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Commodity Prices and Growth: An empirical investigation

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Commodity Prices and Growth: An Empirical Investigation*

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Abstract

Whereas empirical evidence on the effect of higher commodity prices on the *long-run* growth of commodity exporters is ambiguous, time series analyses using vector autoregressive (VAR) models have found that commodity booms raise income in the *short run*. In this paper we adopt panel error correction methodology to analyze global data for 1963 to 2008 to disentangle the short and long run effects of international commodity prices on output per capita. Our results show that commodity booms have unconditional positive short-term effects on output, but non-agricultural booms in countries with poor governance have adverse long-term effects which dominate the short-run gains. Our findings have important implications for non-agricultural commodity exporters with poor governance, especially in light of the recent wave of resource discoveries in low-income countries.

Keywords: commodity prices; natural resource curse; growth

JEL classification: O13, O47, Q33

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1. Introduction

Empirical evidence on the long-run effect of natural resources on economic growth is ambiguous. Some studies find that resources are a "blessing" (Alexeev and Conrad, 2009; Brunnschweiler and Bulte, 2008; Lederman and Maloney, 2007; Sala-i-Martin et al., 2004), while others find that they are a "curse" (Sachs and Warner, 1995 and 2001; Gylfason et al., 1999; Sala-i-Martin and Subramanian, 2003). The theoretical literature offers several explanations for this mixed evidence. First, it may be due to adverse effects of natural resources that occur in some countries but not in others. Mehlum et al. (2006), for example, argue that resource rents invite non-productive lobbying and rent seeking in countries with weak "grabber-friendly" institutions but not in countries with strong "producer-friendly" institutions. An alternative explanation is put forward by Robinson et al. (2006), who argue that commodity booms lead to inefficient redistribution by governments in return for political support but that this occurs only in countries where government accountability is lacking. In addition to these political economy channels, other adverse effects of natural resources may also occur in some countries but not in others. Examples include Dutch disease (Corden and Neary, 1982; Torvik, 2001), a deterioration in governance (Baland and Francois, 2000; Tornell and Lane, 1999; Torvik, 2002), conflict (Collier and Hoeffler, 2004), excessive borrowing (Mansoorian, 1991; Manzano and Rigobon, 2007), volatility (Sala-i-Martin and Subramanian, 2003), and lower levels of education (Gylfason, 2001).

But the mixed evidence on the long-run effect of resources on growth could also be due to methodological problems that apply to much of the resource literature. Manzano and Rigobon (2007), for example, show that the resource curse effect identified in the cross-sectional growth regressions of Sachs and Warner (1995) disappears when using a panel and employing a fixed effects estimator to control for unobserved country characteristics.[4] But even when using fixed effects, many of the proxies used for natural resource dependence or abundance are likely to suffer from endogeneity, which makes it difficult to interpret the estimated effects as causal. Brunnschweiler and Bulte (2008), for example, point out that the Sachs and Warner proxy for resource dependence (the share of resources in GNP) is endogenous. They show that, when instrumenting for resource dependence, its negative effect on growth disappears whereas subsoil resource wealth (abundance) positively affects growth.

Whereas the resource literature predicts an ambiguous effect of commodity booms on *long-run* growth, empirical studies by Deaton and Miller (1995) for Africa and Raddatz (2007) for

[4] Lederman and Maloney (2007) find that the Sachs and Warner results also disappear when using the unmanipulated data in cross section, or when employing a system GMM estimator.

low-income countries use vector autoregressive (VAR) models and find that higher commodity prices significantly raise income in the *short run*. Since the purpose of these studies is not to estimate long-run effects of commodity booms, an unexplored possibility is that these positive short-run effects are followed by others, beyond the horizon of the VAR models, whose sign is conditional upon country characteristics and which potentially more than offset initial benefits.

In this paper we adopt panel error correction methodology to analyze global data for 1963 to 2008 to disentangle the short and long run effects of international commodity prices on output per capita. This approach has several advantages compared to earlier studies. First, unlike many of the resource proxies used in previous literature, international commodity prices are typically unaffected by the behaviour of individual countries and so can be viewed as a relatively exogenous source of variation in a country's resource revenues (Deaton and Miller, 1995).[5] Secondly, the error correction model allows us to jointly estimate short-run *growth* effects of booms and long-run *level* effects within one regression framework, which to our knowledge has not previously been done. It avoids that the long-run estimates are contaminated by short-run effects, a problem that is typically ignored in the resource literature. Thirdly, the size of our panel dataset represents a substantial increase in the number of observations used in estimation when compared to the cross-sectional datasets that are used in much of the resource literature, but also when compared to earlier studies that employ panel data, such as Lederman and Maloney (2007) and Manzano and Rigobon (2007). Also, the length of the panel (46 years) is long compared to earlier panel data studies, which is helpful when estimating long-run relationships as a large time span of the data gives more observations on long-run fluctuations (Hakkio and Rush, 1991). Our estimations include fixed effects and regional time dummies to control for unobserved country characteristics and common regional shocks and we allow the effects of commodity prices to vary across different types of commodities. We also address potential sources of endogeneity that have sometimes been neglected in previous literature.

Our results show that commodity booms have positive short-term effects on output, but conditional adverse long-term effects. The adverse long-term effects are confined to "high-rent", non-agricultural commodities. Within this group, we find that the adverse long-term effects are avoided by countries with sufficiently good governance.

[5] We relax this assumption of exogeneity when we address concerns over endogeneity in Section 5.

The conditional adverse long-term effects we identify in our estimations are consistent with the above mentioned theories of Mehlum et al. (2006) and Robinson et al. (2006). They also, we believe, lend more credibility to cross-section results in three previous studies that suffer from the same methodological problems as much of the resource literature (e.g., the use of endogenous resource proxies, the failure to control for unobserved country characteristics and the relatively short length of the sample period). The first of these three studies, by Boschini et al. (2007), shows that in countries where resources are highly appropriable, as determined by both the type of resources and institutional quality, resources lower growth, while in countries with less appropriable resources, they promote growth. The other two studies, by Mehlum et al. (2006) and Iimi (2007), show that natural resources raise growth in countries with good institutions but lower growth in countries with bad institutions.[6]

Our finding that positive short-term effects of booms on output co-exist with conditional adverse long-term effects has important implications for non-agricultural commodity exporters with weak institutions, many of which are located in Sub-Saharan Africa. Global commodity prices remain markedly higher than those before the post-2000 boom and are, if past behaviour is repeated, likely to have strongly adverse long term effects, so that the recent acceleration in growth rates of Africa's commodity exporting economies is particularly misleading. However, if our tentative diagnosis of the root cause of the adverse effects as being due to errors in governance is correct, then this prognosis could be avoided by improvements in the quality of governance.

The rest of this paper is structured as follows. Section 2 describes the empirical analysis. Section 3 reports the estimation results and simulates the short and long run effects of higher commodity export prices on output. Section 4 investigates whether the adverse long-term effect of booms occurs conditional on governance. Section 5 addresses endogeneity concerns. Section 6 concludes.

2. The Empirical Analysis

In this section we describe our econometric model and the variables used in estimation. Data description and sources can be found in Appendix I. The short-run and long-run effects of commodity export prices on GDP per capita are analyzed within the framework of a neoclassical growth model. In this framework, long-run steady state output growth is driven

[6] Collier and Hoeffler (2009) show that the effect of natural resources on growth depends on different features of democracy. Whereas electoral competitition worsens the effect, checks and balances improve it. Others have argued that the effect of natural resources depends on inequality (Dunning, 2008) or human capital (Bravo-Ortega and De Gregorio, 2007).

by exogenous technological progress, while the growth rate during the transition to the steady state is a function of the determinants of the steady state *level* of output and the initial level of output. The predictions of the neoclassical growth model have been studied empirically by Mankiw et al. (1992), Barro and Sala-i-Martin (1995) and Caselli et al. (1996), amongst others. As explained by Bond et al. (forthcoming), these studies relate growth to investment and other explanatory variables, while conditioning on the initial level of output. In a panel data setting, this suggests a specification of the form

for and , where denotes the logarithm of real GDP per capita in country in year , is the growth rate of real GDP per capita between and , is an vector of variables that are expected to affect the long-run steady state level of GDP per capita, is a country-specific fixed effect, and is a time trend.

Equation (1) allows the researcher to study the potential determinants of the steady state level of output, as well as the hypothesis of conditional convergence, i.e. the idea that countries converge to parallel equilibrium growth paths. However, it does not allow the growth rate during the transition to the steady state to be subject to short-run business cycle fluctuations driven by shocks to the economic environment, as for example studied by real business cycle macroeconomic theory.[7] To account for such fluctuations, we augment equation (1) by contemporaneous and lagged changes in . We also add a lagged dependent variable to account for persistence in growth rates.[8] This results in the following core estimating equation of our empirical analysis:

Equation (2) above can be rewritten as an error correction model:

[7] Mendoza (1995) and Kose and Riezman (2001) use calibrated general-equilibrium small-open-economy models based on real business cycle theory to study the impact of terms-of-trade shocks on output fluctuations. Both studies find that terms-of-trade shocks account for around 50% of actual GDP variability.

[8] The fixed effects (within groups) estimator is biased in "small , large " panels with explanatory variables that are not strictly exogenous, such as a lagged dependent variable (Nickell, 1981). However, this bias becomes negligible as grows large. Bond (2006), based on calculations of this inconsistency, and Monte Carlo experiments, concludes that the bias poses a huge problem with , remains non-negligible for or , and can quite comfortably be ignored when or . The average number of time series observations in the core specifications of our analysis ranges from 29 to 38, suggesting that the bias is small. Moreover, additional estimations (not reported) showed that our results on the short-run and long-run effects of commodity export prices on output are robust and, if anything, become stronger when making the bias arguably negligible by excluding all countries with fewer than 30 time series observations.

where , , , for , , and

. In equation (3) above, output “error-corrects”, i.e. responds to deviations from long-run equilibrium (captured by the term between brackets) in a way that gradually brings the economy back to its steady state. This error-correction process implies that the coefficient in equation (3), which equals the coefficient in equation (2), should be negative, while the size of this coefficient captures the speed of (conditional) convergence. The long-run steady state equilibrium is attained when the term between brackets in equation (3) equals zero so that . If we assume that in long-run equilibrium the determinants of output take the constant value , the steady state growth rate is given by so that , and the long-run equilibrium condition can be written as

(4)

Equations (3) and (4) above show how the coefficients from our estimating equation (2) can be mapped into long-run effects on the steady-state *level* of output and short-run effects on the *growth rate* of output. In particular, the long-run effects of the variables in the vector in equation (2) are captured by the vector in equations (3) and (4) and, given that —, can be computed from the coefficients in equation (2). By contrast, the short-run effects directly follow from the coefficients in equation (2): for the speed of convergence, for the short-run effect of growth in the previous year, and for the short-run effects of changes in the -variables. In addition, equations (3) and (4) also emphasize the importance of the country-specific fixed effect and the time trend in equation (2). The fixed effect, , controls for any country-specific time-invariant unobservables that affect the steady state level of output, as can be seen from in equation (4). It also controls for country-specific time-invariant unobservables that affect growth during the transition to the steady state, as can be seen from in equation (3). The time trend allows the steady state growth rate in equations (3) and (4) to be different from zero, which is important given that average output typically increases over time (Durlauf et al., 2005). It also allows for a short-term trend in growth rates during the transition to the steady state, as can be seen from in equation (3).

Four comments are in order. First, the inclusion of a simple linear time trend in equation (2) above restricts the steady state growth rate to be the same for all countries in all periods. To allow for a more heterogeneous steady state growth path, we will also experiment with the inclusion of an vector of regional time dummies instead of a time trend, where represents the number of regions. The regional time dummies capture year-specific fixed effects for each of the following geographical regions: (i) Central and Eastern Europe and Central Asia, (ii) East Asia and Pacific and Oceania, (iii) Latin America and Caribbean, (iv) North Africa and Middle East, (v) South Asia, (vi) Sub-Saharan Africa, and (vii) Western Europe and North-America. The inclusion of these regional time dummies in equation (2) allows the steady state growth rate to differ across the seven regions and, within each region, across years. In addition, the dummy variables also control for common regional macroeconomic shocks that may affect short-term growth.

Secondly, although error correction models inherently use the concept of *long run*, its meaning can vary substantially across applications. While for some economic questions, the long run may correspond to several months, for others it may be a matter of years or even decades (Hakkio and Rush, 1991). In our analysis, we think of the long run as the point in time when the ramifications of political and economic policy decisions that were made in response to commodity booms have occurred. One example is the occurrence of Dutch disease, that is, the loss of positive externalities such as learning-by-doing due to a decline of the non-resource exports sector. A second example is the output loss from lobbying and rent-seeking that occupy otherwise productive entrepreneurs. The remaining candidate explanations for the resource curse already described in the introduction represent additional examples. What many of these ramifications have in common is that they only gradually come about so that their cumulative effects on output only slowly materialize. A classic example is the case of the Netherlands (from which the term Dutch disease originates). In the late 1950s, the Dutch discovered large natural gas fields which quickly led to a substantial increase in government revenues. As a result, the Dutch guilder appreciated in real terms which led to a sharp decline in the competitiveness of the non-resource exports sector. Subsequently, during the 1960s and 1970s, employment in manufacturing steadily declined (Kremers, 1986). While this is just one example, it suggests that, when analyzing the long-run effects of commodity booms on output, the long run may correspond to a period of one or two decades. As we show in Section 3 below, the length of this time horizon is to some extent comparable with the long run implied by our main error-correction estimates, which suggest a speed of adjustment of 5 to 6 percent per year and hence a half life of shocks of around 11 to 13 years. Hakkio and Rush (1991) argue that, when estimating long-run relationships, the total sample period (span) needs to be sufficiently long relative to the length of the long run for the data. In our analysis, the sample period is 46 years which is quite long for a cross-country panel dataset. Nevertheless, it would be useful to repeat our analysis in the future when more years of data become available.[9]

Thirdly, unless stated otherwise, we compute robust standard errors clustered by year to account for heteroskedasticity and cross-country correlation in the error terms. The latter is likely to be important, as many of the countries in our sample are subject to common macroeconomic shocks.

Fourthly, most studies that estimate panel growth regressions use five-year or ten-year averages to eliminate cyclical fluctuations that could contaminate estimates of longer-term effects (Durlauf et al., 2005). Since our main goal is to analyze both the short-run and long-run effects of commodity export prices, we are interested in econometrically modelling not just long-run growth but also short-run output deviations. Rather than using averaged or Hodrick-Prescott filtered data, we therefore prefer to use original annual data and control for a range of shocks that cause short-run deviations from potential output. As already discussed, we also include separate year-specific fixed effects for seven geographical regions to control for common regional macroeconomic shocks. Controlling for these various shocks is likely to account for an important part of the cyclical variation in growth rates, thereby reducing the likelihood of contamination of the long-run results, while still allowing us to estimate the effect of commodity export prices on *short-run* growth.

Having discussed our econometric model, we next describe the right-hand side variables included in the vector , starting with our indicators of commodity export prices.

2.1 Constructing indicators of commodity export prices

To estimate the long-run effect of commodity export prices on GDP per capita, the vector in equation (2) includes a commodity export price index constructed using the methodology of Deaton and Miller (1995) and Dehn (2000). We collected quarterly world commodity price indices and 1990 commodity export and import values for as many commodities as data availability allowed. Table 1a lists the 50 commodities in our sample. We averaged the quarterly price indices across all observations in a calendar year to obtain an annual price index for each commodity *c*, . For each country, we calculated the total 1990 net export value (exports minus imports) of all commodities for which the country is a net exporter. We then constructed weights by dividing a country's individual 1990 net export value for each commodity by this total:

[9] As shown by Hakkio and Rush (1991), while increasing the *span* of the data is helpful when identifying long-run relationships, switching to more frequently sampled data (*periodicity*), such as for example quarterly data in our case, is not very helpful.

In the above equation, *x* and *m* stand for exports and imports of commodity *c* by country *i* in 1990, respectively. The 1990 weights, , are held fixed[10] over time and applied to the annual price indices of the same commodities to form a geometrically weighted index. We deflated this index by the annual average of the quarterly export unit value, , rescaled it so that 1980 = 100, and calculated the natural logarithm. This resulted in an annual country-specific logged index of commodity export prices, , which is formally defined as:

To allow the effect of commodity export prices to be larger for countries with larger exports, we weight by the 1990 share of net commodity exports in GDP, , which is defined as:

We use the weighted index, , in our estimations. To investigate whether the effects of commodity prices vary across different types of commodities, we experiment with sub-indices for non-agricultural and agricultural commodities. These sub-indices were constructed in the same way as the composite index and are represented by and , respectively, where the superscripts and stand for non-agricultural and agricultural commodities.

Having described how the commodity export price index is constructed, we next discuss the interpretation of its long-run coefficient in the estimating equation. Partitioning and as and , respectively, equation (4) can be rewritten as:

(8)

Equation (8) implies that, everything else equal, a increase in leads to a change in the log of real GDP per capita, which is approximately equal to a *percent* change in real GDP per capita. In this paper, we interpret as a long-run elasticity: if country *i*'s commodity export prices *simultaneously* increase by one percent, its level of GDP per capita will change by percent, holding other things constant. To see that this interpretation is consistent with the construction of the commodity export price index, it is

[10] Deaton and Miller (1995) argue that keeping the weights constant over time is important because it ensures that endogenous supply responses to price changes are excluded from the analysis. It also means that we lose some changes in the composition of primary exports but, as recognized by Deaton and Miller (1995), this loss is inevitable if we are to exclude endogenous quantity changes. Moreover, the loss is likely to be limited as the pairwise correlations between the 1990 weights and the same weights for 1970, 1980, and 2000, are 0.74, 0.87, and 0.84, respectively, indicating that the weights of individual resources in a country's primary exports are relatively persistent over time.

important to recall that the weights ( ) used to construct the geometrically weighted index are held fixed over time. More specifically, let us consider a case with two commodities. The index of commodity export prices, , can then be written as:

Since the weights of the commodities, and , are fixed over time, equation (9) implies that a simultaneous one *percent* increase in the real prices of commodities 1 and 2 will lead to a +0.01 increase in . Since we know from equation (8) that, everything else equal, a +0.01 increase in leads to a *percent* change in real GDP per capita, this shows that our interpretation of as a long-run elasticity is consistent with the construction of the commodity export price index. Indeed, if country *i*'s commodity export prices *simultaneously* increase by one percent, its level of GDP per capita will change by percent, everything else equal. The role played by is noteworthy. It suggests that the elasticity is higher in countries where commodity exports represent a larger share of GDP.

Having discussed the interpretation of the long-run coefficient, we now take a first casual look at the evolution of commodity export prices and economic growth rates over time. Figure 1a shows average economic growth ( ) and the average level of the commodity export price index ( ) over the period 1961-2008 for sixteen large *non-agricultural* commodity exporters in our sample. The time series for the commodity index contain some well-known features. While non-agricultural commodity prices were least volatile in the 1960s, 1973-1974 and 1978-1979 saw large spikes due to the energy crises caused by problems in the Middle East. Prices fell again during the 1980s and 1990s although permanently staying above the levels of the 1960s, and then rose to unprecedented levels after 1999. Commodity booms and busts seem to roughly coincide with accelerations and decelerations in economic growth, respectively, consistent with a positive short-run effect of higher commodity prices on growth. However, while growth rates of non-agricultural commodity exporters went up around the time of the commodity booms of the 1970s, they fell substantially in the years thereafter and on average remained low throughout the 1980s and 1990s. In fact, while these economies on average grew by 2.6 percent per year during the period 1961-1972, they only grew by an average 0.5 percent per year during the period 1982-1999, after which growth picked up during the global commodity boom. Figure 1b shows average growth and the average level of the commodity export price index for twenty-one large *agricultural* commodity exporters in our sample. Again, the time series for the index reflect some well-known events. Agricultural commodity prices were relatively stable during the 1960s, rose sharply in the mid-1970s, then fell dramatically until the mid-1990s, and sharply increased again after 2003. Commodity booms and busts at times seem to roughly coincide with accelerations and decelerations in economic growth, respectively, although this is clearly not always the case. While global food prices tumbled between 1984 and 1993, growth rates of agricultural commodity exporters on average went up.

2.2 Control variables

The vector also includes a range of long-run control variables. Earlier studies that estimate an empirical version of the neoclassical growth model, as we do in this paper, typically include variables that proxy for those suggested by the (augmented) Solow model, as explained in Durlauf et al. (2005) and Bond et al. (forthcoming). All our specifications therefore include initial output per capita (already separately included in equation (2)), the investment share of GDP, population growth and a measure of human capital (the average number of years of secondary schooling for the population aged 15 and above).

Since there is no general consensus on which growth determinants should be included in a growth model, we follow the approach described in Durlauf et al. (2005) and applied for example in Levine and Renelt (1992). In particular, while we always include the Solow regressors described above, we experiment with the inclusion of a wide range of additional long-run control variables based on the ranking of 67 proxies in Sala-i-Martin et. al. (2004). A list of these additional controls can be found in Appendix I.[11]

To estimate the *short-run* effect of commodity export prices on GDP growth, we also include contemporaneous and lagged *changes* in the commodity export price index. In addition, we experiment with the inclusion of several short-run controls. A list of these short-run controls can be found in Appendix I.

Our dataset consists of all countries and years for which data are available, and covers around 120 countries between 1963 and 2008. Table 1b reports summary statistics.

3. Estimating the short and long run effects of commodity prices

[11] We augmented this list with an oil import price index to control for the effect of oil prices on oil-importing countries.

Table 2 reports the results of estimating equation (2).[12] The first column reports the results for the most parsimonious specification with the Solow regressors but without the additional short-run and long-run controls. Table 2, column (2), shows the results when adding the robustly significant long-run controls from Sala-i-Martin et al. (2004), while column (3) shows the results when adding the robustly significant short-run controls as well. Table 2, columns (4) to (6), report the results when replacing the trend in the specifications of columns (1) to (3) by the regional time dummies.[13] Clearly, the selection of control variables is an important issue. As we show in Section 5, our results are robust to the inclusion of each of the other additional long-run and short-run controls listed in Appendix I.

The estimating equation (2) is only appropriate if there is a long-run levels relationship between GDP per capita and the long-run right-hand side variables. In addition, estimating equation (2) in a single-equation framework without additional equations for the long-run right-hand side variables is only appropriate if these variables are weakly exogenous. Tests for weak exogeneity and the existence of a long-run levels relationship, discussed in Appendix II, suggested that these assumptions are justified.

In all of the specifications of Table 2, columns (1) to (6), the long-run coefficient of the commodity export price index is negative and statistically significant at 1 or 5 percent, consistent with a long-run resource curse effect. Higher commodity export prices significantly reduce the long-run level of real GDP in commodity exporting countries. The size of the coefficients implies a substantial effect. Figure 2a, based on the coefficient in Table 2, column (5), shows this effect as a function of a country's dependence upon commodity exports. For example, in 1990 in both Zambia and Nigeria commodity exports constituted 35 percent of GDP. The results in Figure 2a therefore predict a long-run elasticity of -0.52.[14] In other words, a 10 percent increase in the prices of their commodity exports leads to a 5.2 percent lower long-run level of GDP per capita. We should note that a reduction in constant-price GDP is not the same as a reduction in real income. The higher

[12] As explained in the previous section, the estimated long-run coefficients correspond to –, while the short-run coefficients correspond to , and .

[13] We selected the long-run controls from Sala-i-Martin et al. (2004) by separately adding each of the controls listed in Appendix I to the specifications of columns (1) and (4). Five controls were significant at 5 percent in both specifications. When jointly adding these five controls, three remained significant (trade to GDP, population ages 0-14 and inflation (log)) and we included these in the specifications of columns (2), (3), (5) and (6). The short-run controls were selected in a similar way by separately adding each of the short-run controls listed in Appendix I to the specifications in columns (2) and (5). Three variables were significant at 5 percent in both specifications and remained so when jointly added (coup, civil war and natural disaster). We considered possible long-run effects of these variables using two approaches. First, we experimented with indicators of the cumulative number of coups, war years and natural disasters (see Appendix I) but found no significant effects. Second, we re-estimated the specifications of Table 2, columns (3) and (6), with 1, 2 or 3 lags of the short-run indicators but found no evidence that the adverse effects of coups, civil wars and natural disasters last beyond the year(s) of occurrence. These results are consistent with the findings of Murdoch and Sandler (2002) and Sala-i-Martin et al. (2004) and suggest that coups, civil wars and natural disasters reduce short-term growth but not the long-term level of GDP per capita.

[14] Recall that the commodity export price index ( ) is weighted by the share of net commodity exports in GDP ( ). So for Zambia and Nigeria, the long-run elasticity equals the long-run coefficient, -1.475, multiplied by the share of net commodity exports in GDP, 0.35.

export price directly raises real income for a given level of output and this qualitatively offsets the decline in output. The magnitude of this benefit from the terms of trade follows directly from the change in the export price and the share of exports in GDP. Thus, in the examples above, the terms of trade gain directly raises income by 3.5 percent for given output. Even so, this is less than the decline in output of 5.2 percent, so that the resource curse ends up reducing both output and income relative to counterfactual.[15]

Having discussed the long-run effects of commodity export prices, we now turn to the other variables in the specifications of Table 2, columns (1) to (6). The coefficients of the Solow regressors (except for population growth) always have the expected signs but in most cases are not significant. Only the coefficient of lagged GDP per capita is robustly significant at 1 percent. The size of this coefficient, which captures the speed of adjustment to equilibrium, or conditional convergence, indicates that output returns to the long-run equilibrium at a speed of around 3 to 6 percent per year. Hence, the adverse long-run effect of higher commodity export prices on output is not instantaneous but comes about gradually. After the commodity price increase, output "corrects" by 3 to 6 percent of the remaining deviation from its new, lower, long-run level each year, implying a prolonged phase of slower growth. This adjustment process continues until output reaches the new long-run equilibrium growth path and the resource curse effect is complete. The coefficients of the additional long-run controls (trade to GDP, population ages 0-14 and inflation) from Sala-i-Martin et al. (2004) have the expected signs and are always significant at 1 percent.

Having discussed the long-run effects and the adjustment phase, we now turn to the short-run effects. The contemporaneous as well as the first and second lag of the change in the commodity export price index enter positive. This effect is always significant at 1 percent for the first lag, while in half of the specifications significant at 1 or 5 percent for the contemporaneous change. These results indicate that an increase in the growth rate of commodity export prices initially has a positive effect on GDP growth. Thus, the short-run dynamics of a commodity boom are quite contrary to the long-run effects. During the first few years after a boom, the positive short-run effect coexists with the gradual adjustment to the adverse long-run effect. To illustrate the *net* effect, Figure 2b (based on the results in Table 2, column 5) shows the impulse response functions of an increase in the growth rate of

[15] We re-estimated the specifications of Table 2, columns (1) and (4), *without* investments, population growth and education (while keeping the sample size constant). The results were very similar, suggesting that the long-run effect of commodity export prices does not work *through* these three long-run controls. We also re-estimated the specifications in Table 2, columns (2) and (5), *without* trade, population ages 0-14 and inflation (while keeping the sample constant). The estimated long-run effect of commodity prices was of similar magnitude although somewhat larger, which might imply that a small fraction of the long-run effect of commodity prices is explained by these three additional long-run controls. If so, the estimated long-run effect of commodity export prices in the specifications that include these additional controls should probably be viewed as a somewhat conservative estimate of the true effect.

commodity export prices for different levels of commodity exports to GDP. In the year of the price increase and the two subsequent years, the short-run positive effect dominates the adjustment to the long run and growth goes up. The effect of a 10 percentage points increase in prices in period t cumulates to 0.26 percentage points of GDP growth after year t+2 in countries with commodity exports that represent 10 percent of their GDP. This growth gain amounts to 0.52, 0.78, and 1.05 percentage point for countries with commodity exports to GDP shares of 20, 30 and 40 percent, respectively. The positive net short-run effect of commodity export prices is consistent with the findings of Deaton and Miller (1995) and Raddatz (2007).[16] Further, the short run effects on output are reinforced by the direct gain in income through the improvement in the terms of trade, so that real incomes rise strongly. However, our results also indicate that this short-run gain is temporary. Over time, the growth acceleration is reversed as the short-run effect of the boom dies out and output gradually adjusts to its new, lower, long-run level.

Table 2, columns (1) to (6), also report the coefficients of the other short-run GDP determinants. The coefficient of the lagged dependent variable is positive and robustly significant at 1 percent. We experimented with additional lags but found that these are unimportant. Next, coups and civil wars have highly significant adverse effects on growth. A coup appears to cut growth by around 3.1 percentage points in the same year, while the negative impact of civil war is estimated to be 1.0 percentage point for each year of the war, consistent with Collier (1999). We investigated whether this varies during the course of the war but could find no significant effect. Finally, natural disasters significantly reduce growth by around 0.5 percentage points.

The specifications in Table 2, columns (1) to (6), all include country fixed effects to control for unobserved heterogeneity. To assess the importance of this heterogeneity, Table 2, column (7), for comparison reports the results when excluding the fixed effects from the specification in column (5). The coefficient of the lagged dependent variable is somewhat higher than in the fixed effects specification of column (5). This is consistent with the observation of Bond (2002, 2006) that, in the presence of unobserved individual-specific time-invariant effects, the OLS estimator of the coefficient for the lagged dependent variable is biased upwards due to the positive correlation of this variable with the individual effects. In contrast to the fixed effects bias, the OLS bias does not disappear as the number of time periods increases so that OLS remains inconsistent for panels with large , such as ours. The

[16] Raddatz (2007) documents that a 14 percent increase in commodity export prices results in a 0.9 percent increase in GDP after four years. Both Raddatz (2007) and Deaton and Miller (1995) do not distinguish between short-run and long-run effects of commodity prices.

coefficient of the lagged level of GDP per capita is also substantially higher than in column (5) and also likely to be biased upwards due to a positive correlation with the individual effects. These results indicate the presence of unobserved heterogeneity and, given the large time dimension of our panel, support the choice of a fixed effects estimator over an OLS estimator without fixed effects. Nevertheless, it is reassuring that the short-run and long-run coefficients of the commodity export price index in Table 2, column (7), are similar to the coefficients in column (5), both in terms of size and significance.[17]

Having found that higher commodity export prices significantly reduce the long-run level of real GDP in commodity exporting countries, we next investigate whether this adverse long-run effect is common to all the commodities in our index. We decompose the composite commodity export price index into two sub-indices: one for non-agricultural commodities only and one for agricultural commodities only. Table 3 shows the results when we replace the composite index in Table 2, columns (1) to (6), by the two sub-indices. For non-agricultural commodities we again find strong evidence of an adverse long-run effect. In all six specifications, the coefficient is negative and statistically significant at 1 or 5 percent.[18] For Zambia and Nigeria, whose commodity exports are overwhelmingly non-agricultural, the results in Table 3, column (5), predict a long-run elasticity of -0.59. In other words, a 10 percent increase in the price of oil leads to a 5.9 percent lower long-run level of GDP. By contrast, the coefficient for agricultural commodity export prices is always positive and insignificant. This suggests that higher agricultural export prices are not a curse analogous to non-agricultural commodities.[19]

4. The adverse long-term effect of booms conditional on governance

The results in the previous section point indirectly at governance as being important in explaining the resource curse. This is because of the sharp distinction we have found between the agricultural and non-agricultural commodities. This distinction closely corresponds to whether or not the activity generates rents. Agricultural commodities can be produced in many different locations and so competitive entry will drive profits to normal levels. The

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[17] To assess the importance of *parameter* heterogeneity, we split the sample of Table 2, column (5), into pre- and post-1985 subsamples, this being the year in the middle of the sample period. We then re-estimated the specification for both subsamples separately and for each coefficient performed a Wald test of equality across the two subsamples. We found that eight out of the twelve coefficients, including the long-run coefficient of the commodity export price index, do not significantly differ across the two subsamples. The coefficients of the contemporaneous and second lag of the change in the commodity export price index are significantly higher in the pre-1985 period but the coefficient of the first lag remains positive and significant at 1 percent in both subsamples.

[18] Given the economic importance of oil, we experimented with a further decomposition of non-agricultural commodities into oil and other non-agricultural commodities. An -test on the coefficients of these two sub-indices did not reject the null hypothesis of equal coefficients. This suggests that we can analyze oil and other non-agricultural commodities as a common aggregate.

[19] We also investigated whether the positive *short-run* effects of higher commodity export prices differ across non-agricultural and agricultural commodities but found no evidence that this is the case.

rents on land used for export crops should therefore be no higher than that used for other crops, once allowance is made for differences in investment, such as the planting of trees. In contrast, the non-agricultural commodities are all extractive, the feasibility of production being dependent upon the presence of the resource in the ground. Hence, the extractive industries all generate rents as a matter of course. Mehlum et al. (2006) and Robinson et al. (2006) argue that rents lead to rent-seeking and inefficient redistribution in countries with weak "grabber-friendly" governance but not in countries with strong "producer-friendly" governance. This suggests that the adverse long-term effect of booms occurs *conditional on* weak governance.

To investigate this possibility, we split the countries in our sample in two groups according to their mean International Country Risk Guide (ICRG) composite risk rating between 1984 and 2009.[20] The ICRG is a commercial rating service whose continued viability has been dependent upon client firms regarding it as having value. There is therefore some reasonable presumption that it has informational content. The first group, which for convenience we will call the "good governance" group, consists of the countries with a mean ICRG score of 75 or higher. This group contains countries like Australia, Canada, and Norway, but also Botswana. The second "bad governance" group consists of the countries with a mean ICRG score below 75 and contains for example Venezuela, Libya and the Republic of the Congo.

We next investigate whether the long-run effect of commodity export prices differs between the good governance and bad governance countries. We begin with the composite index and then focus on the decomposition into agricultural and non-agricultural commodities since it is only the latter where we find evidence of an adverse long-term effect. We introduce governance by constructing an interaction term of the commodity price index with a dummy variable that, for all years in our sample period 1963-2008, takes a value of 1 for good governance countries and 0 for bad governance countries.[21] We then add this interaction to the specifications in Table 2, columns (2) and (5), and Table 3, columns (2) and (5). The results are reported in Table 4.[22] In column (1), the commodity export price index enters negative and is significant at 1 percent, indicating that there is indeed an adverse long-

[20] Since the ICRG is an ordinal variable it is best introduced into the quantitative analysis through a threshold.

[21] We compute the governance dummy for *all* years in our sample period 1963-2008 (not just for the post-1984 period for which the ICRG rating is available). The main reason is that we want to include the commodity booms of the 1970s, which, except for the recent boom, represented the largest increases in global commodity prices since 1960 (as can be seen from Figure 1a). This obviously has drawbacks if a country's mean post-1984 ICRG score is not a good indicator for its quality of governance in earlier years. To assess the importance of this problem, we constructed separate mean ICRG scores for the periods 1984-1993, 1994-2003 and 2004-2009. The correlations between the mean scores ranged from 0.82 to 0.91, suggesting a relatively high persistence over time. This lends some plausibility to the claim that average post-1984 ICRG scores are probably good proxies for governance over the entire sample period 1963-2008. But even if they are not and the governance indicator is thus subject to measurement error, this is unlikely to lead to a substantial bias in our estimates, as we show in Section 5 when we address endogeneity (including measurement error) by employing instrumental variables.

[22] Since we only include countries for which the mean ICRG score is available, the number of observations drops from 3793 to 3371.

term effect of booms in countries with bad governance. The interaction term of the index with the good governance dummy enters positive but at this stage is significant at 10 percent only.

In Table 4, column (2), we again replace the trend in the specification of column (1) by the regional time dummies. The composite commodity export price index again enters negative and is significant at 1 percent, while its interaction with good governance is again positive but is now significant at 5 percent.

In Table 4, columns (3) and (4), we again decompose the composite index into sub-indices for non-agricultural and agricultural commodities. As previously, the direct effect of the non-agricultural export price index is negative and significant at 1 percent, suggesting that badly governed countries suffer from an adverse long-run effect of higher non-agricultural commodity prices. However, the interaction term of the index with the good governance dummy enters positive and is now significant at 5 and 1 percent, respectively. This indicates that the long-run effect of non-agricultural export prices is different for good governance countries. For such countries, the net long-run effect is given by the linear combination of the two coefficients, which is positive (although not significant). This suggests that countries with good governance do not suffer from an adverse long-term effect of booms and may even be able to transform them into sustainable higher output. These findings support the hypothesis that the adverse long-term effect of commodity booms occurs conditional on bad governance. The agricultural index enters positive but is insignificant, while its interaction enters negative and is also insignificant. This indicates that the effects of higher agricultural export prices in countries with good and bad governance are not significantly different. It also supports our earlier finding that higher agricultural export prices do not lead to any adverse long-term effect.

We next re-estimate the specifications in Table 4 using the initial 1985 composite ICRG scores rather than the average scores and find that the results for the composite index and the two sub-indices are robust.[23]

Finally, to further explore the non-linear effect of non-agricultural commodity prices, Table 5 reports the results of separate regressions for countries with bad governance and countries with good governance. Columns (1) and (3) show the results for the subsample of bad governance countries when including a time trend and regional time dummies, respectively. In both cases the coefficient of the non-agricultural index is negative and significant at 1 percent. This is consistent with the earlier finding of an adverse long-term effect of booms in

[23] The first year for which ICRG scores are available is 1984 but the coverage is better for 1985. Given that 1984 and 1985 scores are highly correlated (> 0.98), we use 1985 scores. We again separate the countries into "good governance" (1985 score > 70) and "bad governance" (1985 score ≤ 70). The proportion of good governance countries is equal across the average ICRG and 1985 ICRG samples (22%).

countries with bad governance. Table 5, columns (2) and (4), show the results for the subsample of countries with good governance. The coefficient of the non-agricultural index now enters positive and is insignificant in column (2) and significant at 10 percent in column (4). These results support our earlier finding that the adverse long-term effect of booms is absent in countries with good governance and that, if anything, the long-run effect of higher export prices is positive, as one would expect. Although the coefficients are at most marginally significant and should be viewed with caution, their size suggests a substantial effect. For a country like Norway, which in 1990 had non-agricultural commodity exports that represented 15 percent of its GDP, the results in Table 5, columns (2) and (4), predict a long-run elasticity of between 0.21 and 0.35. In other words, a 10 percent increase in the price of non-agricultural commodities leads to a 2.1 to 3.5 percent higher long-run level of Norway's GDP per capita.[24] These results provide evidence that the adverse long-term effect of commodity booms occurs conditional on bad governance. Countries with sufficiently good governance do not suffer from this adverse effect, and instead may even benefit from higher commodity prices, both in the short run and in the long run.

5. Endogeneity and robustness

A possible concern with the results in the previous sections is that the commodity export price index, , is endogenous, i.e. correlated with the error term in equation (2). Let us first consider the potential endogeneity of . As argued by Deaton and Miller (1995), one of the advantages of using international commodity prices is that they are typically not affected by the actions of individual countries. Also, by keeping the weights of individual commodities constant over time, endogenous supply responses to price changes are excluded from the analysis. Nonetheless, countries that are major exporters of one or more commodities may have an influence on the world price of those commodities, which could lead to biased estimates. To address this concern, we express each country's exports of a given commodity as a share of the total world exports of that commodity and repeat this for all other commodities in our sample. This yields a list of commodity export shares that reflect the importance of individual exporters in the global markets for individual commodities. We found that of the 120 countries in our sample, 24 countries export at least one commodity for which their share in world exports exceeds 20 percent. We investigate whether the inclusion of these major exporters affected our results by re-estimating the specifications in Tables 2, 3

[24] The results in Table 5 are robust to using the initial 1985 composite ICRG scores instead of the average scores, except for a much more significant positive effect of non-agricultural export prices in countries with good governance.

and 4 without these 24 countries, but find no evidence that this is the case. The long-run coefficients of the composite and non-agricultural indices and their interactions with governance are similar to the original coefficients, both in terms of size and significance. The short-run positive effects of commodity prices are robust as well. Hence, our results do not seem to be biased by the possible influence of major exporters on world prices.[25]

We next address the endogeneity of the share of commodity exports in GDP and the good governance dummy, using instrumental variables. We use the estimated 2000 values of sub-soil assets in thousands of current US dollars per capita by the World Bank (2006) as an instrument for the share of non-agricultural exports in GDP, these being the commodities that appear to generate the adverse long-term effect on output.[26] The estimates are based on the net present value of a country's expected benefits over a horizon of 20 years and include 13 commodities, 12 of which are included in our non-agricultural index. Since the share of non-agricultural exports in GDP, , only enters our specifications as part of the non-agricultural price index, , we use as an instrument for , where is the 2000 value of sub-soil assets. For to be a valid instrument, it should be correlated with and not correlated with the error term. While the first requirement is likely to be met, the second requirement is less likely to be fulfilled. Slow-growing countries are likely to have smaller stocks of *discovered* resources due to overexploitation and lower investment in geological exploration. This implies that weighting by the value of sub-soil assets per capita, , may over-weight *fast-growing* countries. Although this could bias the results, the direction of the bias is likely to be opposite to any OLS bias, as the use of non-agricultural exports in GDP may imply over-weighting *slow-growing* countries with underdeveloped non-resource sectors. Comparing the 2SLS and OLS coefficients can therefore bound the potential bias. In addition to the non-agricultural price index, , we also need to instrument for its interaction with good governance, , where represents the dummy for good governance. The best instrument for governance is probably the settler mortality rate used by Acemoglu et al. (2001), but it is only available for 4 out of the 21 good governance countries in our sample. We therefore use three alternative variables, taken from Hall and Jones (1999): the fraction of the population speaking English, the fraction of the population speaking one of the major languages of Western Europe (English, French, German, Portuguese, or Spanish),

[25] We repeated this robustness check using a threshold of 10 percent instead of 20 percent. 38 out of the 120 countries export at least one commodity for which their share in world exports exceeds 10 percent. Again, our findings in Tables 2, 3 and 4 were for the most part robust to the exclusion of these 38 countries. The only result that did not survive was the interaction effect of the composite and non-agricultural indices with good governance. This was due to the fact that only 9 of the 21 good governance countries remained in the sample.

[26] The World Bank estimates of natural capital were earlier used by Brunnschweiler and Bulte (2008), who argue that measures of resource wealth are less prone to endogeneity than measures of resource exports over GDP.

and a country's distance from the equator, measured as the absolute value of latitude in degrees divided by 90 to place it on a 0 to 1 scale. We first run a cross-sectional probit regression of the governance dummy, , on these three variables for the 96 countries in the sample of Table 4 for which we have data. All three variables enter with the expected positive sign, while latitude and the fraction of the population speaking English are significant at 1 percent and 5 percent, respectively. The pseudo R-squared of the probit regression is 0.42. Letting denote the fitted values of the probit regression, extrapolated to all years in the sample of Table 4, we then interact with the instrument for the non-agricultural price index, , discussed above. This yields an additional instrument, , which we use to instrument for the interaction of the non-agricultural index with the good governance dummy, . We next use the constructed instruments to perform two-stage-least-squares estimation. For comparison, Table 6, columns (1) and (4), first report the OLS results. The short and long run effects of non-agricultural commodity prices are consistent with the results in Table 4. Table 6, columns (2) and (5), report the second-stage results of a 2SLS procedure in which the lagged level and differences of the non-agricultural index (, , and), and the lagged interaction of the index with the dummy for good governance ( ) are treated as endogenous explanatory variables. As excluded instrumental variables we use the corresponding lagged level and differences of the instrument for the non-agricultural index (, , and , the lagged instrument for the interaction of the index with good governance ( ) and, in addition, the lagged level and differences of the unweighted index (, , and . As can be seen from the diagnostic tests at the bottom of columns (2) and (5), the Kleibergen-Paap rk Wald *F* statistic is below 10, which indicates a possible weak instrument problem. The Angrist-Pischke first-stage *F* statistics for individual endogenous regressors (not reported) indicated that the problem is due to the weak identification of the *short-run* coefficients. Since this might reflect a problem of multicollinearity, we re-estimate the specifications of Table 6, columns (2) and (5), without the contemporaneous and twice lagged differences of the non-agricultural index.[27] The results are reported in Table 6, columns (3) and (6). The Kleibergen-Paap rk Wald *F* statistic in both cases is now above 10 and exceeds the critical values for a

[27] We use the lagged level and lagged difference of the instrument for the non-agricultural index (and , the lagged instrument for the interaction of the index with good governance ( ) and the lagged level and lagged difference of the unweighted index (and as excluded instruments.

maximal IV relative bias of 5%, 10%, 20% or 30%, indicating the absence of a weak instrument problem (Stock and Yogo, 2005).[28] The long-run coefficient of the non-agricultural export price index is negative and significant at 1 percent in all four 2SLS specifications in Table 6, columns (2), (3), (5) and (6).[29] The size of the coefficients is smaller than the size of the coefficients in columns (1) and (4), possibly indicating an endogeneity bias. However, individual endogeneity tests (not reported) in all four cases did not reject the null hypothesis of consistent OLS estimates for the non-agricultural export price index.[30] Given that any potential biases in the OLS and 2SLS estimates are likely to have opposite signs, the failure to reject exogeneity implies that such biases are probably small. The coefficients of the interaction of the index with the good governance dummy are similar to the coefficients in columns (1) and (4), and are significant at 1 percent. As previously, endogeneity tests did not reject the null of exogeneity. The short-run coefficients of the non-agricultural index enter with positive signs and gain in both size and significance compared to the OLS estimates in columns (1) and (4), while endogeneity tests did again not reject the null of exogeneity. Taken together, these results suggest that biases in the OLS estimates of the short- and long-run effects of non-agricultural commodity export prices are probably small.[31]

Another potential source of endogeneity relates to the possibility that the estimated long-run coefficients partly reflect expectational and adjustment parameters rather than just the long run parameters of interest (Alogoskoufis and Smith, 1991). In particular, if the expected period *t* level of a right-hand side variable differs from the realized level in period *t-1*, then failing to account for these expectations can cause their effect to be wrongfully attributed to the long run effect of the variable. However, this problem is not likely to be important in our analysis because any effect of expectations is likely to be controlled for by the inclusion of the contemporaneous first-difference of the commodity price indices (a good proxy for the expected change in commodity prices).[32]

[28] In all first-stage regressions of Table 6, columns (2), (3), (5) and (6), the relevant instrument (e.g. for , for , etc.) enters with the expected sign and is significant at 1 percent, while in all but one of the cases, the relevant additional instrument (e.g. for , for , etc.) is also significant. To save space, we do not report these first-stage results.

[29] Since the variables we used to construct the instruments (value of sub-soil assets, latitude, fraction of population speaking English, and fraction of population speaking major European language) are all time-invariant and thus only generate cross-sectional variation, we cluster the (robust) standard errors by country in all specifications of Table 6, using the command "xtivreg2" written for Stata by Mark Schaffer (Schaffer, 2007) to perform 2SLS. Failing to account for within-group correlation of errors in models with mixtures of individual and grouped data can result in estimated standard errors that are biased downwards (Moulton, 1990, Donald and Lang, 2007).

[30] We use the endogeneity test statistic of Baum et al. (2007), which under conditional homoskedasticity is numerically equal to a Hausman test statistic.

[31] The Kleibergen-Paap rk *LM* statistics, reported at the bottom of Table 6, reject the null of underidentification in all four 2SLS specifications, while the Sargan-Hansen test statistics do not reject the null of valid instruments, except in column (2). The joint endogeneity test statistics do not reject the null hypothesis that the specified endogenous regressors can be treated as exogenous.

[32] Alogoskoufis and Smith (1991) point out that long run coefficients can also be biased for reasons other than expectations. They discuss an example in which the estimated error-correction model is subject to multiple interpretations. These do not, however, relate to the

Finally, we consider the possibility that our estimates are biased due to the way in which the commodity export price index was constructed. In particular, as explained in Section 2.1, we take into account only those commodities for which a country is a net exporter. To assess the potential bias from not using information on commodities for which a country is a net *importer*, we constructed a second set of commodity price indices (composite, non-agricultural and agricultural). This set of indices was constructed in the same way as the indices already used in our analysis but instead of using only the commodities for which a country is a net *exporter*, we use only the commodities for which a country is a net *importer*. We then added the relevant net commodity import price variables as additional controls to the specifications of Tables 2 to 5. This ensures that the information on negative net exports is now used in the analysis and any short-run or long-run effects of commodity import prices are controlled for. The results of estimating these augmented specifications (not reported) showed that all our findings for commodity *export* prices are robust to controlling for import prices, while *import* prices themselves do not seem to affect output. This suggests that our estimates are not biased due to the way in which the commodity price index was constructed.

Having addressed concerns of endogeneity, we next test for robustness by separately re-estimating the specification of Table 6, column (3), with each of the twenty-four additional long-run and short-run controls listed in Appendix I for which we have data since 1963 and found that the results are robust.[33]

6. Conclusions

We have adopted panel error correction methodology to estimate the short and long run effects of international commodity prices on output per capita. Our results show that commodity booms have short-term effects on output which are unconditionally positive, but long-term effects which depend upon the type of commodity and the quality of governance. With poor governance, booms in non-agricultural commodities have large adverse long-term effects on output which dominate the short term gains, and which can more than offset the direct gain in income arising through the terms of trade.

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interpretation of the long run coefficients of the right-hand side variables. Since our interest lies in estimating the effect of such right-hand side variables (in particular, the commodity export price indices), the issue of multiple interpretations does not arise.

[33] The long-run coefficient of the non-agricultural export price index in all cases remained negative and significant with values ranging from -1.027 to -2.202, while the long-run coefficient of the interaction of the index with good governance in all cases remained positive and significant with values ranging from 3.063 to 4.502. The short-run coefficients of the index in all cases remained positive as well, while the coefficient of the first lagged difference of the index remained significant.

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Table 1a Commodities

		Non-agricultural		
Aluminum	Gasoline	Natural gas	Phosphatrock	Uranium
Coal	Ironore	Nickel	Silver	Urea
Copper	Lead	Oil	Tin	Zinc
		Agricultural		
Bananas	Cotton	Oliveoil	Pulp	Sugar
Barley	Fish	Oranges	Rice	Sunfloweroil
Butter	Groundnutoil	Palmkerneloil	Rubber	Swinemeat
Cocoabeans	Groundnuts	Palmoil	Sisal	Tea
Coconutoil	Hides	Pepper	Sorghum	Tobacco
Coffee	Jute	Plywood	Soybeanoil	Wheat
Copra	Maize	Poultry	Soybeans	Wool

Table 1b Summary statistics

	Obs.	Mean	St. Dev.	Min.	Max.
Real GDP per capita (log)	4345	7.66	1.56	4.39	10.91
Investment share of GDP	4345	0.21	0.07	-0.24	0.71
Δ Population (log)	4345	0.02	0.01	-0.08	0.16
Secondary schooling (average number of years)	4345	1.78	1.24	0.03	7.48
Trade to GDP	4284	0.70	0.44	0.05	4.38
Population ages 0-14 (share of total population)	4345	0.35	0.10	0.13	0.51
Inflation (log (1 + inflation rate))	3844	0.12	0.27	-0.15	5.48
Commodity export price index ( )	4345	0.36	0.40	0.00	1.97
Unlogged unweighted index, 1980 = 100 (———)	4345	85.33	29.71	15.80	313.42
Net commodity exports to GDP ( )	4345	0.08	0.09	0.00	0.42
Non-agricultural commodity export price index	4345	0.21	0.36	0.00	1.96
Unlogged unweighted non-agri index (1980 = 100)	4345	88.38	33.21	15.71	371.29
Net non-agricultural commodity exports to GDP	4345	0.05	0.09	0.00	0.40
Agricultural commodity export price index	4345	0.16	0.23	0.00	1.03
Unlogged unweighted agri index (1980 = 100)	4345	93.68	28.28	31.46	308.35
Net agricultural commodity exports to GDP	4345	0.04	0.05	0.00	0.22
Dummy good governance	3861	0.28	0.45	0	1
Δ GDP per capita (log)	4345	0.02	0.05	-0.63	0.32
Δ Commodity export price index	4345	0.00	0.02	-0.27	0.33
Δ Unlogged unweighted index (1980 = 100)	4345	0.89	15.89	-119.23	349.04
Coup	4017	0.02	0.15	0	2
Civil war	4023	0.07	0.26	0	1
Natural disaster	3908	0.24	0.57	0	4

Table 2 Estimation results: baseline specifications

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
			Estimates of long-run coefficients				
Investment share of GDP	1.571**	0.290	0.066	2.348***	0.521	0.365	8.374**
	(0.753)	(0.436)	(0.447)	(0.837)	(0.464)	(0.441)	(3.954)
Δ Population (log)	-1.825	8.529	8.881	0.600	8.715	7.955	36.362
	(13.474)	(9.984)	(9.125)	(13.133)	(10.473)	(9.051)	(53.650)
Secondary schooling	0.103	0.013	0.042	0.085	0.046	0.049	0.079
	(0.073)	(0.042)	(0.045)	(0.068)	(0.043)	(0.046)	(0.156)
Trade to GDP		0.554***	0.636***		0.503***	0.532***	1.550**
		(0.130)	(0.146)		(0.116)	(0.136)	(0.583)
Population ages 0-14		-5.401***	-5.419***		-4.364***	-4.651***	-14.414***
		(1.054)	(1.023)		(1.038)	(0.916)	(5.257)
Inflation (log)		-0.306***	-0.240***		-0.243***	-0.201***	-1.547**
		(0.091)	(0.074)		(0.072)	(0.060)	(0.711)
Commodity export price index	-2.341**	-1.765***	-1.989***	-1.705**	-1.475***	-1.329***	-1.838***
	(0.884)	(0.568)	(0.599)	(0.735)	(0.478)	(0.480)	(0.668)
			Estimates of short-run coefficients				
GDP per capita (log)$_{t-1}$	-0.030***	-0.049***	-0.054***	-0.033***	-0.051***	-0.060***	-0.006***
	(0.005)	(0.006)	(0.007)	(0.006)	(0.007)	(0.008)	(0.002)
Δ GDP per capita (log)$_{t-1}$	0.217***	0.222***	0.200***	0.165***	0.195***	0.180***	0.256***
	(0.045)	(0.029)	(0.030)	(0.046)	(0.029)	(0.030)	(0.031)
Δ Commodity export price index$_t$	0.217***	0.140**	0.130*	0.148**	0.074	0.085	0.103
	(0.052)	(0.067)	(0.073)	(0.060)	(0.072)	(0.082)	(0.080)
Δ Commodity export price index$_{t-1}$	0.229***	0.214***	0.221***	0.214***	0.210***	0.230***	0.178***
	(0.046)	(0.051)	(0.050)	(0.039)	(0.049)	(0.052)	(0.048)
Δ Commodity export price index$_{t-2}$	0.118**	0.125	0.123	0.078	0.101	0.111	0.054
	(0.057)	(0.088)	(0.095)	(0.051)	(0.077)	(0.086)	(0.075)
Coup$_t$			-0.031***			-0.032***	
			(0.009)			(0.009)	
Civil war$_t$			-0.010***			-0.010**	
			(0.004)			(0.004)	
Natural disaster$_t$			-0.006***			-0.005***	
			(0.002)			(0.002)	
Country fixed effects	YES	YES	YES	YES	YES	YES	NO
Regional time dummies	NO	NO	NO	YES	YES	YES	YES
Time trend	YES	YES	YES	NO	NO	NO	NO
Observations	4345	3793	3315	4345	3793	3315	3793
R-squared (within)	0.11	0.15	0.17	0.26	0.29	0.30	0.31

Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by year and are reported in parentheses. Columns (1) to (6) report the R-squared *within*, while column (7) reports the R-squared. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 3 Estimation results: decomposing the long-run effect of commodity export prices

	(1)	(2)	(3)	(4)	(5)	(6)
			Estimates of long-run coefficients			
Investment share of GDP	1.596**	0.348	0.119	2.359***	0.553	0.383
	(0.723)	(0.419)	(0.433)	(0.816)	(0.451)	(0.432)
Δ Population (log)	-1.755	7.986	8.358	0.622	8.538	7.761
	(12.937)	(9.585)	(8.783)	(12.763)	(10.219)	(8.866)
Secondary schooling	0.084	0.002	0.026	0.072	0.038	0.039
	(0.068)	(0.040)	(0.043)	(0.066)	(0.042)	(0.045)
Trade to GDP		0.560***	0.650***		0.526***	0.561***
		(0.127)	(0.140)		(0.114)	(0.133)
Population ages 0-14		-5.150***	-5.189***		-4.203***	-4.522***
		(1.020)	(0.991)		(1.017)	(0.904)
Inflation (log)		-0.299***	-0.238***		-0.246***	-0.206***
		(0.087)	(0.072)		(0.071)	(0.059)
Non-agricultural commodity export price index	-2.570***	-1.976***	-2.249***	-1.934**	-1.685***	-1.579***
	(0.845)	(0.558)	(0.582)	(0.727)	(0.481)	(0.486)
Agricultural commodity export price index	1.974	1.707	1.497	2.490	1.936	1.803
	(2.337)	(1.550)	(1.618)	(2.280)	(1.843)	(1.777)
			Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.031***	-0.051***	-0.056***	-0.034***	-0.053***	-0.061***
	(0.006)	(0.007)	(0.007)	(0.006)	(0.007)	(0.008)
Δ GDP per capita $(\log)_{t-1}$	0.217***	0.221***	0.198***	0.165***	0.194***	0.178***
	(0.045)	(0.029)	(0.029)	(0.046)	(0.029)	(0.030)
Δ Commodity export price index_t	0.220***	0.144**	0.131*	0.151**	0.080	0.088
	(0.051)	(0.066)	(0.072)	(0.058)	(0.071)	(0.080)
Δ Commodity export price index_{t-1}	0.226***	0.210***	0.215***	0.211***	0.207***	0.225***
	(0.046)	(0.050)	(0.048)	(0.040)	(0.047)	(0.050)
Δ Commodity export price index_{t-2}	0.116**	0.122	0.117	0.076	0.100	0.106
	(0.057)	(0.089)	(0.097)	(0.051)	(0.078)	(0.087)
Coup_t			-0.031***			-0.032***
			(0.009)			(0.009)
Civil war_t			-0.011***			-0.011**
			(0.004)			(0.004)
Natural disaster_t			-0.006***			-0.005***
			(0.002)			(0.002)
Country fixed effects	YES	YES	YES	YES	YES	YES
Regional time dummies	NO	NO	NO	YES	YES	YES
Time trend	YES	YES	YES	NO	NO	NO
Observations	4345	3793	3315	4345	3793	3315
R-squared within	0.11	0.15	0.17	0.26	0.29	0.30

Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by year and are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 4 Estimation results: the adverse long-term effect of booms conditional on governance

	(1)	(2)	(3)	(4)
	Estimates of long-run coefficients			
Investment share of GDP	-0.200	0.108	-0.091	0.161
	(0.585)	(0.534)	(0.571)	(0.534)
Δ Population (log)	-1.184	-0.024	-1.506	-0.046
	(4.480)	(4.064)	(4.364)	(4.053)
Secondary schooling	-0.027	0.019	-0.041	0.006
	(0.045)	(0.042)	(0.043)	(0.042)
Trade to GDP	0.601***	0.500***	0.614***	0.560***
	(0.139)	(0.138)	(0.135)	(0.144)
Population ages 0-14	-4.824***	-3.541***	-4.608***	-3.434***
	(0.957)	(1.002)	(0.948)	(1.019)
Inflation (log)	-0.338***	-0.262***	-0.329***	-0.270***
	(0.105)	(0.077)	(0.101)	(0.076)
Commodity export price index	-2.034***	-1.704***		
	(0.693)	(0.597)		
Commodity export price index * good governance	1.871*	2.219**		
	(1.113)	(1.035)		
Non-agricultural commodity export price index			-2.212***	-1.919***
			(0.661)	(0.575)
Non-agricultural commodity export price index * good governance			2.827**	3.226***
			(1.169)	(1.005)
Agricultural commodity export price index			1.641	2.588
			(2.131)	(2.728)
Agricultural commodity export price index * good governance			-1.378	-2.752
			(4.270)	(4.265)
	Estimates of short-run coefficients			
GDP per capita (log)$_{t-1}$	-0.043***	-0.048***	-0.045***	-0.048***
	(0.005)	(0.006)	(0.006)	(0.006)
Δ GDP per capita (log)$_{t-1}$	0.239***	0.214***	0.238***	0.211***
	(0.031)	(0.033)	(0.031)	(0.033)
Δ Commodity export price index$_t$	0.147**	0.071	0.148**	0.075
	(0.068)	(0.070)	(0.067)	(0.070)
Δ Commodity export price index$_{t-1}$	0.207***	0.210***	0.203***	0.205***
	(0.052)	(0.052)	(0.051)	(0.051)
Δ Commodity export price index$_{t-2}$	0.124	0.108	0.120	0.104
	(0.094)	(0.091)	(0.096)	(0.092)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	YES	NO	YES
Time trend	YES	NO	YES	NO
Observations	3371	3371	3371	3371
R-squared within	0.17	0.33	0.17	0.34

Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by year and are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 5 Estimation results: subsamples good and bad governance

	(1) bad governance	(2) good governance	(3) bad governance	(4) good governance
	Estimates of long-run coefficients			
Investment share of GDP	0.230	-3.244**	0.343	-3.270
	(0.588)	(1.503)	(0.560)	(2.229)
Δ Population (log)	1.095	-2.923	3.103	-0.079
	(5.493)	(6.085)	(4.852)	(6.810)
Secondary schooling	-0.085	-0.034	-0.019	-0.065
	(0.081)	(0.040)	(0.076)	(0.063)
Trade to GDP	0.639***	0.663**	0.427**	0.982**
	(0.148)	(0.302)	(0.180)	(0.441)
Population ages 0-14	-5.556***	-2.838	-3.675***	-3.351
	(1.118)	(1.913)	(1.148)	(2.653)
Inflation (log)	-0.298***	-3.402**	-0.230***	-2.656
	(0.102)	(1.415)	(0.071)	(1.765)
Non-agricultural commodity export price index	-2.079***	1.398	-1.783***	2.340*
	(0.649)	(1.128)	(0.566)	(1.277)
	Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.047***	-0.030***	-0.053***	-0.026**
	(0.007)	(0.008)	(0.008)	(0.010)
Δ GDP per capita $(\log)_{t-1}$	0.232***	0.307***	0.197***	0.314***
	(0.037)	(0.052)	(0.041)	(0.048)
Δ Non-agricultural commodity export price index_{t}	0.135*	0.163	0.073	0.027
	(0.078)	(0.098)	(0.080)	(0.050)
Δ Non-agricultural commodity export price index_{t-1}	0.185***	0.091**	0.180***	0.159**
	(0.059)	(0.045)	(0.063)	(0.064)
Δ Non-agricultural commodity export price index_{t-2}	0.110	-0.031	0.091	0.010
	(0.105)	(0.097)	(0.107)	(0.103)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	NO	YES	YES
Time trend	YES	YES	NO	NO
Observations	2396	975	2396	975
R-squared within	0.16	0.31	0.33	0.57

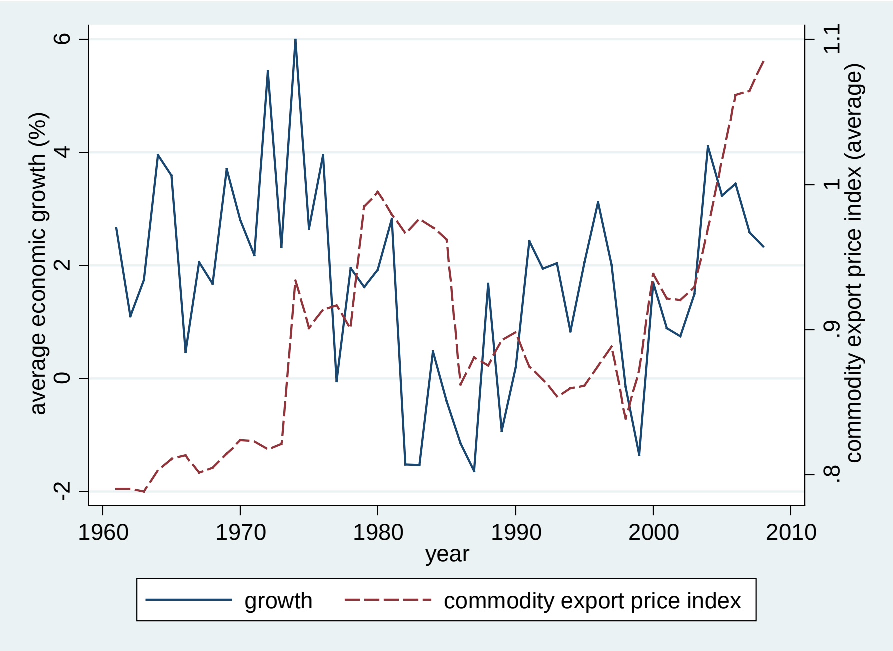
Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by year and are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 6 Estimation results: instrumental variables estimation

	(1)	(2)	(3)	(4)	(5)	(6)
			Estimates of long-run coefficients			
Investment share of GDP	-0.163	0.028	0.049	0.078	0.381	0.304
	(0.630)	(0.630)	(0.623)	(0.584)	(0.532)	(0.541)
Δ Population (log)	-1.583	1.949	1.437	-0.003	1.988	2.284
	(4.795)	(4.965)	(4.966)	(4.461)	(4.301)	(4.339)
Secondary schooling	-0.041	-0.019	-0.036	0.010	0.023	0.016
	(0.057)	(0.049)	(0.048)	(0.052)	(0.041)	(0.040)
Trade to GDP	0.639***	0.716***	0.746***	0.539***	0.527***	0.544***
	(0.147)	(0.129)	(0.128)	(0.155)	(0.129)	(0.127)
Population ages 0-14	-4.799***	-4.025***	-4.248***	-3.570***	-2.899***	-3.010***
	(1.209)	(0.992)	(1.022)	(1.028)	(0.804)	(0.858)
Inflation (log)	-0.333***	-0.229***	-0.249***	-0.262***	-0.164**	-0.165**
	(0.102)	(0.082)	(0.087)	(0.084)	(0.069)	(0.071)
Non-agricultural index	-2.107***	-1.551***	-1.583***	-1.774**	-1.030***	-0.948***
	(0.736)	(0.412)	(0.368)	(0.735)	(0.375)	(0.368)
Non-agricultural index * good governance	2.903**	2.707***	3.277***	3.143***	3.164***	3.376***
	(1.120)	(0.604)	(0.583)	(1.134)	(0.412)	(0.435)
			Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.044***	-0.051***	-0.050***	-0.048***	-0.058***	-0.058***
	(0.006)	(0.006)	(0.006)	(0.007)	(0.008)	(0.008)
Δ GDP per capita $(\log)_{t-1}$	0.244***	0.223***	0.223***	0.216***	0.196***	0.196***
	(0.036)	(0.042)	(0.042)	(0.035)	(0.039)	(0.040)
Δ Non-agricultural index_t	0.140***	0.197***		0.080*	0.116**	
	(0.046)	(0.059)		(0.048)	(0.053)	
Δ Non-agricultural index_{t-1}	0.178***	0.201***	0.178***	0.172***	0.212***	0.188***
	(0.052)	(0.063)	(0.054)	(0.058)	(0.064)	(0.055)
Δ Non-agricultural index_{t-2}	0.104**	0.128**		0.087	0.132**	
	(0.051)	(0.062)		(0.057)	(0.064)	
Country fixed effects	YES	YES	YES	YES	YES	YES
Regional time dummies	NO	NO	NO	YES	YES	YES
Time trend	YES	YES	YES	NO	NO	NO
Method	OLS	2SLS	2SLS	OLS	2SLS	2SLS
Nr. of endogenous regressors		5	3		5	3
Nr. of excluded instruments		9	5		9	5
Kleibergen-Paap rk *LM* statistic		21.6***	19.1***		19.8***	19.9***
Kleibergen-Paap rk Wald *F* statistic		8.2	12.3		6.1	10.4
Sargan-Hansen (*P*-value)		0.02	0.69		0.28	0.18
Endogeneity test (*P*-value)		0.34	0.45		0.59	0.46
Observations	3371	2908	2940	3371	2908	2940
R-squared within	0.17	0.15	0.14	0.33	0.34	0.33

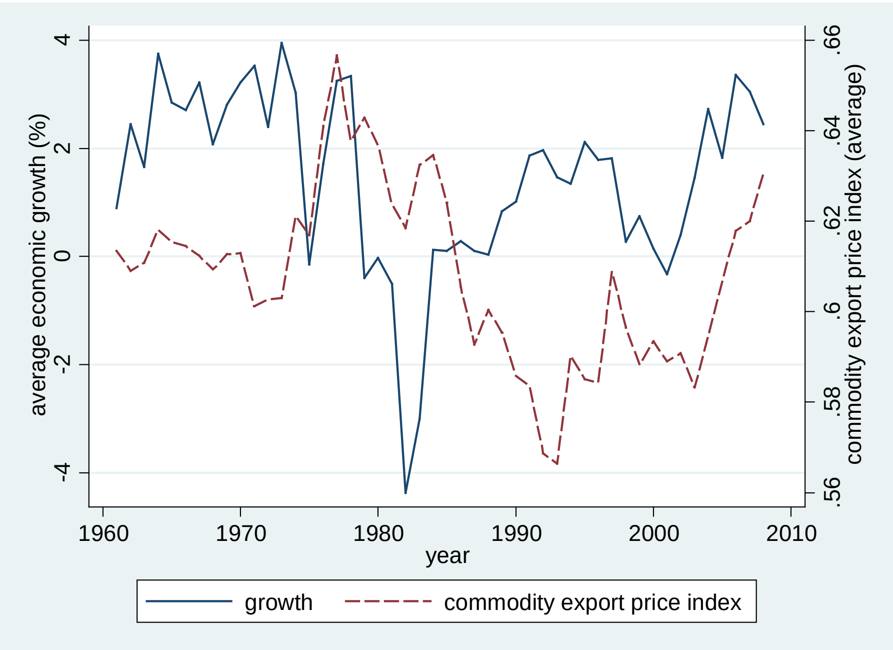
Notes: The dependent variable is the first-differenced log of real GDP per capita. Columns (1) and (4) report OLS results. Columns (2), (3), (5) and (6) report the second-stage results of 2SLS procedures. The Kleibergen-Paap rk *LM* and Wald *F* statistics correspond to tests of underidentification and weak identification, respectively. For the specifications in columns (3) and (6), the critical values of the Kleibergen-Paap rk Wald *F* statistic for a maximal IV relative bias of 5%, 10%, 20% and 30%, are 9.53, 6.61, 4.99 and 4.30, respectively, while for the specifications in columns (2) and (5), the critical values are not available (Stock and Yogo, 2005). The two reported *P*-values correspond to the Sargan-Hansen *J* test of valid overidentifying restrictions and the endogeneity test statistic of Baum et al. (2007). Robust standard errors (clustered by country) are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Figure 1a Commodity prices and growth: Non-agricultural commodity exporters



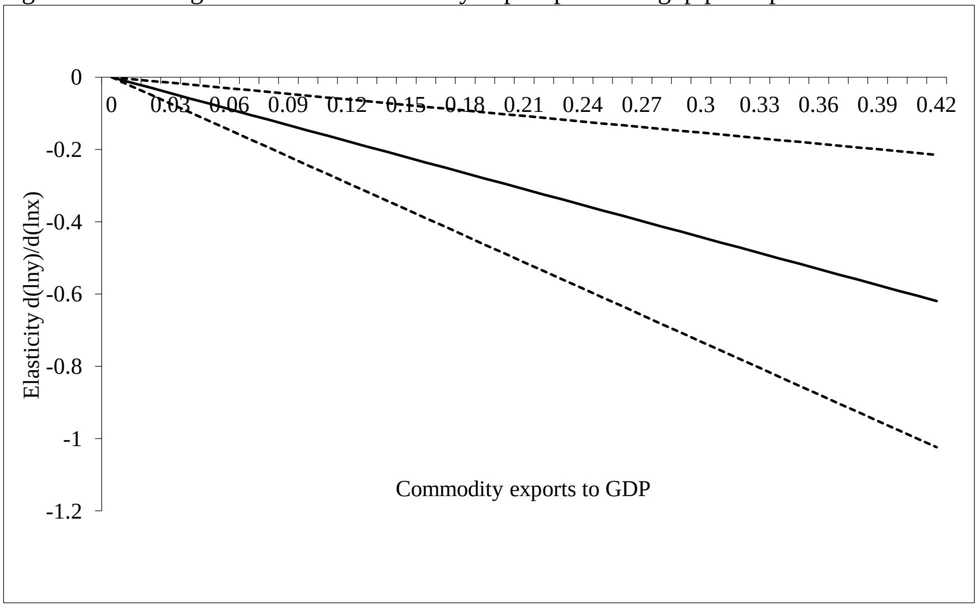
Notes: Figure 1a shows average economic growth ( ) and the average level of the commodity export price index ( ) over the period 1961-2008 for 16 countries: Algeria, Bolivia, Cameroon, Chile, the Republic of Congo, Ecuador, Gabon, Guyana, Malaysia, Norway, Papua New Guinea, Singapore, Syrian Arab Republic, Trinidad and Tobago, Venezuela and Zambia. This selection of countries was based on (i) their presence in our estimation sample, (ii) 1990 non-agricultural commodity exports that constituted more than 10 percent of GDP, and (iii) available data for growth and commodity prices over the entire period 1961-2008.

Figure 1b Commodity prices and growth: Agricultural commodity exporters



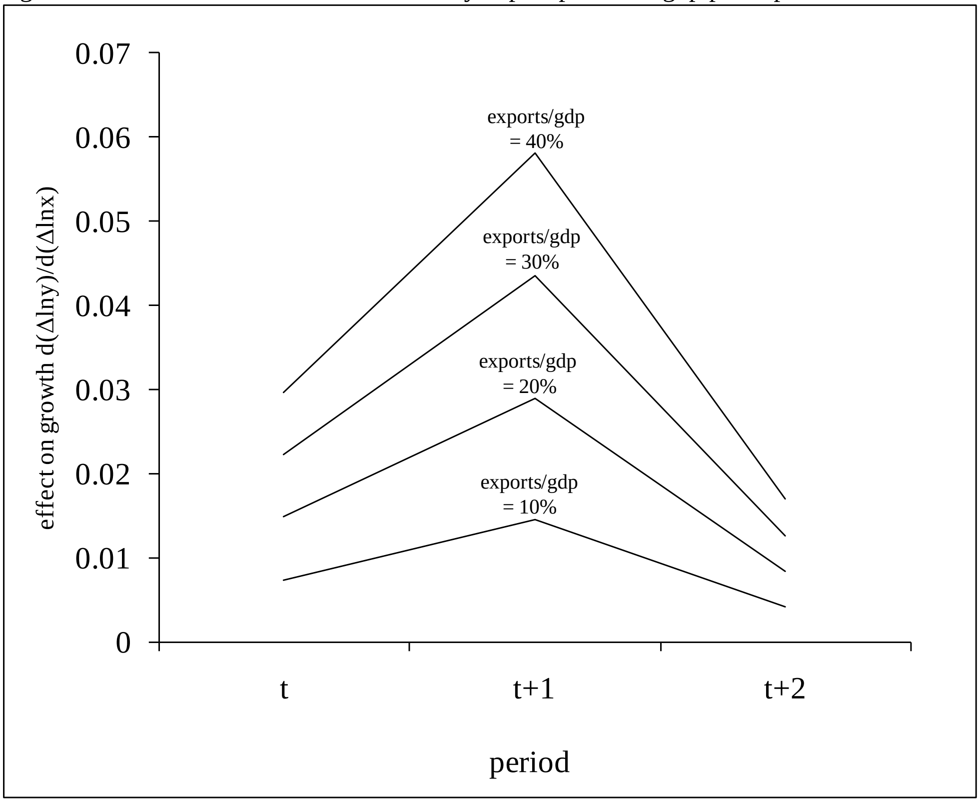
Notes: Figure 1b shows average economic growth ( ) and the average level of the commodity export price index ( ) over the period 1961-2008 for 21 countries: Belize, Burundi, Colombia, Costa Rica, Cote d'Ivoire, Ecuador, El Salvador, Fiji, Ghana, Guatemala, Guyana, Honduras, Kenya, Malawi, Malaysia, Nicaragua, Panama, Papua New Guinea, Paraguay, Sri Lanka and Togo. This selection of countries was based on (i) their presence in our estimation sample, (ii) 1990 agricultural commodity exports that constituted more than 5 percent of GDP, and (iii) available data for growth and commodity prices over the entire period 1961-2008.

Figure 2a The long-run effect of commodity export prices on gdp per capita



Notes: Figure 2a is based on the estimation results in Table 2, column (5). The solid line denotes the elasticity of gdp per capita with respect to commodity export prices. The dashed lines illustrate the 95% confidence interval. The range of values on the horizontal axis corresponds to the range of values in the estimation sample.

Figure 2b The short-run effect of commodity export prices on gdp per capita



Notes: Figure 2b is based on the estimation results in Table 2, column (5). The four lines denote the impulse response functions of an increase in the growth rate of commodity export prices in period t for different levels of commodity exports to GDP. A value of 0.03 on the vertical axis implies that a 10 percentage point increase in the growth rate of commodity export prices leads to a 0.30 percentage point increase in the gdp per capita growth rate.

Appendix I: Data description and sources

Real GDP per capita in constant 2000 US $ (World Development Indicators (WDI))

Commodity export price index Commodity export and import values for 1990 from UNCTAD Commodity Yearbook 2003 and UN International Trade Statistics Yearbook 1993 and 1994. Quarterly commodity price indices from International Financial Statistics (IFS, series 74 for butter and coal, 76 for all others), except for natural gas and gasoline indices, which are from Energy Information Administration’s (EIA) Annual Energy Review 2005 (Column 1 in Tables 5.24 and 6.7). Four price series (coal, plywood, silver, and sorghum) had short gaps in the early periods. Following Dehn (2000), we filled these gaps by holding the price constant at the value of the first available observation. Four price series (palmkerneloil, bananas, tobacco, and silver) had 1, 2, or 3 missing values in the middle. These gaps were filled by linear interpolation. Price series with larger gaps were not adjusted. Where gaps for relatively unimportant commodities (share of net exports in total net exports < 10% or share of net exports in GDP < 1%) would cause missing observations, these price series were left out. Export unit values from IFS (series 74..DZF). from WDI (GDP in current US dollars). The sub-indices for non-agricultural and agricultural commodities were constructed in the same way.[34]

Investment share of GDP gross fixed capital formation as a share of GDP (WDI)

Δ Population (log) (WDI)

Secondary schooling average number of years of secondary schooling for the population aged 15 and above, linearly interpolated (Barro and Lee, 2010)

Good governance 1 for countries with a 1984-2009 mean ICRG composite risk rating of 75 or higher, 0 otherwise (The PRS Group)

A: list of additional long-run controls based on Sala-i-Martin et al.(2004)[35]

Primary schooling average number of years of primary schooling for the population aged 15 and above, linearly interpolated (Barro and Lee, 2010)

Price level of investment from Penn World Tables v.6.3 (log of variable “PI”)

Life expectancy at birth, total (years) (WDI)

Government consumption general government final consumption expenditure as a percentage of GDP (WDI)

[34] To ensure that when replacing the composite commodity export price index by the sub-indices the sample remains the same, we exclude commodities with incomplete time series.

[35] Some of the 67 proxies in Sala-i-Martin et al. (2004) are not listed here because they do not vary over time and hence do not lend themselves for our fixed effects panel estimations.

Real exchange rate overvaluation index (Global Development Network Growth Database)
Trade openness exports plus imports of goods and services as a share of GDP (WDI)
Political rights (Freedom House)
Government share of real GDP per capita from Penn World Tables v.6.3 (variable "kg")
Tertiary schooling average number of years of tertiary schooling for the population aged 15 and above, linearly interpolated (Barro and Lee, 2010)
Public investment as a share of GDP (Global Development Network Growth Database)
Population ages 0-14 as a share of the total population (WDI)
Fertility rate, total (births per woman) (WDI)
Civil liberties (Freedom House)
Number of coups d'etat since 1960 (Banks' Cross-National Time-Series Data Archive)
Number of revolutions since 1960 (Banks' Cross-National Time-Series Data Archive)
Population ages 65 and above as a share of the total population (WDI)
Military expenditure as a share of GDP (WDI)
Population (log) (WDI)
Public spending on education as a share of GDP (WDI)
Inflation log (1 + (annual % change in consumer prices/100)), data from WDI.
Inflation squared inflation   inflation
Number of interstate war years since 1960 (Gleditsch, 2004)
Number of civil war years since 1960 (Gleditsch, 2004)
Oil import price index constructed by interacting the log of the annual average deflated oil price index (see description of the commodity export price index above for sources) with a dummy variable for net oil importers. Net oil imports are crude oil imports plus total imports of refined petroleum products minus crude oil exports minus total exports of refined petroleum products (EIA Annual Energy Review 2002).
Number of large natural disasters since 1960 ($\geq$ 0.5% of pop. affected, or damage $\geq$ 0.5% of GDP, or $\geq$ 1 death per 10000, criteria established by the IMF) Geological disasters: earthquakes, landslides, volcano eruptions, tidal waves; Climatic disasters: floods, droughts, extreme temperatures, wind storms; Human disasters: famines, epidemics (WHO CRED).

B: list of short-run controls

Number of coups d'etat in a given year (Banks' Cross-National Time-Series Data Archive)
Interstate war 1 for interstate war in a given year, 0 otherwise (Gleditsch, 2004)
Civil war 1 for civil war in a given year, 0 otherwise (Gleditsch, 2004)

Number of large natural disasters in a given year (see description above)

Δ Oil import price index contemporaneous, first and second lagged differences in the oil import price index (see description above)

Appendix II: Testing for weak exogeneity and the existence of a long run relationship

A: long run relationship

The estimating equation (2) is only appropriate if there is a long-run levels relationship between GDP per capita and the long-run right-hand side variables. Testing for the existence of a levels relationship is often done using cointegration techniques. However, cointegration requires that all the long-run variables in our specifications are integrated of order 1, which seems unlikely to be the case.[36] We therefore use a new approach to testing the existence of a levels relationship developed by Pesaran et al. (2001), which can be used irrespective of whether the level variables are stationary or non-stationary. This "bounds test" is based on a standard -statistic for the null hypothesis that the coefficients of the lagged level variables, corresponding to and in our estimating equation (2), are equal to zero. Pesaran et al. (2001) show that the asymptotic distribution of the -statistic is non-standard under the null of no level relationship, i.e. : and in equation (2). They report two sets of critical values for the two polar cases in which the lagged level variables in are either all I(1) or all I(0). They then propose a bounds testing procedure. If the computed -statistic lies *below* the two relevant critical values, the null hypothesis of no level relationship cannot be rejected, regardless of whether the variables in are I(1) or I(0). If the -statistic lies *in between* the two critical values, the result is inconclusive and rejection of the null depends on whether the variables in are I(1) or I(0). Finally, if the -statistic lies *above* the two critical values, the null hypothesis is rejected, regardless of whether the variables in are I(1) or I(0). In addition to the bounds test based on the -statistic, Pesaran et al. (2001) propose a second bounds test, based on a standard -statistic for the null hypothesis that the coefficient of the lagged level of the dependent variable, corresponding to in equation (2), is equal to zero. Following Pesaran et al. (2001), we re-estimated the specification of Table 2, column (2), and computed the relevant - and -statistics. Both statistics satisfied the test, being larger than the two relevant critical values corresponding to the 1 percent significance level. We repeated this exercise by separately re-estimating the

[36] Breitung and Das (2005) and Pesaran (2007) second-generation unit root tests for the long-run variables in some cases rejected and in some cases did not reject the null of a unit root.

specification of Table 2, column (2), with 1 lag, 2 lags and 3 lags of *all* the differenced long-run variables and found similar results.[37] Based on these results, we reject the null of no long-run levels relationship irrespective of whether the variables in are I(1) or I(0).

B: weak exogeneity

Estimating equation (2) in a single-equation framework without additional equations for the long-run right-hand side variables is only appropriate if these variables are weakly exogenous. As explained by Urbain (1992) and Enders (2004), a sufficient condition for right-hand side variables to be weakly exogenous for the long-run parameters is that they are not "error-correcting", or in other words, that the right-hand side variables do not themselves "respond to the discrepancy from long-run equilibrium". Engle and Granger (1987) therefore argue that a simple way to test for weak exogeneity is to estimate an error-correction model for each right-hand side variable and test the statistical significance of the speed of adjustment parameter using a traditional *t*-test. If the speed of adjustment parameter is insignificant, the variable does not respond to deviations from long-run equilibrium and can thus be viewed as weakly exogenous. Following Engle and Granger (1987), we test for weak exogeneity by estimating error-correction models for each of the eight long-run variables, i.e. for GDP per capita, the investment share of GDP, population growth, secondary schooling, trade to GDP, population ages 0-14, inflation, and the commodity export price index. Since this involves cross-equation restrictions, we follow Engle and Granger (1987) and Enders (2004) and use the lagged residuals from a long-run equilibrium regression in levels[38] as an instrument for the deviation from long-run steady state equilibrium. In particular, for each long-run variable, we regress the first-difference of that variable on the lagged residual from the equilibrium regression, and several lagged differences of each of the long-run variables, while also including country fixed effects and a time trend. As a first robustness check, we run these eight error-correction models with zero lags, one lag, two lags, and three lags of the differenced long-run variables. As a second robustness check, we rerun the models with the regional time dummies instead of the linear time trend. This yields a total of sixty-four error-correction specifications, eight for each of the eight long-run variables. Following Engle and Granger (1987) and Enders (2004), we use the statistical significance of the speed of

[37] The values of the *F*-statistics were 19.27, 17.22, 12.32 and 9.76, respectively, with corresponding critical values of 3.34 and 4.63 for I(0) and I(1) variables, respectively. The values of the *t*-statistics were -7.62, -6.97, -7.06 and -7.06, respectively, with corresponding critical values of -3.96 and -5.49 for I(0) and I(1) variables, respectively.

[38] Consistent with the long-run equilibrium condition in equation (4), we regress log real GDP per capita on the investment share of GDP, population growth, secondary schooling, trade to GDP, population ages 0-14, log inflation, and the commodity export price index, while also including country fixed effects and a time trend.

adjustment parameter (i.e., the coefficient of the lagged residuals) as a test for weak exogeneity. For the eight error-correction specifications with the first-difference of GDP per capita as the dependent variable, the speed of adjustment parameter is always negative and statistically significant at 1 percent. The size of the coefficients suggests a speed of adjustment of around 5 percent per year. These results confirm that GDP per capita "error corrects", i.e. responds to the discrepancy from long-run equilibrium. The speed of adjustment parameter in the specifications for the other long-run variables is never robustly significant and in the vast majority of cases the size of the coefficient is very close to zero. These results suggest that the variables other than GDP per capita do not respond to deviations from long-run equilibrium and can thus be viewed as weakly exogenous.