	0.314	GMB--C	0.017
DUM6094	-0.104	0.000	GHA--@TREND	0.005	0.027	GHA--C	-0.164
DUM6093	0.028	0.000	GIN--@TREND	-0.003	0.221	GIN--C	0.243
DUM6092	-0.080	0.000	GNB--@TREND	0.001	0.741	GNB--C	-0.103
DUM6091	0.066	0.000	HTI--@TREND	0.011	0.000	HTI--C	-0.212
DUM6090	0.022	0.003	KEN--@TREND	0.007	0.001	KEN--C	-0.154
DUM6089	-0.027	0.000	KGZ--@TREND	0.009	0.000	KGZ--C	-0.281
DUM6088	0.104	0.000	LAO--@TREND	0.012	0.000	LAO--C	-0.368
DUM6087	-0.070	0.000	MDG--@TREND	0.006	0.008	MDG--C	-0.116
DUM6086	0.069	0.000	MWI--@TREND	0.004	0.082	MWI--C	0.104
DUM6083	0.021	0.006	MLI--@TREND	0.005	0.028	MLI--C	-0.103
DUM6082	-0.030	0.000	MRT--@TREND	0.007	0.001	MRT--C	-0.141
DUM6081	-0.033	0.000	MOZ--@TREND	0.000	0.984	MOZ--C	0.202
DUM6080	0.051	0.000	MMR--@TREND	-0.003	0.188	MMR--C	0.272
DUM6079	0.026	0.003	NER--@TREND	0.008	0.000	NER--C	-0.296
DUM6078	-0.047	0.000	RWA--@TREND	0.012	0.000	RWA--C	-0.370
DUM6077	0.146	0.000	SEN--@TREND	0.007	0.001	SEN--C	-0.155
DUM6076	-0.053	0.000	SLE--@TREND	0.015	0.000	SLE--C	-0.298
DUM6075	-0.162	0.000	SOM--@TREND	0.013	0.000	SOM--C	-0.285
DUM6074	0.162	0.000	TJK--@TREND	-0.022	0.000	TJK--C	0.897
DUM6072	-0.022	0.008	TZA--@TREND	0.005	0.016	TZA--C	-0.020
DUM6071	-0.022	0.016	TGO--@TREND	0.003	0.146	TGO--C	-0.086
DUM6070	0.117	0.000	UGA--@TREND	-0.004	0.044	UGA--C	0.351
DUM6067	0.014	0.108	UZB--@TREND	0.014	0.000	UZB--C	-0.380
DUM6066	0.015	0.084	VNM--@TREND	0.005	0.011	VNM--C	-0.074
DUM6063	0.032	0.001	YEM--@TREND	0.002	0.337	YEM--C	0.061
DUM6062	-0.054	0.000	ZMB--@TREND	0.003	0.150	ZMB--C	0.096
DUM6061	0.036	0.000	ZWE--@TREND	0.008	0.000	ZWE--C	-0.159

Effects Specification: Cross-section fixed (dummy variables)

R-squared	0.887	Mean dependent var	0.034	S.D. dependent var	0.421
Adjusted R-squared	0.875	Akaike info criterion	-0.884	Prob(F-statistic)	0
S.E. of regression	0.149	Schwarz criterion	-0.404	F-statistic	76.86
Sum squared resid	25.7	Hannan-Quinn criter.	-0.704	Durbin-Watson stat	2.008
Log likelihood	684.0				

Appendix 2 Simultaneous equation estimation with full heterogeneity

Regression results for equation (1) with country-specific coefficients for trends, constants and lagged dependent variables

Country	*lag.dep.var.*	*p-value*	*trend*	*p-value*	*constant*	*p-value*	*Obs.*	*adj.R2*	*DW (a)*
AFG	-0.168	0.000	-0.255	0.000	12.050	0.000	3	1.00	1.54
BGD	0.953	0.000	0.000	0.928	-0.013	0.888	48	0.94	1.56
BEN	0.399	0.000	-0.002	0.016	0.073	0.007	45	0.36	1.93
BFA	0.754	0.000	-0.001	0.317	0.068	0.119	41	0.63	1.88
BDI	0.525	0.000	-0.006	0.080	0.066	0.375	36	0.35	1.98
KHM	-0.012	0.949	-0.002	0.081	0.081	0.045	14	0.00	1.82
CAF	0.010	0.965	-0.028	0.023	1.016	0.048	8	0.49	1.93
TCD	0.386	0.000	-0.008	0.000	0.307	0.000	48	0.82	2.04
COM	0.183	0.108	-0.004	0.113	0.124	0.170	28	0.02	1.80
ZAR	0.493	0.000	0.008	0.000	-0.208	0.000	48	0.70	1.67
ERI	0.703	0.000	-0.001	0.883	-0.011	0.972	15	0.51	1.47
ETH	0.630	0.000	-0.009	0.001	0.319	0.001	27	0.72	1.98
GMB	0.550	0.000	-0.008	0.000	0.261	0.000	42	0.78	1.56
GHA	0.510	0.000	-0.006	0.001	0.004	0.916	48	0.64	1.87
GIN	0.578	0.000	-0.009	0.012	0.374	0.015	22	0.91	1.35
GNB	0.675	0.000	-0.005	0.095	-0.032	0.677	37	0.65	1.98
HTI	0.375	0.003	0.004	0.093	-0.086	0.054	33	0.09	1.76
KEN	0.956	0.000	0.001	0.353	-0.023	0.351	48	0.79	1.92
KGZ	0.737	0.000	0.003	0.574	-0.164	0.437	15	0.55	2.00
LAO	0.537	0.048	0.010	0.026	-0.436	0.033	9	0.05	1.79
MDG	0.827	0.000	-0.001	0.290	0.016	0.568	48	0.80	1.78
MWI	0.745	0.000	-0.004	0.018	0.237	0.002	48	0.81	1.77
MLI	0.555	0.000	-0.003	0.002	0.042	0.127	40	0.64	1.54
MRT	0.640	0.000	-0.001	0.427	0.037	0.366	45	0.52	1.74
MOZ	0.675	0.000	-0.008	0.003	0.389	0.002	28	0.75	1.85
MMR	0.731	0.000	-0.011	0.000	0.421	0.000	44	0.95	1.84
NER	0.874	0.000	-0.001	0.471	-0.041	0.457	39	0.73	2.12
RWA	0.413	0.000	0.007	0.006	-0.424	0.000	39	0.35	1.96
SEN	0.685	0.000	-0.001	0.483	0.004	0.866	48	0.62	1.73
SLE	0.418	0.001	0.010	0.050	-0.249	0.044	28	0.33	1.39
SOM	0.656	0.000	0.005	0.071	-0.143	0.021	29	0.63	1.49
TJK	0.717	0.000	-0.033	0.000	1.190	0.000	18	0.72	1.97
TZA	0.385	0.011	0.003	0.495	-0.027	0.875	16	0.11	2.05
TGO	0.168	0.105	-0.015	0.000	0.268	0.000	45	0.59	1.98
UGA	0.637	0.000	-0.014	0.001	0.599	0.001	26	0.77	1.17
UZB	0.441	0.002	0.016	0.003	-0.623	0.005	14	0.68	1.12
VNM	0.389	0.011	0.003	0.150	-0.135	0.136	18	0.20	1.51
YEM	-0.055	0.770	0.000	0.151	0.000	0.101	13	-0.06	2.10
ZMB	0.724	0.000	-0.005	0.040	0.258	0.002	48	0.74	1.69
ZWE	0.697	0.000	0.001	0.602	-0.027	0.489	29	0.54	1.10

(a) Estimation Method: Seemingly Unrelated Regression with one equation per country. Sample: 1961-2008. Countries: 40. Periods: 48. Total system (unbalanced) observations 1278. Linear estimation after one-step weighting matrix.

(b) Durbin-Watson statistic; in case of endogeneity it is only indicative. It is not used for an exact test here; that would require a Breusch-Godfrey test.

Appendix 3	**Simultaneous equation estimation with full heterogeneity**									
	Regression results with country-specific coefficients for trends									
	1962-2007		1962 - 2008		1962 - 2009		1962 - 2012		1962 - 2012	
	Trend	p-value	Trend	p-value	Trend	p-value	Trend	p-value	Adj. R^2	DW stat
	coefficient		coefficient		coefficient		coefficient			
BGD	-0.002	0.520	-0.002	0.443	-0.002	0.340	-0.002	0.348	0.94	1.73
BEN	**-0.006**	**0.000**	**-0.006**	**0.000**	**-0.006**	**0.000**	**-0.006**	**0.000**	0.81	1.90
BFA	-0.001	0.416	-0.001	0.302	-0.001	0.329	-0.001	0.165	0.57	1.93
BDI	**-0.005**	**0.022**	**-0.005**	**0.026**	**-0.005**	**0.025**	-0.003	0.156	0.31	1.97
KHM	-0.001	0.257	0.003	0.100	0.003	0.082	0.003	0.012	-0.06	2.00
CAF	-0.007	0.707	**-0.046**	**0.000**	**-0.039**	**0.005**	-0.001	0.916	-0.41	2.26
TCD	**-0.007**	**0.000**	**-0.007**	**0.000**	**-0.008**	**0.000**	**-0.008**	**0.000**	0.82	1.87
COM	**-0.031**	**0.000**	**-0.031**	**0.000**	**-0.032**	**0.000**	**-0.032**	**0.000**	0.26	2.21
ZAR	0.005	0.007	0.004	0.012	0.004	0.012	0.003	0.035	0.33	1.88
ERI	**-0.014**	**0.001**	**-0.015**	**0.000**	**-0.015**	**0.000**	**-0.013**	**0.001**	0.97	2.05
ETH	**-0.006**	**0.037**	**-0.006**	**0.048**	-0.003	0.323	0.001	0.532	0.57	1.84
GMB	**-0.008**	**0.000**	**-0.008**	**0.000**	**-0.008**	**0.000**	**-0.008**	**0.000**	0.80	1.92
GHA	**-0.006**	**0.000**	**-0.006**	**0.000**	**-0.006**	**0.000**	**-0.006**	**0.000**	0.62	1.93
GIN	**-0.008**	**0.020**	0.016	0.000	0.009	0.000	0.006	0.002	0.77	1.31
GNB	**-0.008**	**0.012**	**-0.008**	**0.008**	**-0.007**	**0.017**	**-0.007**	**0.023**	0.63	1.89
HTI	0.003	0.632	-0.009	0.131	-0.005	0.357	**-0.013**	**0.003**	0.63	2.21
KEN	0.000	0.928	0.000	0.893	0.000	0.807	0.000	0.916	0.78	1.98
KGZ	0.005	0.331	0.004	0.365	0.003	0.475	0.001	0.741	0.38	2.14
LAO	-0.012	0.360	0.009	0.319	0.009	0.341	0.008	0.417	-0.21	1.48
MDG	-0.002	0.148	-0.001	0.143	**-0.002**	**0.088**	**-0.002**	**0.047**	0.80	1.85
MWI	**-0.011**	**0.000**	**-0.011**	**0.000**	**-0.011**	**0.000**	**-0.011**	**0.000**	0.90	1.99
MLI	**-0.002**	**0.013**	**-0.002**	**0.012**	**-0.003**	**0.006**	**-0.003**	**0.008**	0.63	1.92
MRT	-0.002	0.116	0.000	0.781	0.001	0.215	0.001	0.137	0.80	1.82
MOZ	**-0.009**	**0.002**	**-0.009**	**0.002**	**-0.009**	**0.001**	**-0.008**	**0.002**	0.87	1.90
MMR	**-0.012**	**0.000**	**-0.012**	**0.000**	**-0.012**	**0.000**	**-0.012**	**0.000**	0.95	1.85
NER	-0.002	0.134	-0.002	0.169	-0.002	0.130	-0.002	0.166	0.73	1.99
RWA	0.007	0.000	0.007	0.000	0.007	0.000	0.008	0.000	0.63	1.89
SEN	-0.001	0.455	-0.001	0.410	-0.001	0.361	-0.001	0.125	0.58	1.84
SLE	0.000	0.871	0.000	0.955	0.000	0.877	0.000	0.959	0.41	1.58
SOM	0.009	0.001	0.009	0.001	0.008	0.001	0.008	0.002	0.66	1.77
TJK	**-0.022**	**0.005**	**-0.022**	**0.004**	**-0.022**	**0.007**	**-0.021**	**0.006**	0.83	1.72
TZA	0.030	0.000	0.028	0.000	0.023	0.000	0.020	0.000	0.86	1.53
TGO	**-0.011**	**0.000**	**-0.011**	**0.000**	**-0.012**	**0.000**	**-0.011**	**0.000**	0.66	1.90
UGA	**-0.012**	**0.006**	**-0.014**	**0.001**	**-0.013**	**0.001**	**-0.012**	**0.000**	0.81	2.15
UZB	0.010	0.062	0.015	0.020	0.015	0.017	0.016	0.015	-0.03	1.99
VNM	0.007	0.000	0.007	0.000	0.007	0.000	0.008	0.000	0.73	1.27
ZMB	-0.002	0.411	-0.002	0.474	-0.002	0.382	0.001	0.754	0.72	1.91
ZWE	0.002	0.121	0.002	0.098	0.003	0.013	0.002	0.010	0.70	2.14
Det. Res. Cov.		5.60E-187		3.60E-185		3.20E-184		1.30E-182		
Countries		38 low-income, 27 high-income OECD								
Periods		46		47		48		49		
Observations		2325		2382		2436		2569		
Linear estimation after 1-step weight. matrix. Simultaneous updating leads to serial correlat.										
Results for constants, lagged dependent variables and OECD countries not shown.										
Results in bold show significantly negative trends.										

Table 1					
Common trend in panels of exports-as-capacity-to-import/exports					
with fixed effects and lagged dependent variables (a)					
Panel (a)					
Estimation in levels (a)					
Income group	High income OECD	High income Non-OECD	Upper middle income	Lower middle income	Low income
Constant	0.0020	0.0001	0.0063	0.0030	0.0125
(p-value)	(0.27)	(0.99)	(0.03)	(0.01)	(0.11)
coeff.lag.dep.(-1)	1.05	1.03	0.91	0.89	0.80
(p-value)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
coeff.lag.dep.(-2)	-0.151	-0.140	-0.064	-0.034	0.049
(p-value)	(0.00)	(0.03)	(0.06)	(0.26)	(0.06)
Coeff. Trend	-0.00009	0.00026	-0.00004	-0.00003	-0.00064
(p-value)	(0.21)	(0.30)	(0.70)	(0.54)	(0.02)
long-run coeff (b)	-0.0009	0.0023	-0.0003	-0.0002	-0.0042
Adj.R^2	0.916	0.95	0.834	0.831	0.874
DW (c)	1.94	1.89	1.90	1.98	1.99
Number of countries	27	19	39	44	40
Total observations	1152	338	1110	1420	1238
Prob. fixed effects redundant (F-stat.)	0.77	0.00	0.14	0.04	0.00
Period	1962-2008	1962-2008	1962-2008	1962-2008	1962-2008
(a) Dependent variable: LOG(ECM/EX). Method:Fixed effects. Pooled EGLS;PCSE: Period SUR					
(b) Coefficient of trend divided by (1- sum of coefficients of lagged dependent variables).					
This value is the stable growth rate to which the system converges.					
(c) Durbin-Watson statistic. Although it is not the adequate statistic for rigorous tests under endogeneity, its size indicates that there can be no serious serial correlation bias. See Epple and McCallum 2006.					
(d) F-statistic					
Panel (b) Unit roots					
Income group	High income OECD	High income Non-OECD	Upper middle income	Lower middle income	Low income
Test\p-values					
Null: Unit root (assumes common unit root process)					
Levin, Lin & Chu t*	0.58	0.0014	0.62	0.66	0.03
Breitung t-stat	0.03	0.98	0.998	0.98	0.00
Null: Unit root (assumes individual unit root process)					
Im, Pesaran and Shin W-stat	0.060	0.264	0.0001	0.137	0
ADF - Fisher Chi-square	0.008	0.093	0	0.002	0
PP - Fisher Chi-square	0.122	0.005	0.0001	0.001	0.0001
Observations	1141-1179	294-334	1035-1128	1337-1423	1184-1275
Countries	27	18	38	41	39

Table 1 (continued)					
Panel (c)					
Estimation in first differences					
Income group	High income OECD	High income Non-OECD	Upper middle income	Lower middle income (c)	Low income (d)
Constant	-0.000065	0.0108	0.0056	-0.0007	-0.0175
(p-value)	(0.75)	0	(0.00)	(0.86)	(0.00)
coeff.lag.dep.(-1)	0.144	0.114979	-0.02	-0.13	-0.18
(p-value)	0.00	0.1	(0.56)	(0.00)	(0.00)
coeff.lag.dep.(-2)	-0.198	-0.225	-0.110	-0.138	-0.157
(p-value)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
long-run coeff (b)	-0.0001	0.0097	0.0050	-0.00062	-0.0148
Adj.R^2	0.058	0.10	-0.003	0.012	0.032
DW (a)	2.02	2.04	1.93	1.95	2.03
Number of countries	27	19	39	43	39
Period	1963-2008	1963-2008	1963-2008	1966-2008	1970-2008
Total observations	1125	318	1071	1240	987
Prob. fixed effects redundant (b)	0.47	0.37	0.93	0.56	0.76
Dependent Variable: D(LOG(ECM/EX))	Method: Pooled EGLS (Cross-section weights) for first three regressions; panel least squares first the last two. PCSE: Period SUR.				

(a) Durbin-Watson statistic. Although it is not the adequate statistic for rigorous tests under endogeneity, its size indicates that there can be no serious serial correlation bias. See Epple and McCallum 2006.

(b) F-statistic. Dropping fixed effects changes neither the sign nor the size of the coefficients or p-values strongly.

(c) There are three more lags included, because lag five is still significant, and serial correlation should be minimized.

(d) There are four more lags included, because lag six is still significant, and serial correlation should be minimized.

Table 2	**A system GMM estimate dealing with endogeneity**		
(a)	*Fixed effects (b)*	*OLS (c)*	*System GMM (d)*
constant	0.044	0.0085	-
	(0.0012)	(0.37)	-
LOG(ECM(-1)/EX(-1))	0.825	0.915	0.842
	(0.0000)	(0.00)	(0.00)
trend	-0.0017	-0.0005	-0.0014
	(0.0003)	(0.09)	(0.01)
long-run trend	-0.010	-0.006	-0.009
Countries	40	40	39
Periods	1961-2008	1961-2008	1964-2008
Observations	1278	1278	1158
s.e. of regression	0.157	0.159	0.156
s.e. of fixed effects	0.066	-	0.062
variance ratio (e)	0.178	-	0.156

Source: Author's estimates

(a) For all three regressions, p-values in parentheses and PCSE Period SUR.

(b) Panel least squares with fixed effects, equivalent to the least-squares dummy variable estimator or within estimator. Adj. R sq.: 0.861. Durbin-Watson stat.: 2.07

(c) Ordinary least squares. Adj. R-sq.: 0.858. Durbin-Watson: 2.15

(d) Fixed effects with orthogonal deviations of Arellano and Bover (1995). Instrument rank: 3. J-statistic: 1.14. Sargan p-value: 0.285. Instrument specification: LOG(ECM(-2)/EX(-2)) LOG(ECM(-3)/EX(-3)), trend. The Sargan-difference test for the last lag is the same as the Sargan test, because an instrument rank of 2 equal to the number of estimated coefficients yields a J-statistic of zero.

(e) Variance ratio is the squared fraction of s.e. of fixed effects and s.e. of regression.

Table 3 **Number of countries with individual trends in exports-as-capacity-to-import/exports (a)**

Levels

Group	*Signif. Neg.*	*Signif.pos.*	*Insign.*	*Total*	*% sign.neg*	*coeff. lag.dep.(b)*
High income OECD	11	8	8	27	0.41	0.84
High Income Non-OECD	1	2	16	19	0.05	-0.06
Upper Middle income	4	8	27	39	0.10	0.81
Lower Middle Income	11	7	26	44	0.25	0.76
Lower income	15	3	22	40	0.38	0.73

(a) Least squares with country-specific fixed effects and common coefficient of the lagged dependent variable.

(b) Period SUR PCSE; p-val. is 0.0000 in all cases. For high income countries four lags are significant; for all other samples only one lag.

First differences

Group	*Signif. Neg.*	*Signif.pos.*	*Insign.*	*Total*	*% sign.neg*	*coeff. lag.dep.(b)*
High income OECD	10	9	8	27	0.37	-0.05
High Income Non-OECD	0	4	15	19	0.00	-0.27
Upper Middle income	0	1	38	39	0.00	-0.12
Lower Middle Income	2	3	37	43	0.05	-0.42
Lower income	7	3	29	39	0.18	-0.77

(a) Least squares with country-specific fixed effects and common coefficient of the lagged dependent variable.

(b) See Table 1, panel (c) for details of estimation.

Table 4 **Trends in net-barter terms of trade (a)**

Income group	High OECD	High Non-OECD	Upper Middle	Lower Middle	Low
Constant	0.677	0.129	0.551	0.688	0.789
(p-value)	0.000	0.366	0.000	0.000	0.000
coeff.lag.dep.(-1)	1.064	0.981	0.870	0.854	0.908
(p-value)	0.000	0.000	0.000	0.000	0.000
coeff.lag.dep.(-2)	-0.322	-	-	-	-0.062
(p-value)	0.000	-	-	-	0.110
coeff.lag.dep.(-3)	0.192	-	-	-	-
(p-value)	0.001	-	-	-	-
coeff.lag.dep.(-4)	-0.078	-	-	-	-
(p-value)	0.037	-	-	-	-
Coeff. Trend	0.000	0.000	0.001	-0.001	-0.002
(p-value)	0.253	0.535	0.001	0.068	0.003
long-term trend	-0.002	-0.015	0.010	-0.004	-0.012
Adjusted R-squared	0.790	0.887	0.825	0.764	0.883
Durbin-Watson stat (c)	1.972	1.736	1.756	2.002	2.055
prob. fixed eff. redundant	0.874	0.000	0.824	0.226	0.019
Periods	1984-2008	1981-2007	1981-2007	1981-2007	1982-2007
Method: Panel Est.GLS	no ind. effects	fixed eff.	no ind. effects	no effects (b)	fixed eff.(d)
Panel corrected s.e.	Period SUR	Period SUR	Period SUR	Period SUR	Period SUR
Countries	22	13	27	34	32
Observations	491	167	611	734	693

(a) Dependent Variable: LOG(NBT)

(b) Signs and significance also hold with fixed effects.

(c) Durbin-Watson statistic. Although it is not the adequate statistic for rigorous tests under endogeneity, its size indicates that there can be no serious serial correlation bias. See Epple and McCallum 2006.

(d) When using first differences and time fixed effects, lagged dependent variables are insignificant and the growth rate is a negative constant of -0.009875.

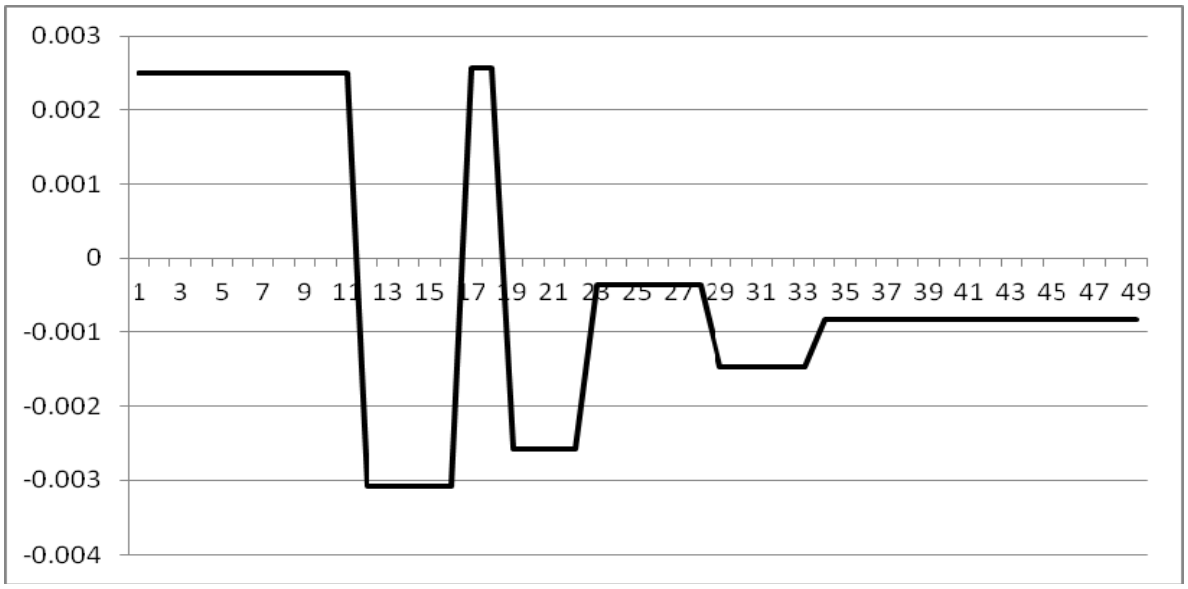

Fig. 1: Time-dependent coefficients of trends for low-income countries (EGLS fixed effects cross), 1960-2008.

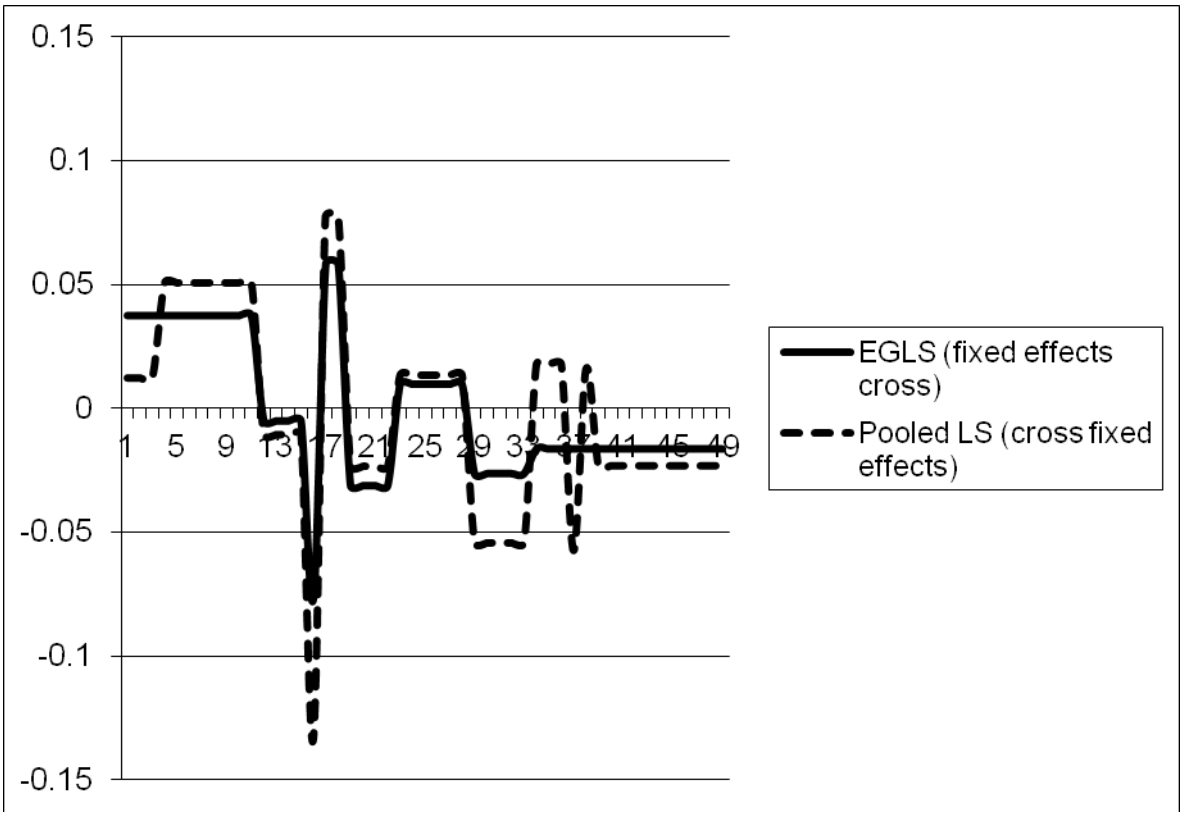

Fig. 2. Intercept dummies replacing a significant downward trend in terms of trade of low-income countries, 1960-2008.

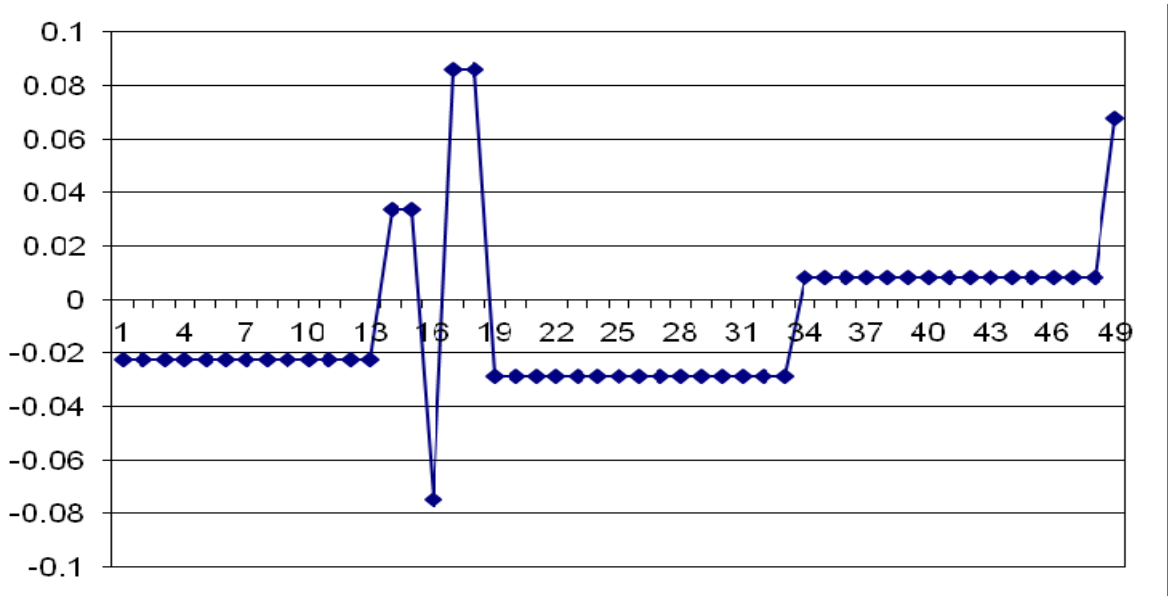

Fig. 3. Cumulated values of constant and intercept dummies for first-difference estimates for lower-middle income countries, 1960-2008.

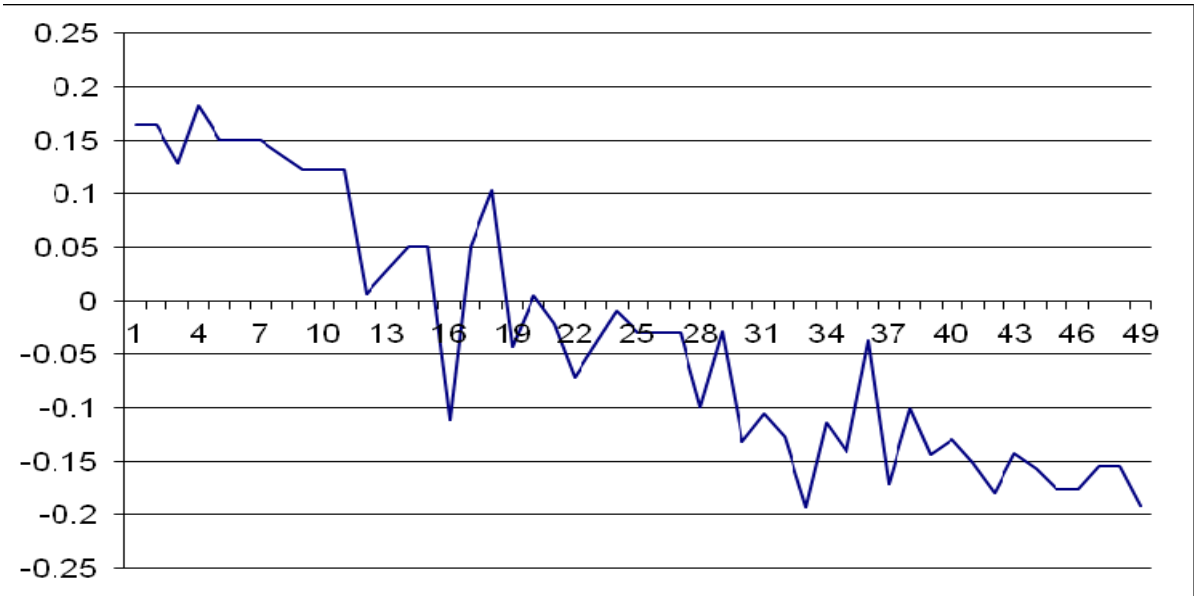

Fig. 4. Period-specific intercepts from time dummies in the presence of country-specific time trends show falling terms of trade, 1960-2008.

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