Commodity Prices, Growth, and the Natural Resource Curse: Reconciling a Conundrum*

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December, 2009

Abstract

Currently, evidence on the 'resource curse' yields a conundrum. While a large literature describes and explains the curse, initial cross-section econometric results have now been overturned and time series analyses using vector autoregressive (VAR) models have found that commodity booms raise the growth of commodity exporters. This paper adopts panel cointegration methodology to explore longer term effects than permitted using VARs. We find strong evidence of a conditional resource curse. 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 weak 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

* We thank three anonymous referees, Chris Adam, Kofi Adjepong-Boateng, Robin Burgess, Phillip Crowson, Mardi Dungey, Jan Willem Gunning, John Page, Klaus Schmidt-Hebbel, Nicolas Van de Sijpe, Ron Smith, Måns Söderbom, and seminar/conference participants at the WIDER Conference "Aid: Principles, Policies, and Performance", the LSE/UCL Development and Growth seminar, the CSAE Conference, the G-20 Workshop "Commodity Cycles and Financial Stability", the Treasury/DfID seminar "Economics of Africa", the University of the Witwatersrand, the Money, Macro and Finance Research Conference, the joint AfDB-WB Workshop "Good Governance and Sustainable Management of Petroleum and Mineral Resources", the OxCarre Launch Conference, the University of Gothenburg, and the Global Development Network's Conference "Natural Resources and Development" for many helpful comments. We acknowledge support of the UK Economic and Social Research Council and the UK Department for International Development.

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1. Introduction

A large literature suggests that there is a 'resource curse': natural resource abundant countries tend to grow slower than resource scarce countries, although this may be conditional upon country characteristics.[4] However, whereas the resource curse literature predicts a negative effect of commodity booms on growth, empirical studies by Deaton and Miller (1995) for Africa and Raddatz (2007) for low-income countries find quite the contrary: higher commodity prices significantly raise growth.

The resource curse literature and the studies of the effects of commodity prices use different methodologies, but both suffer from acknowledged limitations. The former is largely reliant upon cross-sectional growth regressions in which average growth over recent decades is regressed on a measure of resource abundance and a selection of controls. This methodology is unable to disentangle the dynamics of the resource curse and suffers from potential omitted variable bias. Alexeev and Conrad (2009) show that once allowance is made for some important omitted variables, the unconditional version of the resource curse hypothesis falls apart. For the resource curse to be more than just a series of idiosyncratic events in particular countries, it is therefore "crucial to move from cross-country to panel data evidence" (Van der Ploeg, 2006).[5] However, the approach pioneered by Deaton and Miller (1995), namely vector autoregressive (VAR) models, cannot address long-run effects. The unexplored possibility for a systematic resource curse is thus that these positive short-run effects are followed by others, beyond the horizon of the VAR models, whose sign is

[4] This empirical finding is documented in amongst others Sachs and Warner (1995a, 2001), Gylfason et al. (1999), and Sala-i-Martin and Subramanian (2003). Van der Ploeg (2006) provides a survey of the resource curse literature. Sala-i-Martin et al. (2004) propose a Bayesian Averaging of Classical Estimates (BACE) approach to test the robustness of cross-country growth regression results and find that, contrary to claims made in earlier literature, countries with a large mining sector tend to grow faster. Alexeev and Conrad (2009) and Brunnschweiler and Bulte (2008) also find that natural resources positively affect growth. Finally, Dunning (2008), Mehlum et al. (2006), Robinson et al. (2006), and Collier and Hoeffler (2009) argue that depending on country-specific characteristics such as inequality or institutions, natural resources can be a curse in some countries and a blessing in others.

[5] Lederman and Maloney (2007) and Manzano and Rigobon (2007) use panel data with two to four time series observations and show that the resource curse effect in cross-sectional growth regressions disappears when employing system GMM or fixed effects estimators, respectively.

conditional upon country characteristics and which potentially more than offset initial benefits.

In this paper we adopt panel cointegration methodology to analyze global data for 1963 to 2003 to disentangle the short and long run effects of international commodity prices on output per capita. An advantage of using international commodity prices to analyze the effects of natural resources is that they are typically unaffected by the behaviour of individual countries (Deaton and Miller, 1995), although we relax this assumption when we address concerns over endogeneity. Our estimations include country fixed effects and regional time dummies to control for unobserved heterogeneity 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.

We find strong evidence in support of the conditional resource curse hypothesis. 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.[6] Within this group, we find that the resource curse is avoided by countries with sufficiently good 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 resource curse 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 the Appendix. The short-run and long-run effects of

[6] The effect of different types of resources was earlier studied by Boschini et al. (2007). Using cross-sectional growth regressions, they find that in countries where resources are highly appropriable, as determined by both the type of resources and institutional quality, resource abundance lowers growth, while in countries with less appropriable resources, it promotes growth.

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 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

$$\Delta y_{i,t} = \lambda y_{i,t-1} + \beta_1' x_{i,t-1} + \alpha_i + \delta t + u_{i,t} \tag{1}$$

for $i = 1, \dots N$ and $t = 1, \dots T$, where $y_{i,t}$ denotes the logarithm of real GDP per capita in country i in year t, $\Delta y_{i,t}$ is the growth rate of real GDP per capita between $t-1$ and t, $x_{i,t-1}$ is an $m \times 1$ vector of m variables that are expected to affect the long-run steady state level of GDP per capita, α_i is a country-specific fixed effect, and t 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 $x_{i,t}$ and an $n \times 1$ vector $s_{i,t}$ of n control variables that are expected to have only a short-run effect on growth. We also add a

[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.

lagged dependent variable to account for persistence in growth rates.[8] This results in the following core estimating equation of our empirical analysis:

$$\Delta y_{i,t} = \lambda y_{i,t-1} + \beta_1' x_{i,t-1} + \beta_2 \Delta y_{i,t-1} + \sum_{j=0}^{k} \beta_{3j}' \Delta x_{i,t-j} + \beta_4' s_{i,t} + \alpha_i + \delta t + u_{i,t} \quad (2)$$

Equation (2) above can be rewritten as an error correction model, thereby distinguishing between the short- and long-run effects of the right-hand side variables on output:

$$\Delta y_{i,t} = a_1\left(y_{i,t-1} - \theta' x_{i,t-1} - \mu_i - gt\right) + a_2 \Delta y_{i,t-1} + \sum_{j=0}^{k} a_{3j}' \Delta x_{i,t-j} + a_4' s_{i,t} + a_i + a_5 t + u_{i,t} \quad (3)$$

where $\lambda = a_1$, $\beta_1 = -a_1\theta$, $\beta_2 = a_2$, $\beta_{3j} = a_{3j}$ for $j = 0, \dots k$, $\beta_4 = \alpha_4$, $\alpha_i = a_i - a_1\mu_i$, and $\delta = a_5 - a_1 g$. In equation (3) above, output responds to the deviation from long-run steady state equilibrium, captured by the term between brackets, $y_{i,t-1} - \theta' x_{i,t-1} - \mu_i - gt$. Everything else equal, if this deviation is positive, so that $y_{i,t-1} - \theta' x_{i,t-1} - \mu_i - gt > 0$, output will fall. Alternatively, if output lays below its steady state level, so that $y_{i,t-1} - \theta' x_{i,t-1} - \mu_i - gt < 0$, it will rise. In other words, output "error-corrects", i.e. responds to deviations from equilibrium in a way that gradually brings the economy back to its long-run equilibrium. This error-correction process implies that the coefficient a_1 in equation (3), which equals the coefficient λ in equation (2), should be negative, while the size of this coefficient captures the speed with which the economy returns to its long-run equilibrium, or in other words 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 $y_{i,t-1} = \theta' x_{i,t-1} + \mu_i + gt$. If we assume that in long-run equilibrium the determinants of output take

[8] The fixed effects (within groups) estimator is biased in "small T, large N" panels with explanatory variables that are not strictly exogenous, such as a lagged dependent variable (Nickell, 1981). However, this bias becomes negligible as T grows large. Bond (2006), based on calculations of this inconsistency, and Monte Carlo experiments, concludes that the bias poses a huge problem with $T < 10$, remains non-negligible for $T = 10$ or $T = 15$, and can quite comfortably be ignored when $T = 30$ or $T = 40$. The average number of time series observations in the core specifications of our analysis ranges from 28 to 36, suggesting that the bias is small. However, as part of our discussion of endogeneity in section 5, we show that our results on the short-run and long-run effects of commodity export prices on output are robust and even become stronger when making the bias arguably negligible by excluding all countries with fewer than 30 time series observations. We also briefly discuss alternative dynamic panel estimators.

the constant value x_i, the steady state growth rate is given by g so that $y_{i,t} = y_{i,t-1} + g$, and the long-run equilibrium condition can be written as

$$y_{i,t} = \theta' x_i + \mu_i + g + gt \tag{4}$$

Equations (3) and (4) above show how the growth regression 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 $x_{i,t-1}$ in equation (2) are captured by the coefficient vector θ' in equations (3) and (4) and, given that $\theta' = -\frac{\beta_1'}{\lambda}$, can be computed from the estimated coefficients in equation (2). By contrast, the short-run effects directly follow from the estimated coefficients in equation (2): λ for the speed of convergence, β_2 for the short-run effect of growth in the previous year, β_{3j}' $(j = 0, \dots k)$ for the short-run effect of changes in the x-variables, and β_4' for the short-run effect of the control variables in $s_{i,t}$. In addition, equation (3) and (4) also emphasize the importance of the country-specific fixed effect and the time trend in equation (2). The fixed effect, α_i, controls for any country-specific time-invariant unobserved variables that affect the steady state level of output, as can be seen from μ_i 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 a_i in equation (3). The time trend t, on the other hand, allows the steady state growth rate g 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 a_5 in equation (3).

Three comments are in order. First, the inclusion of a simple linear time trend in equation (2) above restricts the steady state growth rate (g) 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 $rT \times 1$ vector of regional time dummies instead of a time trend, where r 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.[9] 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, unless stated otherwise, we compute robust standard errors clustered by year in all our estimations 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.

Thirdly, 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. In particular, as described in Section 2.1 below, we include measures of political shocks, such as coups and civil wars, and natural shocks, such as geological, climatic, and human disasters. We also control for shocks to trade openness, inflation, and international reserves, as well as for the effect of oil price shocks on the output of oil-*importing* economies. And finally, as already discussed, we include separate year-specific fixed effects for seven geographical regions to

[9] This categorization is based on the country classifications of the World Bank and the United Nations, and on the online Central and Eastern European Directory.

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 vectors $x_{i,t}$ and $s_{i,t}$, including our indicators of commodity export prices.

2.1 The variables used in estimation

As explained above, the vector $x_{i,t-1}$ in equation (2) includes variables that are expected to affect GDP per capita both in the short run and long run. First of all, it includes a commodity export price index constructed using the methodology of Deaton and Miller (1995) and Dehn (2000). We collected data on world commodity prices and commodity export and import values for as many commodities as data availability allowed. Table 1a lists the 50 commodities in our sample. For each country, we calculated the total net export value (exports minus imports) of all commodities for which the country is a net exporter. We constructed weights by dividing a country's individual 1990 net export value for each commodity by this total. These 1990 weights are then held fixed over time and applied to the quarterly world price indices of the same commodities to form a geometrically weighted index. We deflated this quarterly index by the export unit value and then calculated the log of its annual average (rescaled so that 1980 = 100). This resulted in an annual country-specific logged index of commodity export prices, $P_{i,t}$. To allow the effect of commodity export prices to be larger for countries with larger exports, we weight $P_{i,t}$ by the 1990 share of net commodity exports in GDP, denoted by w_i, and use the weighted index, $w_i P_{i,t}$, 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 $w_i^N P_{i,t}^N$ and $w_i^A P_{i,t}^A$, respectively, where the superscripts N and A stand for non-agricultural and agricultural commodities. We also include an oil *import* price index to control for the effect of oil prices on oil importing countries. This index was constructed by interacting the log of the deflated oil price index with a dummy variable that takes a value of one if a country is a net oil importer and zero otherwise.

Finally, we include three control variables taken from the empirical growth literature: trade openness, measured as the ratio of trade to GDP; inflation, measured as the log of 1 plus the annual consumer price inflation rate; and international reserves over GDP. Clearly, the selection of control variables is an important issue. As we show, our results are robust to a wide range of additional or alternative controls used in the literature, including indicators of institutional quality, conflict, commodity price volatility, industrial development, investment (as suggested by the empirical studies of the neoclassical growth model), public and private consumption, democracy, the black market premium, the number of assassinations, an alternative measure of trade openness, and exchange rate overvaluation.[10]

The vector $s_{i,t}$ in equation (2) includes control variables that are expected to have only a short-run effect on growth. It includes indicators that capture civil war, the number of coup d'etats, and the number of large natural disasters (geological, climatic, and human disasters).

Our dataset consists of all countries and years for which data are available, and covers around 130 countries between 1963 and 2003. Table 1b reports summary statistics.

2.2 Testing for the existence of a long run relationship

The estimating equation (2) is only appropriate if there is a long-run level relationship between GDP per capita, commodity export prices, oil import prices, trade openness,

[10] These variables are not included in our preferred specification because they were either not robustly significant or severely lowered the number of observations. We include them in section 5 when we address endogeneity concerns. The growth literature also uses a number of time-invariant variables, such as indicators of geography. However, any effect of these variables is already captured by the fixed effects.

inflation, and international reserves. Testing for the existence of a level relationship is often done using cointegration techniques. Cointegration requires that the individual variables are integrated of order 1, and that the residuals of a levels regression of GDP per capita on the other five variables are stationary. We tested these requirements using the Im, Pesaran and Shin (2003) and Maddala and Wu (1999) panel unit root tests and the Pedroni (1999) panel cointegration test.[11] For the first-differenced variables, the unit root tests always rejected the null of non-stationarity at the 1 percent significance level, which confirms that the variables are stationary in differences. For the levels, the Im, Pesaran and Shin (2003) test did not reject non-stationarity, except for inflation, while the Maddala and Wu (1999) test rejected non-stationarity for most of the variables. It is important to note that rejection of the null means that at least one of the series is stationary. It is therefore possible that the tests reject non-stationarity while most of the series are in fact non-stationary. To determine the proportion of countries for which non-stationarity is rejected, we performed separate augmented Dickey-Fuller (ADF) tests for individual countries. The results showed that for 82 to 92 percent of the countries, the ADF tests do not reject non-stationarity of the levels variables, while rejecting non-stationarity of the differenced variables for 70 to 90 percent of the countries. This suggests that the variables are integrated of order 1.

We next performed the Pedroni (1999) panel cointegration test. We first ran a levels regression of GDP per capita on the other five long-run variables and a time trend for each country separately and collected the residuals. We then ran ADF-type regressions for the residuals, again for each country separately. Following Pedroni (1999), we allow the lag order of the dependent variable in the ADF regressions to vary across countries by including the lags that enter significant at 10 percent. We then calculated the mean ADF t-statistic,

[11] Since the Im, Pesaran and Shin (2003) test requires a balanced sample, we apply it to a subsample of 40 countries and 42 years. By contrast, the Maddala and Wu (1999) test, which does not require a balanced sample, is applied to both the balanced subsample and the full unbalanced sample of observations for which we have data on all six variables. Since the oil import price index equals either zero (for net oil exporters) or the country-invariant world oil price index (for net oil importers), a panel unit root test is not appropriate. Instead, we use a Dickey-Fuller test to examine whether the world oil price index contains a unit root.

derived the "group t-statistic", and expressed it in the form of equation (2) on p. 665 in Pedroni (1999). We found a value of -4.53 for this standard normally distributed statistic and hence rejected the null hypothesis of no cointegration at the 1 percent significance level.[12] The results of the unit root and cointegration tests are consistent with the existence of a long-run levels relationship and suggest that the estimating equation (2) is appropriate.

A potential problem with the use of cointegration methods in applied research is that they require knowledge about the time series properties of the underlying variables. Although these properties can be tested, as we have done above, the tests are not without problems and introduce additional uncertainty into the analysis of levels relationships. We therefore supplement them with 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. Following Pesaran et al. (2001), we estimated equation (2) with 1 lag, 2 lags, and 3 lags of the differenced long-run variables and computed the F- and t-statistics. In all three cases, the values of both statistics satisfied the test, being larger than the two relevant critical values corresponding to the 1 percent significance level.[13] As a result, we reject the null of no long-run level relationship. This is reassuring as it confirms the existence of a level relationship regardless of whether the variables in $x_{i,t-1}$ are I(1) or I(0). Hence, the estimating equation (2) is appropriate and inference can be drawn from it, even if one doubts the conclusiveness of the unit root and cointegration test results.

[12] For robustness, we also performed the panel cointegration test without a time trend in the levels regressions and found a similar result.

[13] This "bounds test" is based on a standard F-statistic for the null hypothesis that the coefficients of the lagged level variables, corresponding to $y_{i,t-1}$ and $x_{i,t-1}$ in our estimating equation (2), are equal to zero. Pesaran et al. (2001) show that the asymptotic distribution of the F-statistic is non-standard under the null of no level relationship, i.e. H_0: $\lambda = 0$ and $\beta_1' = 0'$ in equation (2). They report two sets of critical values for the two polar cases in which the lagged level variables in $x_{i,t-1}$ are either all I(1) or all I(0). They then propose a bounds testing procedure. If the computed F-statistic lies *below* the two relevant critical values, the null hypothesis of no level relationship cannot be rejected, regardless of whether the variables in $x_{i,t-1}$ are I(1) or I(0). If the F-statistic lies *in between* the two critical values, the result is inconclusive and rejection of the null depends on whether the variables in $x_{i,t-1}$ are I(1) or I(0). Finally, if the F-statistic lies *above* the two critical values, the null hypothesis is rejected, regardless of whether the variables in $x_{i,t-1}$ are I(1) or I(0). In addition to the bounds test based on the F-statistic, Pesaran et al. (2001) propose a second bounds test, based on a standard t-statistic for the null hypothesis that the coefficient of the lagged level of the dependent variable, corresponding to $y_{i,t-1}$ in equation (2), is equal to zero. On the former test the values of the F-statistics were 17.29, 13.67, and 11.50, with corresponding critical values of 3.93 and 5.23 for I(0) and I(1) variables, respectively. The values of the t-statistics were -7.13, -7.02, and -6.65, with corresponding critical values of -3.96 and -5.13.

2.3 Testing for 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 six long-run variables, i.e. for GDP per capita, trade to GDP, inflation, reserves to GDP, the commodity export price index, and the oil import 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[14] 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, the short-run control variables (civil wars, coup d'états, and natural disasters), 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 six error-correction models with one lag, two lags, and three lags of the differenced long-run variables.[15] As a second robustness

[14] Consistent with the long-run equilibrium condition in equation (4), we regress log real GDP per capita on trade to GDP, inflation, reserves to GDP, the commodity export price index, and the oil import price index, while also including country fixed effects and a time trend.
[15] The lag structure of the differenced long-run variables in our empirical specifications varies from 1 lag for differenced GDP per capita to up to three lags for the other differenced variables.

check, we rerun the models with the regional time dummies instead of the linear time trend. This yields a total of thirty-six error-correction specifications, six for each of the six long-run variables. Following Engle and Granger (1987) and Enders (2004), we use the statistical significance of the speed of adjustment parameter (i.e., the coefficient of the lagged residuals) as a test for weak exogeneity. For the six 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. In the absence of other long-run variables that do the adjustment, this is a necessary condition for a cointegration relationship (Enders, 2004). The weak exogeneity tests for the other long-run variables can be used to assess whether this relationship can be estimated in a single-equation error-correction framework, or whether we need to estimate an error-correction model with more than one equation. The speed of adjustment parameter in the specifications for the other long-run variables is never significant at 5 percent and only in one out of the thirty cases significant at 10 percent. These results clearly indicate 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. As a result, equation (2) can be estimated in a single-equation framework.

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

Table 2 reports the results of estimating equation (2).[16] The first specification includes the commodity export price index. The long-run coefficient is negative and statistically significant at 1 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

[16] As explained in the previous section, the estimated long-run coefficients correspond to $\theta' = -\frac{\beta_1'}{\lambda}$, while the short-run coefficients correspond to λ, β_2, β_{3j}' $(j = 0, \dots k)$ and β_4'.

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 2, column (2), shows the results when we replace the composite index in column (1) by the two sub-indices. For non-agricultural commodities we again find strong evidence of an adverse long-run effect. The coefficient is negative and again significant at 1 percent.[17] By contrast, the coefficient for agricultural commodity export prices is positive and significant at 10 percent. This suggests that higher agricultural export prices are not a curse analogous to non-agricultural commodities: on the contrary, they are more likely than not to be beneficial.

Table 2, column (3), reports the results when replacing the trend in the specification of column (1) by the regional time dummies. The coefficient of the commodity export price index again enters negative and is statistically significant at 5 percent. The coefficient is slightly smaller than in column (1) but implies a substantial long-run resource curse effect. Figure 1a 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 1a therefore predict a long-run elasticity of -0.44.[18] In other words, a 10 percent increase in the prices of their commodity exports leads to a 4.4 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 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

[17] Given the economic importance of oil, we experimented with a further decomposition of non-agricultural commodities into oil and other non-agricultural commodities. An F-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.

[18] Recall that the commodity export price index ($w_i P_{i,t}$) is weighted by the share of net commodity exports in GDP (w_i). So for Zambia and Nigeria, the long-run elasticity equals the long-run coefficient, -1.243, multiplied by the share of net commodity exports in GDP, 0.35.

gain directly raises income by 3.5 percent for given output. Even so, this is less than the decline in output of 4.4 percent, so that the resource curse ends up reducing both output and income relative to counterfactual.

When replacing the composite index by the sub-indices in column (4), the results are also similar to before. The coefficient of the non-agricultural export price index is negative and again significant at 1 percent. For Zambia and Nigeria, whose commodity exports are overwhelmingly non-agricultural, the results predict a long-run elasticity of -0.49. In other words, a 10 percent increase in the price of oil leads to a 4.9 percent lower long-run level of GDP. The coefficient of the agricultural export price index is positive but now insignificant, which is consistent with the absence of a resource curse effect for agricultural commodities.

Having discussed the long-run effects of commodity export prices, we now turn to the other variables in our model. To save space, we only discuss the results in Table 2, column (3). First, the four long-run control variables are statistically significant and enter with the expected signs. Trade to GDP and reserves to GDP enter with a positive sign and are statistically significant at the 1% level, indicating that countries with higher levels of trade liberalization and international reserves tend to have higher long-run GDP levels. Inflation and the oil import price index enter negative and are significant at 5 percent, suggesting that higher inflation and higher oil import prices lead to a lower long-run GDP level.

The coefficient of lagged GDP per capita also has the expected sign and is negative and significant at 1 percent. The size of the 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 6.2 percent per year. Hence, the large adverse long-run effect of higher non-agricultural export prices on output is not instantaneous but comes about gradually. After the non-agricultural price increase, output "corrects" by 6.2 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.

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 largest and statistically significant at 1 percent for the first lag. 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 1b shows the impulse response functions of an increase in the growth rate of commodity export prices for different levels of commodity exports to GDP. In both the year of the price increase and the subsequent year, 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 0.17 percentage points of GDP growth after year t+1 in countries with commodity exports that represent 10 percent of their GDP. This growth gain amounts to 0.34, 0.51, and 0.68 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).[19] 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.

[19] 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.

Table 2, column (3), also reports the coefficients of the other short-run GDP determinants. The coefficient of the lagged dependent variable is positive and significant at 1 percent. We experimented with additional lags but found that these are unimportant. The lagged changes of trade to GDP, inflation and reserves have the expected signs but are not significant.[20] An increase in the oil price has a negative effect on growth in oil importing countries in the same year and the second subsequent year, and a positive effect in the first subsequent year, although these effects are not significant.[21] Next, coups and civil wars have unsurprisingly large and 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 2.3 percentage points 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 0.4 percentage points.

The specifications in Table 2, columns (1) to (4), all include country fixed effects to control for unobserved heterogeneity. To assess the importance of this heterogeneity, Table 2, column (5), for comparison reports the results when excluding the fixed effects from the specification in the previous column. The coefficient of the lagged dependent variable is considerably higher than in the fixed effects specification of column (4). 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 (without fixed effects) remains inconsistent for panels with large T, such as ours. The coefficient of the lagged level of GDP per capita is also substantially higher than in the previous column and also likely to be biased upwards due to a

[20] We do not include the contemporaneous changes in order to limit concerns of endogeneity.
[21] We include the changes in the oil import price index because the commodity export price index also enters with up to two lags.

positive correlation with the individual effects. The small size of this coefficient and its statistical insignificance cause the long-run coefficients to be statistically insignificant and much larger than in the previous column. The higher coefficients of the lagged dependent variable and the lagged level of GDP per capita in Table 2, column (5), indicate the presence of substantial 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.

4. The resource curse 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 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 resource curse 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 2002.[22] 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

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

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 Nigeria.

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 the resource curse. We introduce governance by adding an interaction term of the commodity price index with a dummy that takes a value of 1 for good governance countries and 0 for bad governance countries to the specifications in Table 2, columns (1) to (4). The results are reported in Table 3.[23] In column (1), the commodity export price index enters negative and is statistically significant at 1 percent, indicating that there is indeed a long-run resource curse effect for countries with bad governance. The interaction term of the index with the good governance dummy enters positive but at this stage is not statistically significant.

In Table 3, column (2), 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 statistically 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 1 percent. 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

[23] Since we only include countries for which the mean ICRG score is available, the number of observations drops from 3579 to 3058.

coefficients, which is positive (although not significant). This suggests that countries with good governance do not suffer from a resource curse and may even be able to transform commodity booms into sustainable higher output. These findings support the hypothesis that the resource curse occurs conditional on bad governance. The agricultural index enters positive and is significant at 5 percent, while its interaction with good governance enters negative but is not significant. 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 long-run resource curse effect and are more likely than not to be beneficial.

In Table 3, columns (3) and (4), we again replace the trend in the specifications of columns (1) and (2) by the regional time dummies. Results are similar. In column (3), the composite commodity export price index again enters negative and is significant at 5 percent, while its interaction with good governance is again positive but is now significant at 5 percent. In column (4), the non-agricultural index enters with a negative sign and is significant at 1 percent, while its interaction with the good governance dummy enters positive and is also significant at 1 percent. These results support the findings in columns (1) and (2) and suggest a resource curse conditional upon governance. The agricultural index enters positive but is now insignificant, while its interaction enters negative and is also insignificant, as in column (2).

We next investigate the robustness of these findings by rerunning the specifications in Table 3 using the initial 1985 composite ICRG scores rather than the average scores.[24] The results are similar. In particular, the results for the composite index and the two sub-indices are robust to using this alternative measure of governance.

[24] 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 > 69.5) and "bad governance" (1985 score ≤ 69.5). The proportion of good governance countries is equal across the average ICRG and 1985 ICRG samples (21%).

Finally, to further explore the non-linear effect of non-agricultural commodity export prices, Table 4 reports the results of separate regressions for the countries with bad governance and the countries with good governance. Columns (1) and (3) show the results for the sub-sample of bad governance countries when including a time trend and regional time dummies, respectively. In both cases the non-agricultural index enters with a negative sign, while the coefficient is significant at 1 and 5 percent, respectively. This is consistent with the earlier finding of a resource curse for countries with bad governance. Table 4, columns (2) and (4), show the results for the sub-sample of countries with good governance. In both cases, the non-agricultural index now enters positive, although not significant. These results support our earlier finding that the resource curse effect 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 not significant and should therefore be viewed with caution, their size suggests that the effect is substantial. For a country like Norway, which in 1990 had non-agricultural commodity exports that represented 15 percent of its GDP, the results in Table 4, columns (2) and (4), predict a long-run elasticity of around 0.28. In other words, a 10 percent increase in the price of non-agricultural commodities leads to a 2.8 percent higher long-run level of Norway's GDP per capita.[25] These results provide evidence that the resource curse occurs conditional on bad governance. Countries with sufficiently good governance do not suffer from the curse, and instead may even benefit from higher commodity prices, both in the short run and in the long run.

5. The endogeneity of resource dependence and governance

A possible concern with the results in the previous sections is that the commodity export price index, $w_i P_{i,t}$, is endogenous, i.e. correlated with the error term in equation (2). Let us

[25] The results in Table 4 are robust to using the initial 1985 composite ICRG scores instead of the average scores.

first consider the potential endogeneity of $P_{i,t}$. 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.[26] 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 128 countries in our sample, 22 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 and 3 without these 22 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.[27]

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

[26] Keeping the weights constant over time 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.

[27] We repeated this robustness check using a threshold of 10 percent instead of 20 percent. 34 out of the 128 countries export at least one commodity for which their share in world exports exceeds 10 percent. Again, our findings in Tables 2 and 3 were for the most part robust to the exclusion of these 34 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.

appear to generate the curse.[28] 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, w_i^N, only enters our specifications as part of the non-agricultural price index, $w_i^N P_{i,t}^N$, we use $\widehat{w}_i^N P_{i,t}^N$ as an instrument for $w_i^N P_{i,t}^N$, where $\widehat{w}_i^N$ is the 2000 value of sub-soil assets. For $\widehat{w}_i^N P_{i,t}^N$ to be a valid instrument, it should be correlated with $w_i^N P_{i,t}^N$ and not correlated with the error term. The first requirement is likely to be met, as a country can only be a net exporter of commodities that are available in the country. The second requirement is less likely to be fulfilled. Everything else equal, 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 $P_{i,t}^N$ by the value of sub-soil assets per capita, $\widehat{w}_i^N$, 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, $w_i^N P_{i,t}^N$, we also need to instrument for its interaction with good governance, $G_i \times w_i^N P_{i,t}^N$, where G_i 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), 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

[28] 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.

the governance dummy, G_i, on these three variables for the 97 countries in the sample of Table 3 for which we have data. Latitude and the fraction of the population speaking English enter with the expected positive signs and are statistically significant at 1 percent and 5 percent, respectively, while the fraction of the population speaking one of the major languages of Western Europe also enters with the expected positive sign but is not significant. The pseudo R-squared of the probit regression is 0.61. Letting $\widehat{G}_i$ denote the fitted values of the probit regression, extrapolated to all years in the sample of Table 3, we then interact $\widehat{G}_i$ with the instrument for the non-agricultural price index, $\widehat{w}_i^N P_{i,t}^N$, discussed above.[29] This yields an additional instrument, $\widehat{G}_i \times \widehat{w}_i^N P_{i,t}^N$, which we use to instrument for the interaction of the non-agricultural index with the good governance dummy, $G_i \times w_i^N P_{i,t}^N$. We next use the constructed instruments to perform two-stage-least-squares estimation. For comparison, Table 5, columns (1) and (3), first report the OLS results. The short and long run effects of non-agricultural commodity prices are consistent with the results in Table 3. Table 5, columns (2) and (4), report the second-stage results of a 2SLS procedure in which we instrument for the level and differences of the non-agricultural index, $w_i^N P_{i,t}^N$, and the interaction of the index with the dummy for good governance, $G_i \times w_i^N P_{i,t}^N$. As instruments we use the corresponding level and differences of the instrument for the non-agricultural index, $\widehat{w}_i^N P_{i,t}^N$, and the instrument for the interaction of the index with good governance, $\widehat{G}_i \times \widehat{w}_i^N P_{i,t}^N$.[30] The long-run coefficient of the non-agricultural export price index is negative and significant at 1 percent in both columns.[31] The size of the coefficients is similar to the size of the coefficients in columns (1) and (3), although somewhat larger, indicating that if

[29] Goderis and Ioannidou (2008) perform a similar procedure to construct instruments, following Wooldridge (2002), p. 237.

[30] In all first-stage regressions, the relevant instrument enters with the expected sign and is significant at 5 percent, while in most cases it is significant at 1 percent. To save space, we do not report these first-stage results.

[31] 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 5, 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).

there is an endogeneity bias, it is likely to be small and, if anything, leads to an underestimation of the resource curse effect. Endogeneity tests[32] did not reject the null hypothesis of consistent OLS estimates for the non-agricultural export price index in columns (2) and (4) with p-values of 0.58 and 0.46, respectively. 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 at most marginal. The coefficients of the interaction of the index with the good governance dummy are similar to the coefficients in columns (1) and (3), and are significant at 1 percent. As previously, endogeneity tests did not reject the null of exogeneity with p-values of 0.94 and 0.91, respectively. 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 (3), while endogeneity tests did again not reject the null of exogeneity. This suggests that any bias is likely to be small and if anything leads to a small underestimation of the positive short-run growth effect of higher non-agricultural export prices. These results indicate that the OLS estimates of the short- and long-run effects of non-agricultural commodity export prices are consistent. We next use the OLS specification of Table 5, column (3), to investigate three other sources of endogeneity.

First, our estimates may suffer from dynamic panel bias (Nickell, 1981). In particular, the fixed effects (within groups) estimator that we employ in our analysis requires strict exogeneity, i.e. the explanatory variables are not allowed to depend upon current, future and past values of the idiosyncratic error term. This assumption is necessarily violated in the presence of lagged dependent variables, as these are driven by past shocks to the idiosyncratic error term. However, as explained by Bond (2006), Roodman (2008) and Smith (2006), the assumption of strict exogeneity is crucial for asymptotic properties in the case where N tends to infinity with T fixed, but not in the case where T tends to infinity. Hence,

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

the fixed effects estimator is consistent as the time dimension of the panel becomes large. In fact, as explained by Bond (2006), Nickell (1981), Smith and Fuertes (2006), and Wooldridge (2002), for a coefficient on the lagged dependent variable that is smaller than 1, the bias in the fixed effects estimator is of order $\frac{1}{T-1}$. Bond (2006), based on calculations of this inconsistency, and Monte Carlo experiments, concludes that the bias poses a huge problem with $T < 10$, remains non-negligible for $T = 10$ or $T = 15$, and can quite comfortably be ignored when $T = 30$ or $T = 40$. The average number of time series observations in the specifications of Tables 2 to 5 ranges from 28 to 36, which suggests that the bias is small. However, to investigate whether the fixed effects estimates are biased due to countries with a small number of time series observations, we reran the specification of Table 5, column (3), without the countries for which we have fewer than 30 observations, thereby making the bias arguably negligible. Dropping these countries reduces the sample size from 3058 to 2441 observations. As in the original results, the coefficient on the lagged dependent variable (Δ GDP per capita $(\log)_{t-1}$) enters with a positive sign and is statistically significant at 1 percent. The size of the coefficient is similar to the size of the original coefficient, although somewhat larger. This could indicate a small downward bias in the original coefficient, consistent with the observation of Bond (2002, 2006), Roodman (2008) and Smith (2006) that, in short panels, the fixed effects estimator is likely to be biased downwards. In addition to the lagged dependent variable, the lagged level of GDP per capita also depends on the lagged idiosyncratic error term and its coefficient may therefore also be biased for countries with small T. However, dropping the countries with fewer than 30 observations only leads to a marginally higher coefficient, while the sign and statistical significance of the coefficient do not change, suggesting that any bias in the original coefficient is small. More importantly, even if there is a small bias in the original coefficient of the lagged dependent variable or the lagged level of per capita GDP in Table 5, column (3), this is likely to have led to an underestimation of the effect of commodity export prices on GDP. When dropping the countries with fewer than 30 time series observations, the long-run coefficients of the non-agricultural export price index and its interaction with good governance are larger than the original coefficients in Table 5, while their signs and levels of statistical significance are unchanged. The short-run coefficients of the non-agricultural export price index are also similar to the original short-run coefficients in Table 5. For robustness, we also restricted the larger samples of Table 2, columns (1) to (4), to countries with 30 or more observations and found the same results.[33]

A second potential source of endogeneity relates to parameter heterogeneity. Although the fixed effects estimator we employ in our analysis controls for all time-invariant unobserved country characteristics, some of these characteristics may have changed over the course of the sample period and not accounting for such changes could have affected our estimates. But in addition to the fixed effects, the other parameters in our model may also have changed over time. To assess the importance of parameter heterogeneity, we split the sample of Table 5, column (3), into one subsample for all years prior to 1983 and one subsample for all years since 1983, this being the year in the middle of the sample period 1963-2003. We then re-estimated the specification for both of these subsamples separately and for each coefficient performed a Wald test of equality across the two subsamples. We found that sixteen out of the twenty coefficients, including all six long-run coefficients, do not significantly differ across the two subsamples. The only coefficients that are significantly different are the short-

[33] Alternative dynamic panel estimators that are consistent irrespective of the length of the time series were developed by Arellano and Bond (1991), Holtz-Eakin et al. (1988), Arellano and Bover (1995), and Blundell and Bond (1998), mainly for small T, large N panels. These estimators could be used by rewriting equation (1) in levels and then applying difference- or system-GMM. However, using these estimators in our analysis is not without problems. First, as the number of instruments used in GMM grows rapidly with T, applying GMM to our large T panel is likely to lead to the "problem of too many instruments" (Bowsher, 2002, Roodman, 2009), which causes very inefficient or severely biased estimates. Secondly, applying difference-GMM to the highly persistent level variables in our panel is likely to lead to "weak instruments" and therefore seriously biased estimates, as the lagged levels of the dependent variable are only weakly (or not at all) correlated with the subsequent first differences (Bond, 2002, 2006). This problem is aggravated by the inclusion of the second lagged dependent variable, $\Delta y_{i,t-2}$, as an exogenous regressor in the difference-GMM, as over time this variable consists of the same information as $y_{i,t-2}$, used to instrument for $\Delta y_{i,t-1}$. Hence, only further-lagged instruments add information and these may be less predictive, which could make the first-stage regression quite weak (Bond, 2006). In light of the problems with using GMM in a large T panel with non-stationary series and given that the fixed effects bias is small or negligible for large T, we prefer to use fixed effects estimation, as suggested by Roodman (2008) for large T panels.

run coefficients of the lagged level of GDP per capita (although still significant at 1 percent in both subsamples), the lagged differences of trade to GDP and inflation, and natural disasters. We also found that the negative long-run coefficient of the non-agricultural export price index remains statistically significant at 5 and 10 percent, despite the much smaller size of both subsamples (1232 and 1826 observations, respectively). These results indicate that the resource curse effect occurred in both the 1963-1982 and 1983-2003 periods and that the short-run and long-run effects of non-agricultural export prices were not significantly different across the pre-1983 and post-1983 periods. We also re-estimated the specification in Table 5, column (3), using the full sample but adding a dummy for the 1983-2003 period and interactions of this dummy with each of the four variables for which we find significantly different coefficients across the subsamples. This allows the coefficients of the four short-run variables to differ across the two periods, while using the full sample in estimation. The results showed that our findings on the short- and long-run effects of non-agricultural export prices, as well as the long-run interaction effect with governance, are robust to the inclusion of these variables. Based on these additional estimations, we conclude that parameter heterogeneity over time is limited and that any heterogeneity that was not accounted for in our original results did not affect our main finding that commodity booms have positive short-term and adverse long-term effects. As a result, pooling the two subsamples, as we did throughout the analysis in this paper, seems appropriate.

A third 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. This problem is not likely to be important in our analysis for two reasons. First, in contrast with the mean-reverting variables considered by Alogoskoufis and Smith (1991), the long run right-hand side variables in our analysis are non-stationary. It is therefore not likely that the expected period *t* level of the variables differs much from the realized level in period *t-1* and, as a result, any bias in the estimated long run coefficients is likely to be small. Secondly, and specifically related to commodity export prices, 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).[34]

Finally, we tested the robustness of our results to a wide range of additional or alternative controls used in the empirical growth literature. We separately added indicators[35] of institutional quality, conflict, commodity price volatility, industrial development, investment, public and private consumption, democracy, the black market premium, the number of assassinations, an alternative measure of trade openness, and exchange rate overvaluation to the specification of Table 5, column (3). The results supported our earlier finding that higher non-agricultural export prices have positive short-term effects on output but negative long-run effects in countries with bad governance.

[34] 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 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.

[35] As the indicators could potentially affect GDP both in the short and long run, we include both their lagged level and contemporaneous first-difference. For institutional quality, we use the PRS Group's composite International Country Risk Guide risk rating (extrapolated using the 1985 rating for all years prior to 1986) and the political constraints indicators "polconiii" and "polconv" from Henisz (2002). For conflict, we use the cumulative number of civil war years. For commodity price volatility, we use a variable that captures the pre-1986 mean absolute change in the composite unweighted commodity export price index for the years before 1986 and the post-1985 mean absolute change in the index for the years after 1985. For industrial development, we use manufacturing as a share of GDP and services as a share of GDP, both from the World Development Indicators (WDI). For investment, we use gross capital formation as a share of GDP (WDI). For public and private consumption, we use general government final consumption expenditure as a share of GDP (WDI) and household final consumption expenditure as a share of GDP (WDI), respectively. For democracy, we use the democracy and autocracy indicators "democ", "autoc" and "polity2" from Polity IV. For the black market premium, we use the parallel market exchange rate premium from the Global Development Network Growth Database. For the number of assassinations, we use the variable "s17f1" from the Banks' Cross-National Time-Series Data Archive. For an alternative measure of trade openness, we use a dummy variable for de jure trade openness from Sachs and Warner (1995b). Finally, for exchange rate overvaluation, we use the logged index of real exchange rate overvaluation from the Global Development Network Growth Database.

6. Conclusions and implications for the recent boom in global commodity prices

We find strong evidence of a conditional resource curse. 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.

Our findings have important implications for non-agricultural commodity exporters with weak institutions, many of which are located in Sub-Saharan Africa. Using the baseline estimation results in Table 2, column (3), we simulated the effects of the post-2000 boom in global commodity prices on the growth rate of commodity-exporting economies. We first extended the commodity price series and the commodity export price index to 2009.[36] Both non-agricultural and agricultural commodities experienced a boom during the decade which was abruptly punctured in 2008 by the onset of global crisis. We then evaluated the effects of the recent boom against a counterfactual of commodity export prices fixed at their 1999 pre-boom levels for all years after 1999. In other words, we compare the model predictions based on *actual* prices to the predictions based on a counterfactual in which we fix prices at their 1999 levels, everything else the same. The short-run results of the simulation reflect the estimated positive short-run effect of commodity prices discussed earlier. In particular, the recent increases in commodity prices are expected to have raised growth in commodity-exporting economies in the same year and the next two years. This effect is strongest in the year after the price increase. We use the earlier example of Nigeria to illustrate the size of the effect. In 2000 the world oil price increased by 57 percent. Since Nigeria's commodity exports are dominated by oil and correspond to 35 percent of its GDP, the oil price increase

[36] The 2009 prices are based on the first two quarters of 2009.

represented a strong rise in Nigeria's export revenues. Using our simulation results, we predict that this windfall added 0.73 percentage points to Nigeria's growth rate in the next year (2001). Compared to Nigeria's actual growth of 0.5 percent in 2001 and on average 0.3 percent over the period 1996-2001, this effect is economically significant. However, despite this positive short run effect, the recent commodity boom is likely to have strongly adverse long term effects in countries with bad governance. In the case of Nigeria, the world oil price in 2009 was still around three times as high as ten years earlier. If this higher price level is permanent, our simulation suggests that the recent boom is predicted to lower Nigeria's long run level of GDP by more than 30 percent. Although this prediction should not literally be used to forecast future output in Nigeria, it does indicate that, if past behaviour is repeated, the recent commodity boom is likely to lead to lower long run output in commodity-exporting countries with bad governance. However, if our tentative diagnosis of the resource curse is correct, then this prognosis could be avoided by improvements in the quality of governance.[37]

[37] The other earlier example of an equally resource-rich country is Zambia, which primarily exports copper. The surge in the world price of copper occurred predominantly in 2004, 2005 and 2006, when prices increased by 61%, 22% and 83%, respectively. Our simulation predicts that the windfall out of this boom added 0.7, 0.5 and 0.4 percentage points to Zambia's growth rate in 2005, 2006 and 2007, respectively. But it also suggests that Zambia's long run level of GDP will be around 30 percent lower as a result of the boom.

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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

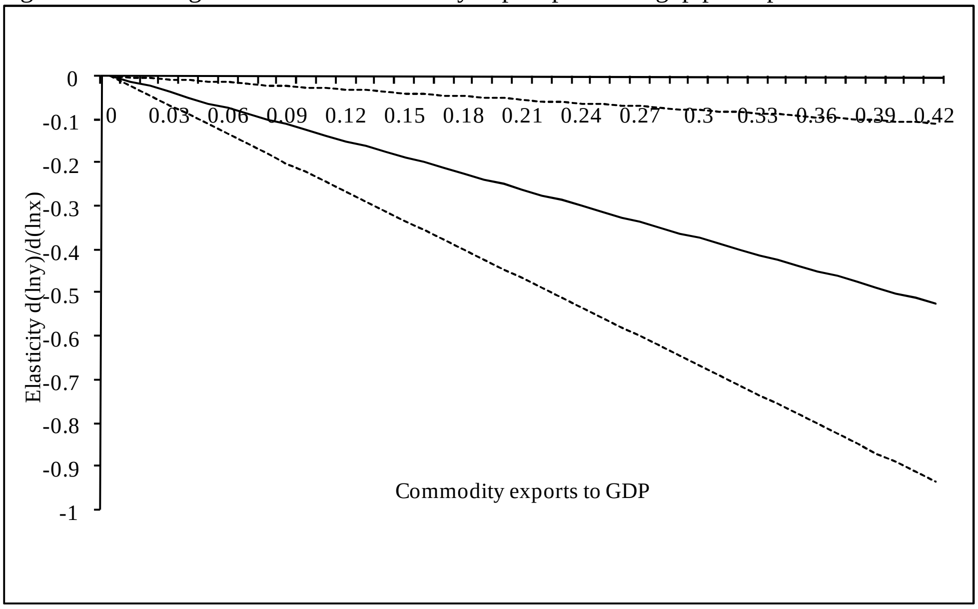
	Obs.	Mean	St. Dev.	Min.	Max.
Real GDP per capita (log)	3579	7.51	1.54	4.31	10.55
Trade to GDP	3579	0.64	0.36	0.06	2.51
Inflation (log (1 + inflation rate))	3579	0.14	0.29	-0.24	5.48
Reserves to GDP	3579	0.09	0.10	0.00	1.24
Commodity export price index	3579	0.33	0.36	0.00	1.85
Unlogged unweighted index (1980 = 100)	3579	83.02	29.29	15.10	235.41
Commodity exports to GDP (net)	3579	0.08	0.09	0.00	0.42
Non-agricultural commodity export price index	3579	0.17	0.33	0.00	1.84
Unlogged unweighted non-agri index (1980 = 100)	3579	85.19	28.58	14.92	260.58
Non-agricultural commodity exports to GDP (net)	3579	0.04	0.08	0.00	0.40
Agricultural commodity export price index	3579	0.16	0.21	0.00	1.03
Unlogged unweighted agri index (1980 = 100)	3579	92.72	28.39	30.45	287.03
Agricultural commodity exports to GDP (net)	3579	0.04	0.05	0.00	0.22
Dummy good governance	3058	0.25	0.43	0	1
Oil import price index	3579	3.11	1.86	0.00	4.96
Δ GDP per capita (log)	3579	0.02	0.05	-0.36	0.30
Δ Trade to GDP	3579	0.00	0.08	-0.88	1.21
Δ Inflation (log (1 + inflation rate))	3579	-0.00	0.19	-3.62	2.52
Δ Reserves to GDP	3579	0.00	0.03	-0.25	0.31
Δ Commodity export price index	3579	0.00	0.02	-0.27	0.41
Δ Unlogged unweighted index (1980 = 100)	3579	-0.55	14.35	-81.33	76.58
Δ Oil import price index	3579	0.02	0.21	-0.68	0.93
Coup	3579	0.03	0.17	0	2
Civil war	3579	0.07	0.26	0	1
Natural disaster	3579	0.26	0.58	0	4

Table 2 Estimation results: baseline specifications

	(1)	(2)	(3)	(4)	(5)
			Estimates of long-run coefficients		
Trade to GDP	0.722***	0.734***	0.475***	0.492***	4.393
	(0.186)	(0.177)	(0.109)	(0.106)	(3.611)
Inflation (log)	-0.206*	-0.198*	-0.186**	-0.188**	-2.299
	(0.113)	(0.106)	(0.074)	(0.074)	(3.024)
Reserves to GDP	0.663**	0.611**	0.648***	0.623***	12.810
	(0.300)	(0.278)	(0.195)	(0.191)	(9.788)
Commodity export price index	-1.778***		-1.243**		
	(0.622)		(0.486)		
Non-agricultural export price index		-2.020***		-1.407***	-6.526
		(0.608)		(0.498)	(4.669)
Agricultural export price index		3.213*		1.004	-4.876
		(1.604)		(1.326)	(4.216)
Oil import price index	-0.171**	-0.192***	-0.134**	-0.153**	-0.064
	(0.075)	(0.070)	(0.065)	(0.063)	(0.260)
			Estimates of short-run coefficients		
GDP per capita (log)$_{t-1}$	-0.045***	-0.047***	-0.062***	-0.063***	-0.002
	(0.006)	(0.007)	(0.008)	(0.008)	(0.001)
Δ GDP per capita (log)$_{t-1}$	0.152***	0.150***	0.134***	0.134***	0.218***
	(0.028)	(0.027)	(0.028)	(0.028)	(0.028)
Δ Trade to GDP $_{t-1}$	0.018	0.016	0.018	0.017	0.033**
	(0.014)	(0.015)	(0.016)	(0.016)	(0.014)
Δ Inflation (log) $_{t-1}$	-0.006	-0.006	-0.003	-0.003	-0.008*
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Δ Reserves to GDP $_{t-1}$	0.092**	0.091**	0.046	0.046	0.083
	(0.045)	(0.045)	(0.047)	(0.048)	(0.052)
Δ Commodity export price index $_t$	0.085*	0.088*	0.038	0.041	0.062
	(0.044)	(0.044)	(0.050)	(0.050)	(0.051)
Δ Commodity export price index $_{t-1}$	0.155***	0.147***	0.206***	0.201***	0.160***
	(0.040)	(0.037)	(0.047)	(0.046)	(0.039)
Δ Commodity export price index $_{t-2}$	0.080	0.074	0.067	0.062	0.020
	(0.110)	(0.113)	(0.107)	(0.109)	(0.117)
Δ Oil import price index $_t$	-0.002	-0.003	-0.005	-0.005	-0.006
	(0.004)	(0.004)	(0.008)	(0.008)	(0.009)
Δ Oil import price index $_{t-1}$	-0.006	-0.005	0.008	0.007	0.002
	(0.004)	(0.004)	(0.006)	(0.006)	(0.006)
Δ Oil import price index $_{t-2}$	-0.005	-0.005	-0.002	-0.003	-0.007
	(0.005)	(0.006)	(0.008)	(0.008)	(0.009)
Coup $_t$	-0.030***	-0.031***	-0.031***	-0.031***	-0.029***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Civil war $_t$	-0.022***	-0.023***	-0.023***	-0.023***	-0.016***
	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)
Natural disaster $_t$	-0.005***	-0.005***	-0.004**	-0.004**	-0.002
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Country fixed effects	YES	YES	YES	YES	NO
Regional time dummies	NO	NO	YES	YES	YES
Time trend	YES	YES	NO	NO	NO
Observations	3579	3579	3579	3579	3579
R-squared (within)	0.14	0.14	0.26	0.26	0.26

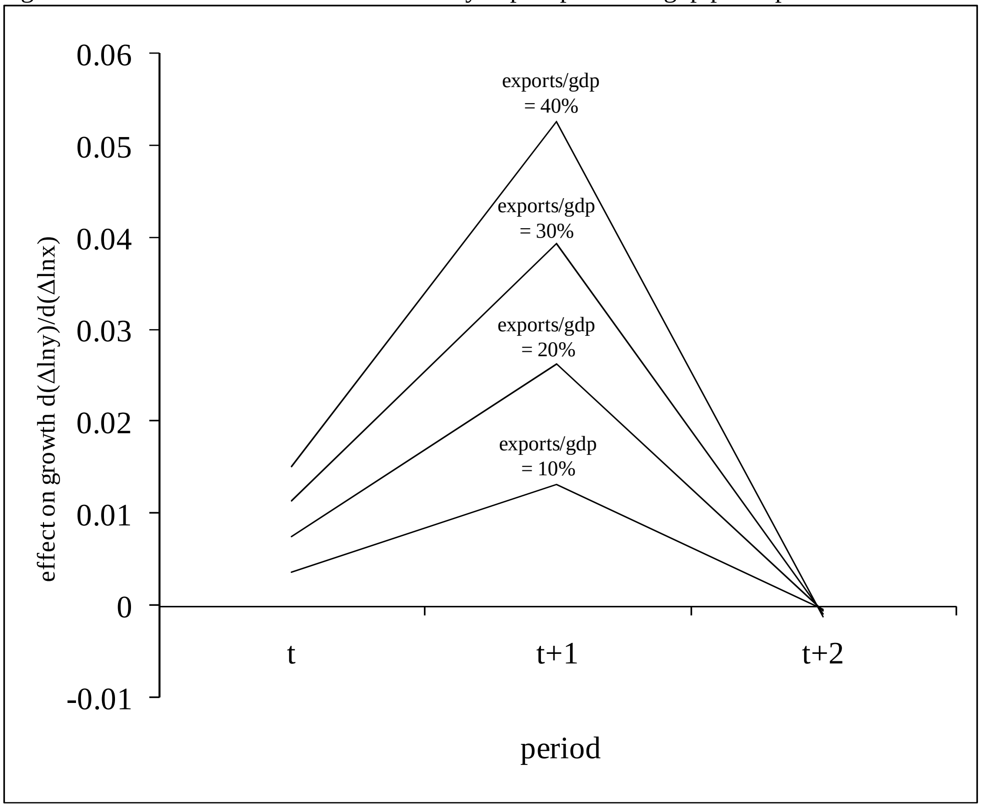
Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by year and are reported in parentheses. Columns (1) to (4) report the R-squared *within*, while column (5) reports the R-squared. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Figure 1a The long-run effect of commodity export prices on gdp per capita



Notes: Figure 1a is based on the estimation results in Table 2, column (3). 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 1b The short-run effect of commodity export prices on gdp per capita



Notes: Figure 1b is based on the estimation results in Table 2, column (3). 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.

Table 3 Estimation results: the resource curse conditional on governance

	(1)	(2)	(3)	(4)
	Estimates of long-run coefficients			
Trade to GDP	0.788***	0.809***	0.478***	0.516***
	(0.202)	(0.189)	(0.120)	(0.118)
Inflation (log)	-0.209*	-0.199*	-0.186**	-0.189**
	(0.119)	(0.110)	(0.075)	(0.075)
Reserves to GDP	0.557	0.507	0.544**	0.511**
	(0.356)	(0.324)	(0.215)	(0.209)
Commodity export price index	-1.911***		-1.275**	
	(0.680)		(0.502)	
Commodity export price index * good governance	1.802		1.778**	
	(1.124)		(0.787)	
Non-agricultural export price index		-2.117***		-1.394***
		(0.644)		(0.511)
Non-agricultural export price index * good governance		2.980***		2.256***
		(1.087)		(0.795)
Agricultural export price index		3.768**		1.313
		(1.860)		(1.691)
Agricultural export price index * good governance		-6.751		-1.498
		(4.277)		(3.098)
Oil import price index	-0.182**	-0.203**	-0.127*	-0.136**
	(0.082)	(0.076)	(0.068)	(0.066)
	Estimates of short-run coefficients			
GDP per capita (log)$_{t-1}$	-0.044***	-0.047***	-0.065***	-0.065***
	(0.006)	(0.007)	(0.009)	(0.009)
Δ GDP per capita (log)$_{t-1}$	0.166***	0.164***	0.153***	0.152***
	(0.032)	(0.031)	(0.034)	(0.033)
Δ Trade to GDP $_{t-1}$	0.002	0.001	0.002	0.001
	(0.018)	(0.018)	(0.020)	(0.020)
Δ Inflation (log) $_{t-1}$	-0.005	-0.005	-0.002	-0.002
	(0.004)	(0.004)	(0.005)	(0.005)
Δ Reserves to GDP $_{t-1}$	0.108*	0.106*	0.055	0.055
	(0.055)	(0.054)	(0.060)	(0.060)
Δ Commodity export price index $_{t}$	0.086*	0.087*	0.024	0.028
	(0.044)	(0.044)	(0.053)	(0.053)
Δ Commodity export price index $_{t-1}$	0.154***	0.148***	0.208***	0.203***
	(0.039)	(0.037)	(0.051)	(0.049)
Δ Commodity export price index $_{t-2}$	0.080	0.074	0.063	0.059
	(0.116)	(0.118)	(0.121)	(0.122)
Δ Oil import price index $_{t}$	0.000	-0.000	-0.003	-0.003
	(0.004)	(0.004)	(0.008)	(0.008)
Δ Oil import price index $_{t-1}$	-0.003	-0.002	0.011	0.010
	(0.004)	(0.004)	(0.007)	(0.007)
Δ Oil import price index $_{t-2}$	-0.005	-0.004	-0.004	-0.004
	(0.005)	(0.006)	(0.009)	(0.009)
Coup $_{t}$	-0.027***	-0.028***	-0.028***	-0.028***
	(0.007)	(0.007)	(0.008)	(0.008)
Civil war $_{t}$	-0.019***	-0.020***	-0.020***	-0.021***
	(0.005)	(0.005)	(0.005)	(0.005)
Natural disaster $_{t}$	-0.005***	-0.005***	-0.004**	-0.004**
	(0.002)	(0.002)	(0.002)	(0.002)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	NO	YES	YES
Time trend	YES	YES	NO	NO
Observations	3058	3058	3058	3058
R-squared within	0.14	0.15	0.28	0.28

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: subsamples good and bad governance

	(1) bad gov.	(2) good gov.	(3) bad gov.	(4) good gov.
	Estimates of long-run coefficients			
Trade to GDP	0.614***	2.190***	0.388***	2.347***
	(0.184)	(0.723)	(0.127)	(0.755)
Inflation (log)	-0.167	-2.096*	-0.174**	-1.957
	(0.101)	(1.094)	(0.069)	(1.196)
Reserves to GDP	1.009**	0.492	0.572*	0.874
	(0.439)	(0.413)	(0.303)	(0.856)
Non-agricultural export price index	-1.811***	1.988	-1.338**	1.896
	(0.634)	(1.346)	(0.504)	(1.610)
Oil import price index	-0.157*	-0.302*	-0.158**	-0.311
	(0.083)	(0.151)	(0.078)	(0.201)
	Estimates of short-run coefficients			
GDP per capita (log)$_{t-1}$	-0.051***	-0.030***	-0.072***	-0.029***
	(0.008)	(0.009)	(0.011)	(0.009)
Δ GDP per capita (log)$_{t-1}$	0.160***	0.235***	0.144***	0.211***
	(0.033)	(0.061)	(0.037)	(0.059)
Δ Trade to GDP $_{t-1}$	0.003	0.023	-0.003	0.075**
	(0.018)	(0.037)	(0.022)	(0.033)
Δ Inflation (log) $_{t-1}$	-0.005	-0.078**	-0.001	-0.055*
	(0.004)	(0.031)	(0.005)	(0.030)
Δ Reserves to GDP $_{t-1}$	0.099	0.084*	0.053	0.121*
	(0.066)	(0.044)	(0.076)	(0.066)
Δ Non-agricultural export price index $_t$	0.070	0.219***	0.006	0.140**
	(0.047)	(0.073)	(0.062)	(0.069)
Δ Non-agricultural export price index $_{t-1}$	0.124***	0.074	0.168***	0.129**
	(0.041)	(0.044)	(0.052)	(0.061)
Δ Non-agricultural export price index $_{t-2}$	0.066	-0.017	0.051	-0.012
	(0.127)	(0.092)	(0.145)	(0.100)
Δ Oil import price index $_t$	-0.000	0.001	-0.008	0.009*
	(0.004)	(0.006)	(0.012)	(0.006)
Δ Oil import price index $_{t-1}$	0.002	-0.015**	0.008	-0.001
	(0.004)	(0.007)	(0.008)	(0.008)
Δ Oil import price index $_{t-2}$	-0.002	-0.007	-0.005	-0.007
	(0.005)	(0.008)	(0.014)	(0.005)
Coup $_t$	-0.028***	-0.058***	-0.029***	-0.050***
	(0.007)	(0.008)	(0.008)	(0.006)
Civil war $_t$	-0.019***	-0.008	-0.022***	0.003
	(0.005)	(0.005)	(0.005)	(0.004)
Natural disaster $_t$	-0.005**	-0.005*	-0.004*	-0.005
	(0.002)	(0.003)	(0.002)	(0.003)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	NO	YES	YES
Time trend	YES	YES	NO	NO
Observations	2302	756	2302	756
R-squared within	0.14	0.36	0.28	0.58

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: instrumental variables estimation

	(1)	(2)	(3)	(4)
	Estimates of long-run coefficients			
Trade to GDP	0.814***	1.057***	0.502***	0.649***
	(0.210)	(0.247)	(0.145)	(0.168)
Inflation (log)	-0.207*	-0.090	-0.187**	-0.137***
	(0.117)	(0.071)	(0.077)	(0.053)
Reserves to GDP	0.556	0.827**	0.529*	0.780***
	(0.391)	(0.417)	(0.276)	(0.260)
Non-agricultural export price index	-2.145***	-2.405***	-1.315***	-1.801***
	(0.348)	(0.518)	(0.343)	(0.636)
Non-agricultural export price index * good governance	3.113***	3.321***	2.213***	2.111***
	(0.529)	(0.554)	(0.533)	(0.581)
Oil import price index	-0.197***	-0.239***	-0.133	-0.229*
	(0.058)	(0.065)	(0.085)	(0.123)
	Estimates of short-run coefficients			
GDP per capita (log)$_{t-1}$	-0.044***	-0.042***	-0.065***	-0.064***
	(0.006)	(0.006)	(0.009)	(0.011)
Δ GDP per capita (log)$_{t-1}$	0.167***	0.125***	0.154***	0.126***
	(0.034)	(0.037)	(0.031)	(0.032)
Δ Trade to GDP $_{t-1}$	0.002	-0.003	0.001	-0.002
	(0.016)	(0.018)	(0.017)	(0.017)
Δ Inflation (log) $_{t-1}$	-0.005	-0.006	-0.002	-0.003
	(0.004)	(0.005)	(0.005)	(0.005)
Δ Reserves to GDP $_{t-1}$	0.110***	0.127**	0.060	0.063
	(0.039)	(0.050)	(0.043)	(0.059)
Δ Non-agricultural export price index $_{t}$	0.071	0.125*	0.027	0.135
	(0.064)	(0.073)	(0.071)	(0.115)
Δ Non-agricultural export price index $_{t-1}$	0.125*	0.189**	0.160**	0.315**
	(0.068)	(0.093)	(0.077)	(0.139)
Δ Non-agricultural export price index $_{t-2}$	0.067	0.115*	0.047	0.262*
	(0.055)	(0.068)	(0.068)	(0.134)
Δ Oil import price index $_{t}$	-0.000	-0.002	-0.003	0.003
	(0.003)	(0.003)	(0.008)	(0.012)
Δ Oil import price index $_{t-1}$	-0.003	-0.003	0.006	0.019
	(0.003)	(0.004)	(0.008)	(0.018)
Δ Oil import price index $_{t-2}$	-0.004	-0.005	-0.005	0.026
	(0.004)	(0.004)	(0.008)	(0.019)
Coup $_{t}$	-0.028***	-0.025***	-0.028***	-0.023***
	(0.009)	(0.009)	(0.007)	(0.006)
Civil war $_{t}$	-0.020***	-0.017***	-0.020***	-0.018***
	(0.005)	(0.006)	(0.006)	(0.006)
Natural disaster $_{t}$	-0.005***	-0.007***	-0.004**	-0.006***
	(0.002)	(0.002)	(0.002)	(0.002)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	NO	YES	YES
Time trend	YES	YES	NO	NO
Method	OLS	2SLS	OLS	2SLS
Observations	3058	2733	3058	2733
R-squared within	0.14	0.15	0.28	0.30

Notes: The dependent variable is the first-differenced log of real GDP per capita. Columns (1) and (3) report OLS results. Columns (2) and (4) report the second-stage results of a two-stages-least-squares procedure in which we instrument for the lagged level, difference, and two lagged differences of the non-agricultural export price index, and for its interaction with the good governance dummy. Robust standard errors (clustered by country) are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Appendix: 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 world commodity price indices from International Financial Statistics (IFS, series 74 for butter and coal, 76 for all others), except for the natural gas and gasoline indices, which are from the 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. The geometrically weighted index was first calculated on a quarterly basis and deflated by the export unit value (IFS, series 74..DZF). We then weighted the log of the annual average (rescaled so that 1980 = 100) index by the share of net commodity exports in GDP (GDP in current US dollars, WDI). The sub-indices for non-agricultural and agricultural commodities were constructed in the same way.[38] The oil import price index was constructed by interacting the log of the annual average deflated oil price index 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). Since these are expressed in

[38] 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.

thousands of barrels per day, we multiply by 365 times the 2001 mean weekly world oil price per barrel (EIA).

Trade openness exports plus imports of goods and services as a share of GDP (WDI).

Inflation log (1 + (annual % change in consumer prices/100)), data from WDI.

International reserves over GDP IFS (1..SZF and AA.ZF) and WDI.

Civil war 1 for civil war, 0 otherwise (Gleditsch, 2004).

Coup d'etat number of extraconstitutional or forced changes in the top government elite and/or its effective control of the nation's power structure (Banks' Cross-National Time-Series Data Archive). Unsuccessful coups are not counted.

Natural disasters nr. of large disasters (≥ 0.5% of pop. affected, or damage ≥ 0.5% of GDP, or ≥ 1 death per 10000, criteria established by the IMF). From WHO CRED. Geological disasters: earthquakes, landslides, volcano eruptions, tidal waves; Climatic disasters: floods, droughts, extreme temperatures, wind storms; Human disasters: famines, epidemics.