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

Collier, Paul and Goderis, Benedikt

University of Oxford

15 June 2008

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

Paul Collier† and Benedikt Goderis‡
Department of Economics, University of Oxford

Abstract

Currently, evidence on the 'resource curse' yields a conundrum. While there is much cross-section evidence to support the curse hypothesis, 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 resource curse. Commodity booms have positive short-term effects on output, but adverse long-term effects. The long-term effects are confined to "high-rent", non-agricultural commodities. We also find that the resource curse is avoided by countries with sufficiently good institutions. We test the channels of the resource curse proposed in the literature and find that it is explained by real exchange rate appreciation and public and private consumption. Our findings have important implications for non-agricultural commodity exporters with weak institutions, especially in light of the current unprecedented boom in global commodity prices.

Keywords: commodity prices; natural resource curse; growth

JEL classification: O13, O47, Q33

---

* We thank Chris Adam, Kofi Adjepong-Boateng, Robin Burgess, Phillip Crowson, Mardi Dungey, 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, and the University of Gothenburg for useful comments. We acknowledge support of the UK Economic and Social Research Council and the UK Department for International Development.

† Centre for the Study of African Economies, Department of Economics, University of Oxford, Manor Road, Oxford OX1 3UQ, UK. Email: Paul.Collier@economics.ox.ac.uk, URL: http://users.ox.ac.uk/~econpco/.

‡ Centre for the Study of African Economies, Department of Economics, University of Oxford, Manor Road, Oxford OX1 3UQ, UK. Email: Benedikt.Goderis@economics.ox.ac.uk, URL: http://www.csae.ox.ac.uk/members/biogs/goderis/goderis.html.

1. Introduction

A large literature suggests that there is a 'resource curse': natural resource abundant countries tend to grow slower than resource scarce countries.[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: commodity booms significantly raise growth. The current African growth acceleration coincident with the commodity boom that began in 2000 is clearly consistent with these findings.

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 control variables.[5] This methodology does not consider commodity prices and is unable to disentangle the dynamics of the resource curse. It is therefore not well-suited for testing the wide range of proposed channels in the theoretical resource curse literature. Further, cross-sectional growth regressions suffer from potential omitted variable bias and it is therefore "crucial to move from cross-country to panel data evidence" (Van der Ploeg, 2007). However, the approach pioneered by Deaton and Miller (1995), namely vector autoregressive (VAR) models, cannot address long-run effects. It is therefore possible that the positive short-run effects are offset by a subsequent resource curse beyond the horizon of the VAR models: the post-2000 upturn would be a false dawn. In this paper we adopt panel cointegration methodology to analyze global data for 1963 to 2003 to disentangle the short and long

[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 (2007) provides a survey of the resource curse literature. Alexeev and Conrad (forthcoming) and Brunnschweiler and Bulte (2008) argue that contrary to the claims made in the literature, natural resources positively affect growth.

[5] Manzano and Rigobon (2006) use panels with two or four time series observations and find that the resource curse effect disappears once one allows for fixed effects.

run effects of commodity prices on growth. Panel data allow for the inclusion of country-specific fixed effects, which effectively control for all unobservable time-invariant country characteristics. In addition, the use of panel data allows for a much larger sample size, as it exploits the within-country variation in the regression variables. We also include regional time dummies, further reducing concerns of omitted variable bias, and we allow the effects of commodity prices to vary across different types of commodities. We investigate all the transmission channels of the resource curse proposed in the literature in a systematic manner and we address potential sources of endogeneity that have sometimes been neglected in previous literature.

We find strong evidence in support of the resource curse hypothesis. In particular, commodity booms have positive short-term effects on output, but adverse long-term effects. The long-term effects are confined to "high-rent", non-agricultural commodities. Within this group, we find that the resource curse is avoided by countries with sufficiently good institutions. When testing the importance of the transmission channels, we find that real exchange rate appreciation, public and private consumption, and to a lesser extent external debt, manufacturing, and services, explain the curse.

Our findings have important implications for non-agricultural commodity exporters with weak institutions, many of which are located in Sub-Saharan Africa. They point at the post-2000 boom in global commodity prices as an important determinant of the recent growth acceleration in Africa's commodity exporting economies. But they also suggest that the commodity boom is, if past behaviour is repeated, likely to have strongly adverse long-term effects, making the recent growth acceleration particularly misleading. However, if our tentative diagnosis of the root cause of the resource curse as being due to errors in governance is correct, then this prognosis could be avoided by improvements in the quality of governance.

The rest of this paper is structured as follows. Section 2 describes the empirical analysis. Section 3 reports the estimation results and simulates the short and long run effects of higher commodity export prices on growth. Section 4 investigates whether the resource curse occurs conditional on governance. Section 5 deals with the endogeneity of resource dependence and governance. Section 6 tests the importance of the proposed transmission channels. Section 7 concludes.

2. The Empirical Analysis

In this section we describe our econometric model and the variables used in estimation. Data description and sources can be found in Appendix A. Panel unit root and panel cointegration tests are discussed in Appendix B. The short-run and long-run effects of commodity export prices on GDP per capita are analyzed using the following error-correction model:

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

for $i = 1,\ldots N$ and $t = 1,\ldots T$, where $y_{i,t}$ is log real GDP per capita in country i in year t, α_i is a country-specific fixed effect, and $z_{i,t}$ is an $rT \times 1$ vector of regional time dummies, where r is the number of regions.[6] $x_{i,t-1}$ is an $m \times 1$ vector of m variables that are expected to affect GDP both in the short run and long run.

[6] The country-specific fixed effect captures all the time-invariant characteristics of the individual countries, which eliminates the possibility of omitted variable bias due to time-invariant unobserved variables. The vector of regional time dummies captures 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. 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.

We include a constructed commodity export price index to test the effect of commodity export prices. To investigate whether the effects vary across different types of commodities, we also experiment with sub-indices 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, and 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 the wide range of additional or alternative controls used in the literature, including indicators of institutional quality, exchange rate overvaluation, external debt, income inequality, commodity price volatility, industrial development, public, private, and total investment, public and private consumption, democracy, capital account openness, the black market premium, the number of assassinations, and an alternative measure of trade openness. These variables are not included in our preferred specification because they were either not robustly significant or severely lowered the number of observations in our sample.[7] Finally, $s_{i,t}$ is an $n \times 1$ vector of n control variables that are expected to have only a short-run effect on growth and 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 1a reports summary statistics for the variables used in estimation.

[7] We include most of them in section 6, when we investigate the transmission channels of the resource curse. The results for the other variables are available upon request. 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 country-specific fixed effect.

2.1 Constructing commodity price indices

The commodity export price index was constructed using the methodology of Deaton and Miller (1995) and Dehn (2000). We collected data on world commodity prices and commodity export values for as many commodities as data availability allowed. Table 1b lists the 50 commodities in our sample. For each country, we calculate the total 1990 value of commodity exports. We construct weights by dividing the individual 1990 export values for each commodity by this total. These 1990 weights are then held fixed over time and applied to the world price indices of the same commodities to form a country-specific geometrically weighted index of commodity export prices. To allow the effect of commodity export prices to be larger for countries with larger exports, we weight the log of the deflated index by the share of commodity exports in GDP. The separate indices for non-agricultural and agricultural commodities were constructed in the same way. The oil import price 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.

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

Table 2 reports the results of estimating equation (1).[8] 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 countries. We next investigate whether this adverse long-run effect is common to all the commodities in our index. We decompose the general commodity export price index into two sub-indices: one for non-agricultural

[8] The long-run coefficients correspond to $-(1/\lambda)\bullet\beta_1$ in equation (1). The short-run coefficients correspond to $\lambda, \beta_2, \beta_3,$ and β_4 in equation (1).

commodities only and one for agricultural commodities only. Table 2, column (2), shows the results when we replace the general 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 statistically significant at 1 percent.[9] By contrast, the coefficient for agricultural commodity export prices is positive and insignificant. 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 adding the regional time dummies to the specification of column (1). The coefficient of the commodity export price index again enters negative and is statistically significant at 1 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. An example of a highly commodity-dependent country is Zambia. In 1990 Zambia's commodity exports represented 35 percent of its GDP. The results in Figure 1a therefore predict a long-run elasticity of -0.44.[10] In other words, a 10 percent increase in the price of Zambian commodity exports leads to a 4.4 percent lower long-run level of GDP per capita. These results clearly suggest the existence of a long-run resource curse. 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 example of Zambia

[9] 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.

[10] Recall that the commodity export price index is weighted by the share of commodity exports in GDP. So for Zambia, the long-run elasticity equals the long-run coefficient, -1.243, multiplied by Zambia's share of commodity exports in GDP, 0.35.

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

When replacing the general index by the sub-indices in column (4), the results are also similar to before. The coefficient of the non-agricultural commodity export price index enters negative and is again significant at 1 percent. The effect is substantial. For a country like Nigeria, which in 1990 had non-agricultural exports that represented 35 percent of its GDP (almost exclusively oil), 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 Nigeria's GDP per capita. The coefficient of the agricultural commodity export price index enters negative but is 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 three 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 enters negative and is significant at 5 percent, suggesting that higher inflation leads to a lower long-run GDP level. The oil import price index, which was included to control for the effect of oil prices on oil importing countries, enters with the expected negative sign but is not statistically significant.

The short-run GDP determinants also enter with the expected sign. 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 has a positive short-run effect on GDP growth. Thus, the short-run dynamics of a commodity boom are quite contrary to the long-run effects. Figure 1b illustrates the short-run effect by showing the impulse response functions of an increase in the growth rate of commodity export prices for different levels of commodity exports to GDP. 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 short-run effect of commodity export prices is consistent with the findings in Deaton and Miller (1995) and Raddatz (2007).[11] 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.

Table 2, column (3), also reports the coefficients of the other short-run GDP determinants. The coefficient of lagged GDP per capita is negative and significant at 1 percent. The size of the coefficient suggests that the speed of adjustment to long-run equilibrium is 6.2 percent per year. The first lag of the dependent variable enters positive and is also 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.[12] 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 on growth in the first subsequent year,

---

[11] 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.

[12] We do not include the contemporaneous changes in order to limit concerns of endogeneity.

although these effects are not significant.[13] Next, the two political shocks, 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 2.2 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.

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 these 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.

[13] Even though the changes in the oil import price index are not significant, we include them because the commodity export price index also enters with up to two lags.

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.[14] The ICRG is a commercial rating service whose continued viability has been dependent upon client firms regarding it as having value. There is therefore some reasonable presumption that it has informational content. The first group, which for convenience we will call the "good governance" group, consists of the countries with a mean ICRG score of 75 or higher. This group contains countries like Australia, Canada, and Norway, but also Botswana. The second "bad governance" group consists of the countries with a mean ICRG score below 75 and contains for example Venezuela, Libya and 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. The results are reported in Table 3.[15] 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 general commodity export price index into sub-indices for non-agricultural and agricultural commodities. As

---

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

[15] We restrict the sample to countries for which the mean ICRG score is available. As a result, the number of observations drops from 3608 to 3087.

previously, the direct effect of the non-agricultural export price index enters negative and is 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 statistically 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 coefficients, which is positive and significant at 5 percent. This suggests that far from suffering from a resource curse, countries with good governance succeed in transforming 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 insignificant, while its interaction with good governance enters negative but is also insignificant. This indicates that the effects of higher agricultural export prices in countries with good and bad governance are not significantly different. It also supports our earlier finding that higher agricultural export prices do not lead to any long-run resource curse effect.

Table 3, columns (3) and (4), report the results when adding the regional time dummies to the specifications of columns (1) and (2). The results are very similar. In column (3), the general commodity export price index again enters negative and is significant at 1 percent, while its interaction with good governance is again positive but is now significant at 1 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 strongly support the findings in columns (1) and (2) and clearly show that the resource curse occurs in badly governed countries but not in countries with good governance.

The agricultural commodity export price index enters positive but is insignificant, while its interaction enters negative and is also insignificant, as in column (2).

We next investigate the robustness of these results by rerunning the specifications in Table 3 using the initial 1985 composite ICRG scores rather than the average scores.[16] The results are very similar. In particular, the results for the composite index and the two sub-indices are robust to using this alternative measure of governance.

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 excluding and including regional time dummies, respectively. In both cases the non-agricultural index enters with a negative sign and is significant at 1 percent. 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. In the specification of column (2) this effect is statistically significant at 5 percent. Not only is the resource curse effect absent in countries with good governance, the long-run effect of higher export prices is now positive, as one would expect. The effect is also economically significant. 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 0.23. In other words, a 10 percent increase in the price of non-agricultural commodities leads to a 2.3 percent higher long-run level of Norway's GDP per capita.[17] These results provide

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

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

strong evidence that the resource curse occurs conditional on bad governance. Countries with sufficiently good governance do not suffer from the curse, and instead 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 indices are endogenous, i.e. correlated with the error term in equation (1). 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 constant over time, supply responses to price changes are excluded from the analysis. Nonetheless, countries that are major exporters of one or more commodities may have an influence on the world price of those commodities, which could lead to biased estimates. To address this concern, we express each country's exports of a given commodity as a share of the total world exports of that commodity and repeat this for all other commodities in our sample. This yields a list of commodity export shares that reflect the importance of individual exporters in the global markets for individual commodities. We found that of the 129 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 in our sample affected our results by re-estimating the specifications in Tables 2 and 3 but without these 22 countries. The results, available upon request from the authors, show that our findings are strongly robust to the exclusion of major exporters of individual commodities. In particular, the long-run coefficients for the commodity export price index and the non-agricultural commodity export price index and their interactions with good governance in the specifications of

Tables 2 and 3 are very similar to the original coefficients and are always significant at 1 percent. The short-run positive effects of commodity prices are strongly robust as well. Hence, our results do not seem to be biased by countries that are major exporters of one or more commodities and that may influence world prices of these commodities.[18]

In addition to world commodity prices, the ratio of commodity exports over GDP is also potentially endogenous. As explained in section 2.1, we weight the commodity price indices by this ratio, which could lead to omitted variable or reverse causality bias.[19] Consider two resource-rich countries: one which has suffered from bad policies, slow growth, and a lack of industrialization, and one which has benefited from good policies, fast growth, and industrialization. The bad policy country will have a higher commodity exports to GDP ratio due to the lack of development of its non-resource sectors. This implies that the estimated effect of higher commodity prices on growth in our estimations could be (partly) due to the higher weights we attach to countries with a poor growth record.

To address this concern we need to instrument for the ratio of non-agricultural commodity exports to GDP, these being the commodities that appear to generate the resource curse. As an instrument, we use the 2000 value of sub-soil assets (minerals) in current US dollars per capita developed by the World Bank (2006).[20] 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. The ratio of non-agricultural commodity exports over GDP does

[18] We repeated this robustness check using a threshold of 10 percent instead of 20 percent. 34 out of the 129 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 generally robust to the exclusion of these 34 countries. The only result that did not survive was the interaction effect of the commodity export price index and the non-agricultural commodity export price index with good governance in Table 3. This was due to the fact that only 10 of the 22 good governance countries remained in the sample.
[19] Any time invariant omitted variables are captured by the fixed effects in our estimations.
[20] These estimates were earlier used by Brunnschweiler and Bulte (2008) to proxy resource abundance.

not enter the specifications by itself but only as a weight of the non-agricultural export price index. We therefore construct an instrument for the index by repeating the procedure in section 2.1 but instead of weighting the (unweighted) non-agricultural index by the ratio of non-agricultural commodity exports over GDP (assumed to be endogenous), we now weight it by the 2000 value of sub-soil assets in current US dollars per capita. For this instrument to be valid, it should be correlated with the ratio of non-agricultural exports over GDP, and it should not be correlated with the error term. The former is likely to hold, as commodity exports (net of imports) are only possible if those commodities are available in a country. The latter requires that the instrument does not itself affect growth, other than through its effect on the endogenous regressor (exclusion restriction), does not depend on growth, and is not correlated with omitted growth determinants. The former is likely to be fulfilled as it is hard to see how a country's resource abundance could affect its exposure to commodity export prices, other than through its relationship with the level of commodity exports. The latter two requirements are less likely to be fulfilled. Slow-growing countries are less likely to invest in geological exploration and are more likely to overexploit the discovered stock of resources. As a result, their stock of discovered resources in the ground may be lower than in fast-growing countries, everything else equal. This means that weighting the non-agricultural export price index by the value of sub-soil assets per capita may imply giving higher weights to fast-growing countries. Although this could potentially bias the results, the direction of the bias is opposite to the bias in the uninstrumented regressions, where higher weights were given to slow-growing countries. Comparing the coefficients of the instrumented and uninstrumented regressions can therefore shed light on the size of the potential bias and the numerical range within which the actual coefficient is likely to be located.

In addition to the export price indices, the dummy for good governance also potentially suffers from endogeneity, which could lead to a biased coefficient of the interaction term. 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 22 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 construct an instrument for the interaction term of the index with the good governance dummy by running a probit regression of the governance dummy on the three variables from Hall and Jones (1999) for the sample in Table 3 and collecting the fitted values.[21] We interact the fitted values of the probit regression with the instrument for the non-agricultural commodity export price index discussed above.[22] This yields an additional instrument for the interaction term of the non-agricultural commodity export price index and the good governance dummy.

We next use our constructed instruments to perform a two-stage-least-squares estimation procedure. For comparison, Table 5, columns (1) and (3), first report the OLS estimation results when replacing the commodity export price index in Table 3, columns (1) and (3), by the non-agricultural commodity export price index. The short and long run effects of non-agricultural commodity prices are consistent with the results for the composite index in Table 3. Table 5, columns (2) and (4), report the

[21] All three variables enter with the expected positive signs and are significant at 1 percent. The pseudo R-squared is 0.69.
[22] Goderis and Ioannidou (2008) perform a similar procedure to construct instruments, following Wooldridge (2002), p. 237.

two-stage-least-squares estimates, in which we instrument for the level and differences of the non-agricultural index (using the level and differences of the first constructed instrument), and for the interaction of the index with good governance (using the second constructed instrument).[23]

The non-agricultural commodity export price index enters with a negative sign and is significant at 1 and 5 percent in columns (2) and (4), respectively. The size of the coefficients is very similar to the size of the coefficients in columns (1) and (3), indicating that if there is an endogeneity bias, it is likely to be small. In fact, we performed Davidson-MacKinnon tests of exogeneity and could not reject the null hypothesis of consistent OLS estimates for the non-agricultural commodity export price index in columns (2) and (4) with p-values of 0.48 and 0.33, 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 also similar to the coefficients in columns (1) and (3), although no longer significant. Again, Davidson-MacKinnon tests did not reject the null of exogeneity of the interaction terms with p-values of 0.44 and 0.46, respectively. The short-run coefficients of the non-agricultural index in Table 5, columns (2) and (4), enter with positive signs and gain in both size and significance compared to the OLS estimates in columns (1) and (3). We performed Davidson-MacKinnon tests for all three short-run coefficients in columns (2) and (4) and could not reject the null of consistent OLS estimates for 5 out of the 6 coefficients, while rejecting exogeneity at 10 percent for the second lag of the differenced non-agricultural index in column (4).[24] This evidence suggests that any bias is likely to be small and if anything leads to a small

[23] To save space, we do not report the results of the first stage. However, in all first-stage regressions, the relevant instrument enters with the expected sign and is statistically significant at 1 percent. The first-stage results are available upon request.

[24] The p-values were 0.87, 0.47, 0.18, and 0.81, 0.15, 0.08 for the short-run coefficients in columns (2) and (4), respectively.

underestimation of the positive short-run growth effect of higher non-agricultural commodity 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 in Table 5, column (3), to test the channels of the resource curse.

6. The channels of the resource curse

The literature offers seven candidate explanations for the resource curse effect: Dutch disease, governance, conflict, excessive borrowing, inequality, volatility, and lack of education. Since the responses appropriate for overcoming the resource curse differ radically as between these routes, their relative magnitude is evidently of importance. In this section we test for the importance of these explanations.

We first explore the possibility that the long-run negative effect reflects the occurrence of Dutch Disease effects. An increase in commodity prices appreciates the real exchange rate, lowering the competitiveness of the non-resource exports sector, and potentially harming long-run output if there are positive externalities to production in this sector (Corden and Neary, 1982; Van Wijnbergen, 1984; Sachs and Warner, 1995a, 1999; Torvik, 2001). This argument is related to recent literature that shows how specialization in natural resources can divert economies away from manufacturing or other skill-intensive activities, thereby slowing down learning-by-doing and reducing incentives for people to educate themselves (Michaels, 2006). To test for the importance of this channel, we add a real exchange rate indicator[25] to the specification in Table 5, column (3). As an appreciation of the real exchange rate

[25] The best available indicator for the real appreciation of a country's currency is probably a real *effective* exchange rate measure. Such measures are available but their coverage for the countries and years in our sample is limited. We therefore use the real exchange rate vis-à-vis the US dollar by adjusting the nominal exchange rate for relative consumer price levels (International Financial Statistics lines rf and 64). However, we experiment with a real effective exchange rate in section 6.1 when we discuss the routes through which governance drives the resource curse.

could potentially affect GDP both in the short and in the long run, we include both the level and the first difference of the index. Further, to allow for the possibility that the effect of a real appreciation is different for resource-abundant countries, we also include interaction terms of the level and differenced exchange rate indicator with the share of non-agricultural exports in GDP. If the negative long-run effect of non-agricultural commodity export prices works (partly) *through* their impact on the real exchange rate, then the estimated direct effect of the export price indices should become smaller once we control for exchange rate appreciation. The results are reported in the two columns in the top left corner of Table 6. In the first column, the level of the real exchange rate enters positive, suggesting that, consistent with Dutch disease, a more appreciated exchange rate (lower level of the indicator) is associated with lower long-run levels of GDP. The interaction of the index with the share of exports in GDP also enters positive, suggesting that this relationship is stronger in resource-abundant countries. However, both coefficients are not statistically significant and should therefore be viewed with caution. The differenced exchange rate variables are also insignificant. Adding the real exchange rate scarcely changes the coefficient of the non-agricultural export price index, as can be seen from the results in the second column for the same sample without the real exchange rate. The long-run coefficient changes from -1.29 to -1.25, which suggests that Dutch Disease does not explain the long-run resource curse effect. Although countries with good governance do not suffer from a resource curse, their long-term gain from higher commodity prices might be negatively affected by Dutch Disease.[26] This gain is captured by the linear combination of the coefficients of the non-agricultural index and its interaction with good governance. This combination changes from 1.00 for the

[26] The term "Dutch Disease" originated in the Netherlands, a "good governance" country with the highest mean composite ICRG rating after Switzerland and Norway. During the 1960s, the high revenue generated by its natural gas discovery led to a sharp decline in the competitiveness of its other, non-booming tradable sector.

specification without the real exchange rate to 1.05 for the specification with the real exchange rate, indicating that Dutch Disease is not important in understanding the effect of higher commodity prices on GDP in good governance countries.

We next explore whether the resource curse induces weak governance. The literature has proposed several such routes. Resource rents may invite non-productive lobbying and rent seeking, as in Tornell and Lane (1999), Baland and Francois (2000), Torvik (2002), and Wick and Bulte (2006). Mehlum et al. (2006) argue that this problem only occurs in countries with grabber-friendly institutions, while countries with producer-friendly institutions do not suffer from a curse. A related literature emphasizes the role of government in the misallocation of resource revenues. Robinson et al. (2006) argue that resource booms have adverse effects because they provide incentives for politicians to engage in inefficient redistribution in return for political support. Again, existing institutions are crucial, as they determine the extent to which politicians can respond to these perverse incentives. The inefficient redistribution can take various forms such as public employment provision (Robinson et al., 2006), subsidies to farmers, labor market regulation, and protection of domestic industries from international competition (Acemoglu and Robinson, 2001). We investigate governance using the same approach as for Dutch disease. There is no agreed composite measure of the quality of governance and so we have investigated a range of commonly used proxies: the Composite International Country Risk Guide (ICRG) risk rating (PRS Group), the parallel market exchange rate premium (Global Development Network Growth Database), civil liberties and political rights (Freedom House), political constraints (Henisz, 2002), democracy, autocracy, and a combined measure of democracy and autocracy (Polity IV), and checks and balances (Database of Political Institutions 2004). To save space, the third and fourth column of the top left corner of Table 6 ("Governance") only report the results for the composite ICRG risk rating[27] since adding any of these other indicators scarcely changes the long run results. The effect of the ICRG rating is positive and significant at 1 percent, both in the short and in the long run, indicating that good governance countries grow faster. For resource-abundant countries, the long-run effect is bigger, although the difference is not significant. While these results indicate that the quality of governance is an important GDP determinant, it only leads to a marginally smaller resource curse effect. The long-run coefficient of commodity prices changes from -1.43 to -1.39, suggesting that the deterioration of governance is not the central explanation of the resource curse. So even though the resource curse only occurs in countries with weak governance, it is not explained by a deterioration of governance in those countries.

We next turn more briefly to five other proposed channels for the resource curse. First, resource abundance can increase the incidence of violence (Collier and Hoeffler, 2004). This can occur through a weakening of the state, easy finance for rebels and warlords (Skaperdas, 2002), or quasi-criminal activities and gang rivalries (Mehlum et al., 2006; Hodler, 2006). Secondly, resource abundance can tempt a government into excessive external borrowing, as in Mansoorian (1991) and Manzano and Rigobon (2006). Thirdly, resource abundance exposes countries to commodity price volatility which could discourage investment (Sala-i-Martin and Subramanian, 2003). Fourthly, resource abundance can lead to increased inequality, which can harm growth (Sokoloff and Engerman, 2000). And finally, as suggested by Gylfason (2001), resource abundance can lower incentives for citizens or the government to invest in education, which can also lower growth. We investigate the importance of

[27] The ICRG rating is only available since 1984, but the coverage is better for 1985. We therefore use the 1985 ratings for all years in our sample prior to 1986, which means we do not capture changes in governance for these years. However, the results for governance are robust to using alternative indicators that are available for all years in our sample.

these channels through the same approach.[28] Controlling for these possible channels does not lead to smaller coefficients for our export price index, suggesting that individually these channels do not explain our resource curse finding.[29]

6.1 Testing the routes through which governance drives the resource curse

Even though the resource curse does not work *through* governance, we have found strong evidence that it works *conditional* on governance. The recent theoretical literature proposes two explanations, each of which implies additional channels of the resource curse. Mehlum et al. (2006) argue that resource rents invite non-productive lobbying and rent seeking, and that the pay-off from these activities is high in countries with grabber-friendly institutions but low in countries with producer-friendly institutions. This leads entrepreneurs in countries with bad institutions away from productive activities into non-productive rent-seeking activities, which in the long run slows down industrial development. We empirically test this theory by adding measures of the importance of the manufacturing and services sectors to our specification.[30] The results are reported in the top right and bottom left corners of Table 6. Controlling for manufacturing and services only leads to marginally smaller coefficients for our export price index. We can therefore conclude that the resource curse does not seem to work through a slower speed of industrial development or lower growth in the services sector.

[28] We use the following indicators for conflict, excessive borrowing, inequality, and volatility, respectively: the cumulative number of civil war years; total external debt to GNP (World Bank's Global Development Finance); gross household income inequality (gini), from the University of Texas Inequality Project (EHII2.3); a variable that captures the pre-1986 mean absolute change in the general unweighted commodity export price index for the years before 1986 and the post-1985 mean absolute change in the general unweighted commodity export price index for the years after 1985. For education we use three variables: the average years of primary, the average years of secondary, and the average years of higher schooling of the population aged 15 and over (Barro and Lee, 2000). Since these variables are only available for 1960, 1965, 1970, 1975, 1980, 1985, 1990, 1995, and 1999, we fill in the missing years by linear interpolation.

[29] The coefficient of the interaction term of the index with good governance is insignificant for the specifications under "Excessive borrowing" in Table 6. This is due to a very low availability of the external debt variable for good governance countries. The same holds for the specifications under "Manufacturing", discussed in section 6.1. To save space, Table 6 does not report the results for education.

[30] Manufacturing and services as a share of GDP were both taken from the World Development Indicators.

The other recently proposed explanation points at inefficient redistribution by the government. Robinson et al. (2006) argue that permanent commodity booms increase incentives for politicians to stay in power. In countries where government accountability is lacking, politicians will use the resource windfall revenues to bias the outcome of elections or in non-democratic regimes political contests. This bias can be induced in many ways but Robinson et al. refer to informal literature that “points to the centrality of public sector employment as a tool for influencing people’s voting behaviour”. Hence, resource rents create inefficiencies by facilitating public employment provision by politicians in return for political support, but only in countries with weak institutions. In countries with strong institutions, the extent to which politicians can use public money to bias elections is limited and therefore the resource curse does not occur. In addition to public employment provision, inefficient redistribution can take place through protection of domestic industries from international competition, subsidies to farmers, and labor market regulation (Acemoglu and Robinson, 2001). We empirically test this theory by adding measures of public consumption and de jure and de facto trade openness to our specification.[31] The results for openness are reported in the third to sixth column of the bottom part of Table 6, while the results for public consumption are reported in the next two columns. Controlling for either de jure or de facto openness hardly changes the long-run effect of the index, but the negative effect of public consumption seems to explain part of the resource curse. The absolute size of the coefficient falls by 22 percent. These results support the argument of Robinson et al. (2006) that commodity booms

[31] For public consumption we use general government final consumption expenditure as a share of GDP, which includes civil servant salaries (World Development Indicators). For de jure trade openness we use a dummy variable that takes a value of 1 if a country has open trade, constructed by Sachs and Warner (1995b). For de facto openness we use the (logged) index of real exchange rate overvaluation from the Global Development Network Growth Database. The index captures the extent to which the real exchange rate is distorted away from a hypothetical free-trade level.

lead to inefficient public sector employment provision which then slows down economic development.

In both explanations of why governance is important, booms lead workers or entrepreneurs away from productive activities into less productive rent-seeking or public sector activities. With this shift, one might expect to see a shift in the pattern of a country's aggregate expenditures as well. As more people secure their income through rent-seeking or public employment and the government allocates more of its revenue to public employment provision, aggregate investment levels will fall and public and private consumption will increase. Gylfason and Zoega (2006) show empirically that natural resource dependence slows down growth through lower levels of investment.

In addition to lowering investment, booms may also lead to a lower quality of investment projects. Robinson and Torvik (2005) develop a theory in which "white elephants", investment projects with negative social surplus, are used as a means of inefficient redistribution aimed at influencing the outcomes of elections. This suggests yet another channel of the curse: inefficient redistribution through inefficient investment projects. We test the importance of these shifts in expenditure by adding measures of private consumption and total investment to our specification.[32] The results are reported in the ninth to twelfth column of the bottom part of Table 6. Controlling for private consumption and total investment takes away part of the long-run resource curse effect. The absolute size of the coefficient falls by 17 percent and 14 percent, respectively. Also note that the positive long-run effect of investment is significantly smaller in resource-rich economies. This is consistent with the theory of

[32] The World Development Indicators database does distinguish between public and private consumption but does not distinguish between public and private investment. As a measure of private consumption we use household final consumption expenditure as a share of GDP. For total investment we use gross capital formation as a share of GDP.

inefficient redistribution through inefficient investment projects in Robinson and Torvik (2005).

In addition to these political economy explanations of why the resource curse works conditional on poor governance, Matsen and Torvik (2005) show that resource-rich countries with high savings rates generally have managed to escape the curse, while countries with low savings rates have not. This suggests that the curse may work through Dutch disease, but only in countries that spend a relatively large part of the windfall. Since bad governance countries tend to be big spenders, this could explain our findings. To test the importance of this channel, we once more add a real exchange rate indicator to our baseline specification. The best available indicators for the real appreciation of a country's currency are probably real *effective* exchange rate measures but their data coverage is limited. Since we now want to investigate the importance of Dutch disease in bad governance countries, we restrict our sample to such countries and add the real effective exchange rate measure with the largest coverage, the (logged) IMF's real effective exchange rate index based on relative consumer prices (IFS line "rec"), to our baseline specification. The results are reported in the last two columns of the bottom part of Table 6. The long-run coefficient of the real effective exchange rate is positive and significant at 5 percent. This suggests that real exchange rate appreciations are associated with higher long-run growth rates. This does probably not imply that causality runs from exchange rate appreciation to GDP, but rather in the opposite way or through an omitted variable. The long-run coefficient is significantly different for resource-abundant countries, as the interaction term enters negative and is significant at 5 percent. This indicates that for countries with non-agricultural exports to GDP levels above 7 percent, the long-run effect of a real appreciation on GDP is negative, consistent with Dutch disease. The short-run effects are not significant.

The low coverage of the real exchange rate indicator lowers the sample size to a level where, in the absence of the exchange rate indicator, the resource curse effect is just insignificant (p-value is 0.101). However, the size of the coefficient (-1.86) is comparable to the coefficients in the larger samples. If anything, the resource curse effect is bigger in the smaller sample than in the larger ones. This gives us some confidence that the smaller sample can be used to test the importance of Dutch disease. Adding the real effective exchange rate index changes the coefficient of the non-agricultural export price index from -1.86 to -0.52, suggesting that Dutch disease does indeed explain an important part of the long-run resource curse effect.[33]

We next investigate whether combinations of indicators can explain the resource curse effect in our estimations. We start with the variables that are individually important in explaining the curse: the real effective exchange rate, public consumption, private consumption, and total investment. We include these variables in our baseline specification and again test their importance by eliminating each of them individually and observing the change in the long-run coefficient, while keeping the sample constant. We find that, once we control for the other variables, the real effective exchange rate, public consumption, and private consumption remain important, while total investment does not. Hence, we rerun the specifications with the real effective exchange rate, public consumption, and private consumption only. The results are reported in the first two columns of Table 7. The three variables together fully explain the resource curse effect, which is now significant again. The

[33] We also ran fixed effects panel regressions of the change in the logged real effective exchange rate index on the change in the non-agricultural export price index. The contemporaneous as well as the first three lags of the change in the export price index entered positive and statistically significant, supporting the Dutch disease hypothesis that commodity booms lead to an appreciation of the real exchange rate.

long-run coefficient changes from -1.96 to +0.78. Hence, once we control for the real effective exchange rate, public consumption, and private consumption as channels of the resource curse, higher commodity prices no longer have a negative long-run effect. If anything, they are a blessing rather than a curse, just as in good governance countries.

As a robustness check we again considered all of the other channels by adding each of the indicators individually to the specification in Table 7, column (1), and observing the long-run coefficient, while again keeping the sample constant. The results support our earlier finding that these other channels cannot account for the curse: controlling for them does not lead to smaller coefficients for our export price index. The only three exceptions are external debt, manufacturing, and services. If we control for each of these three additional variables, the size of the positive long-run effect of commodity prices increases. Table 7, columns (3) and (4), report the results when we add these three variables to the specifications of columns (1) and (2). Controlling for the total of six variables changes the long-run resource curse coefficient from -2.60 to +1.10. We conclude that the resource curse can be explained by real exchange rate appreciation, public and private consumption, and to a lesser extent external debt, manufacturing, and services.

7. Conclusions

We find strong evidence of a resource curse. Commodity booms have positive short-term effects on output, but adverse long-term effects. The long-term effects are confined to “high-rent”, non-agricultural commodities. Within this group, we find that the resource curse is avoided by countries with sufficiently good institutions. We investigate possible transmission channels and find that real exchange rate appreciation, public and private consumption, and to a lesser extent external debt, manufacturing, and services, explain the resource curse. This lends support to the large literature that stresses the importance of Dutch disease in resource-rich economies. It also supports recent theory that points at inefficient redistribution in return for political support as the root of the curse.

Our findings have important implications for non-agricultural commodity exporters with weak institutions, many of which are located in Sub-Saharan Africa. Using our estimation results, we simulated the effects of the post-2000 boom in global commodity prices on the growth rate of Africa's commodity exporting economies. We find that half of the current growth of these economies is attributable to short-term effects of the commodity boom, leaving a residual of underlying growth that remains low. Moreover, the current boom is, if past behaviour is repeated, likely to have strongly adverse long term effects, so that the recent acceleration in growth rates is particularly misleading. However, if our tentative diagnosis of the root cause of the resource curse as being due to errors in governance is correct, then this prognosis could be avoided by improvements in the quality of governance.

References

Acemoglu, Daron and James A. Robinson, "Inefficient Redistribution," *American Political Science Review* 95 (2001), 649-661.

Acemoglu, Daron, James A. Robinson and Simon Johnson, "The Colonial Origins of Comparative Development: An Empirical Investigation," *American Economic Review* 91 (2001), 1369-1401.

Alexeev, Michael and Robert Conrad, "The Elusive Curse of Oil," *Review of Economics and Statistics* (forthcoming).

Baland, Jean-Marie and Patrick Francois, "Rent-Seeking and Resource Booms," *Journal of Development Economics* 61 (2000), 527-542.

Barro, Robert J., and Jong-Wha Lee, "International Data on Educational Attainment: Updates and Implications," CID Working Paper no. 42 (2000).

Brunnschweiler, Christa N. and Erwin H. Bulte, "The Natural Resource Curse Revisited and Revised: A Tale of Paradoxes and Red Herrings," *Journal of Environmental Economics and Management* 55 (2008), 248-264.

Collier, Paul, "On the Economic Consequences of Civil War," *Oxford Economic Papers* 51 (1999), 168-183.

Collier, Paul and Anke Hoeffler, "Greed and Grievance in Civil Wars," *Oxford Economic Papers* 56 (2004), 563-595.

Corden, W. Max, and J. Peter Neary, "Booming Sector and De-Industrialization in a Small Open Economy," *The Economic Journal* 92 (1982), 825-848.

Deaton, Angus S. and Ronald I. Miller, "International Commodity Prices, Macroeconomic Performance, and Politics in Sub-Saharan Africa", *Princeton Studies in International Finance* 79 (1995).

Dehn, Jan, "Commodity Price Uncertainty in Developing Countries," CSAE Working Paper no. 2000-12 (2000).

Gleditsch, Kristian S., "A Revised List of Wars Between and Within Independent States, 1816-2002," *International Interactions* 30 (2004), 231-262.

Goderis, Benedikt and Vasso P. Ioannidou, "Do High Interest Rates Defend Currencies During Speculative Attacks? New Evidence," *Journal of International Economics* 74 (2008), 158–169.

Gylfason, Thorvaldur, "Natural Resources, Education, and Economic Development," *European Economic Review* 45 (2001), 847-859.

Gylfason, Thorvaldur, Tryggvi Thor Herbertsson and Gylfi Zoega, "A Mixed Blessing: Natural Resources and Economic Growth," *Macroeconomic Dynamics* 3 (1999), 204-225.

Gylfason, Thorvaldur and Gylfi Zoega, "Natural Resources and Economic Growth: The Role of Investment," *The World Economy* 29 (2006), 1091-1115.

Hall, Robert E. and Charles I. Jones, "Why Do Some Countries Produce So Much More Output per Worker than Others?," *Quarterly Journal of Economics* 114 (1999), 83-116.

Henisz, Witold J., "The Institutional Environment for Infrastructure Investment," Industrial and Corporate Change 11 (2002), 355-389.

Hodler, Roland, "The Curse of Natural Resources in Fractionalized Countries," *European Economic Review* 50 (2006), 1367-1386.

Im, Kyung S., M. Hashem Pesaran and Yongcheol Shin, "Testing for Unit Roots in Heterogeneous Panels," *Journal of Econometrics* 115 (2003), 53-74.

Maddala, G.S. and Shaowen Wu, "A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test," *Oxford Bulletin of Economics and Statistics* 61 (1999), 631-652.

Mansoorian, Arman, "Resource Discoveries and 'Excessive' External Borrowing," *The Economic Journal* 101 (1991), 1497-1509.

Manzano, Osmel and Roberto Rigobon, "Resource Curse or Debt Overhang?," in Daniel Lederman and William F. Maloney (Eds.), *Natural Resources, Neither Curse nor Destiny* (Stanford: Stanford University Press and World Bank, 2006).

Matsen, Egil and Ragnar Torvik, "Optimal Dutch Disease," *Journal of Development Economics* 78 (2005), 494-515.

Mehlum, Halvor, Karl Moene and Ragnar Torvik, "Institutions and the Resource Curse," *The Economic Journal* 116 (2006), 1-20.

Michaels, Guy, "The Long-Term Consequences of Regional Specialization," CEPR Discussion Paper no. 6028 (2006).

Pedroni, Peter, "Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors," *Oxford Bulletin of Economics and Statistics* 61 (1999), 653-670.

Raddatz, Claudio, "Are External Shocks Responsible for the Instability of Output in Low-Income Countries?," *Journal of Development Economics* 84 (2007), 155-187.

Robinson, James A., and Ragnar Torvik, "White Elephants," *Journal of Public Economics* 89 (2005), 197-210.

Robinson, James A., Ragnar Torvik and Thierry Verdier, "The Political Foundations of the Resource Curse," *Journal of Development Economics* 79 (2006), 447-468.

Sachs, Jeffrey D. and Andrew M. Warner, "Natural Resource Abundance and Economic Growth," NBER Working Paper no. 5398 (1995a, revised 1997, 1999).

Sachs, Jeffrey D. and Andrew M. Warner, "Economic Reform and the Process of Global Integration," *Brookings Papers on Economic Activity* 1 (1995b), 1-118.

Sachs, Jeffrey D. and Andrew M. Warner, "The Big Push, Natural Resource Booms and Growth," *Journal of Development Economics* 59 (1999), 43-76.

Sachs, Jeffrey D. and Andrew M. Warner, "The Curse of Natural Resources," *European Economic Review* 45 (2001), 827-838.

Sala-i-Martin, Xavier and Arvind Subramanian, "Addressing the Natural Resource Curse: An Illustration from Nigeria," NBER Working Paper no. 9804 (2003).

Skaperdas, Stergios, "Warlord competition," *Journal of Peace Research* 39 (2002), 435-446.

Sokoloff, Kenneth L. and Stanley L. Engerman, "Institutions, Factor Endowments, and Paths of Development in the New World", *Journal of Economic Perspectives* 14 (2000), 217-232.

Tornell, Aaron and Philip R. Lane, "The Voracity Effect," *American Economic Review* 89 (1999), 22-46.

Torvik, Ragnar, "Learning By Doing and the Dutch Disease," *European Economic Review* 45 (2001), 285-306.

Torvik, Ragnar, "Natural Resources, Rent Seeking and Welfare," *Journal of Development Economics* 67 (2002), 455-470.

Van der Ploeg, Frederick, "Challenges and Opportunities for Resource Rich Economies," *Journal of Economic Literature* (forthcoming, 2007).

Van Wijnbergen, Sweder J. G., "The Dutch Disease: a Disease after All?," *Economic Journal* 94 (1984), 41-55.

Wick, Katharina and Erwin H. Bulte, "Contesting Resources – Rent Seeking, Conflict and the Natural Resource Curse," *Public Choice* 128 (2006), 457-476.

Wooldridge, Jeffrey, *Econometric Analysis of Cross Section and Panel Data*, MIT Press, Cambridge, MA, 2002.

World Bank, *Where is the Wealth of Nations? Measuring Capital for the 21st Century*, International Bank for Reconstruction and Development/World Bank, Washington DC, 2006.

Table 1a Summary statistics

	Obs.	Mean	St. Dev.	Min.	Max.
Real GDP per capita (log)	3608	7.54	1.55	4.31	10.55
Trade to GDP	3608	0.64	0.36	0.06	2.51
Inflation (log (1 + inflation rate))	3608	0.14	0.29	-0.24	5.48
Reserves to GDP	3608	0.09	0.10	0.00	1.24
Commodity export price index	3608	0.34	0.36	0.00	1.97
Unlogged unweighted index (1980 = 100)	3608	81.07	26.87	15.10	230.05
Commodity exports to GDP	3608	0.08	0.09	0.00	0.45
Non-agricultural commodity export price index	3608	0.18	0.33	0.00	1.84
Unlogged unweighted non-agri index (1980 = 100)	3608	83.01	27.04	14.92	260.58
Non-agricultural commodity exports to GDP	3608	0.04	0.08	0.00	0.40
Agricultural commodity export price index	3608	0.16	0.21	0.00	1.11
Unlogged unweighted agri index (1980 = 100)	3608	92.02	28.54	30.57	337.32
Agricultural commodity exports to GDP	3608	0.04	0.05	0.00	0.24
Dummy good governance	3087	0.26	0.44	0	1
Oil import price index	3608	3.11	1.85	0.00	4.96
Δ GDP per capita (log)	3608	0.02	0.05	-0.36	0.30
Δ Trade to GDP	3608	0.01	0.08	-0.88	1.21
Δ Inflation (log (1 + inflation rate))	3608	-0.00	0.19	-3.62	2.52
Δ Reserves to GDP	3608	0.00	0.03	-0.25	0.31
Δ Commodity export price index	3608	0.00	0.02	-0.27	0.41
Δ Unlogged unweighted index (1980 = 100)	3608	-0.51	13.24	-77.88	73.09
Δ Oil import price index	3608	0.02	0.21	-0.68	0.93
Coup	3608	0.03	0.17	0	2
Civil war	3608	0.07	0.26	0	1
Natural disaster	3608	0.25	0.58	0	4

Table 1b Commodities

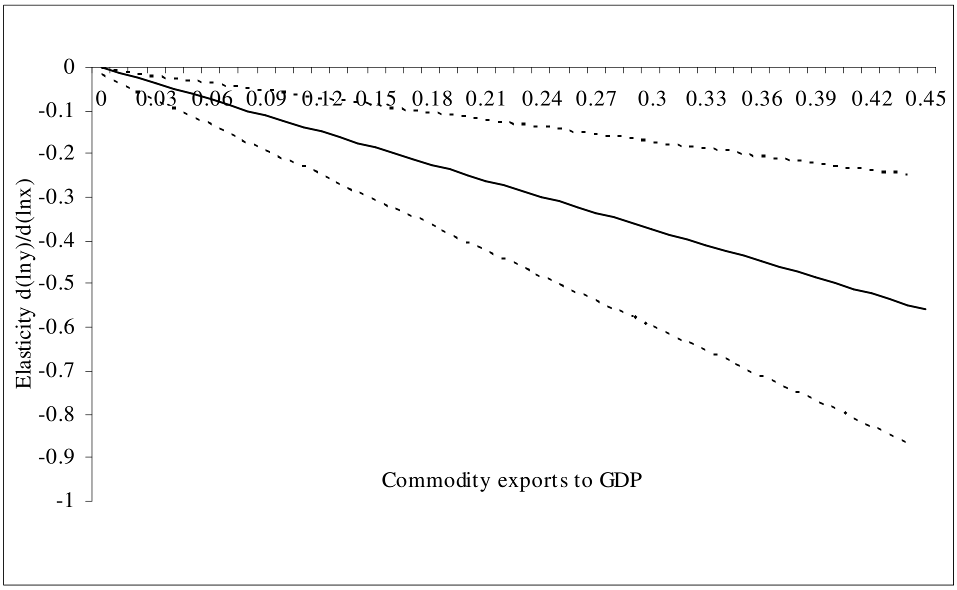
Non-agricultural				
Aluminum	Gasoline	Natural gas	Phosphatrock	Uranium
Coal	Ironore	Nickel	Silver	Urea
Copper	Lead	Oil	Tin	Zinc
Agricultural				
Bananas	Cotton	Oliveoil	Pulp	Sugar
Barley	Fish	Oranges	Rice	Sunfloweroil
Butter	Groundnutoil	Palmkerneloil	Rubber	Swinemeat
Cocoabeans	Groundnuts	Palmoil	Sisal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

Table 2 Estimation results: baseline specifications

	(1)	(2)	(3)	(4)
	Estimates of long-run coefficients			
Trade to GDP	0.851***	0.927***	0.466***	0.483***
	(0.177)	(0.188)	(0.131)	(0.128)
Inflation (log)	-0.215*	-0.216*	-0.185**	-0.188**
	(0.121)	(0.121)	(0.077)	(0.077)
Reserves to GDP	0.869**	0.928**	0.665***	0.642**
	(0.375)	(0.388)	(0.253)	(0.253)
Commodity export price index	-1.947***		-1.243***	
	(0.416)		(0.346)	
Non-agricultural export price index		-2.214***		-1.395***
		(0.387)		(0.352)
Agricultural export price index		1.589		0.920
		(1.971)		(1.198)
Oil import price index	-0.164***	-0.178***	-0.139	-0.157*
	(0.057)	(0.060)	(0.086)	(0.089)
	Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.040***	-0.040***	-0.062***	-0.063***
	(0.005)	(0.005)	(0.008)	(0.008)
Δ GDP per capita $(\log)_{t-1}$	0.150***	0.147***	0.136***	0.135***
	(0.031)	(0.031)	(0.029)	(0.029)
Δ Trade to GDP $_{t-1}$	0.018	0.018	0.019	0.019
	(0.013)	(0.013)	(0.014)	(0.014)
Δ Inflation (log) $_{t-1}$	-0.006	-0.007	-0.003	-0.003
	(0.004)	(0.004)	(0.004)	(0.004)
Δ Reserves to GDP $_{t-1}$	0.091***	0.090***	0.045	0.045
	(0.032)	(0.032)	(0.035)	(0.035)
Δ Commodity export price index $_{t}$	0.085	0.085	0.043	0.046
	(0.059)	(0.058)	(0.063)	(0.063)
Δ Commodity export price index $_{t-1}$	0.146**	0.141**	0.202***	0.197***
	(0.061)	(0.061)	(0.070)	(0.069)
Δ Commodity export price index $_{t-2}$	0.073	0.068	0.066	0.061
	(0.050)	(0.050)	(0.059)	(0.059)
Δ Oil import price index $_{t}$	-0.002	-0.002	-0.004	-0.004
	(0.004)	(0.004)	(0.008)	(0.008)
Δ Oil import price index $_{t-1}$	-0.007**	-0.007**	0.008	0.008
	(0.003)	(0.003)	(0.008)	(0.008)
Δ Oil import price index $_{t-2}$	-0.006	-0.006	-0.002	-0.003
	(0.004)	(0.004)	(0.007)	(0.007)
Coup $_{t}$	-0.031***	-0.031***	-0.031***	-0.031***
	(0.008)	(0.008)	(0.007)	(0.007)
Civil war $_{t}$	-0.022***	-0.022***	-0.022***	-0.023***
	(0.005)	(0.005)	(0.005)	(0.005)
Natural disaster $_{t}$	-0.004***	-0.004**	-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
Observations	3608	3608	3608	3608
R-squared within	0.14	0.14	0.26	0.26

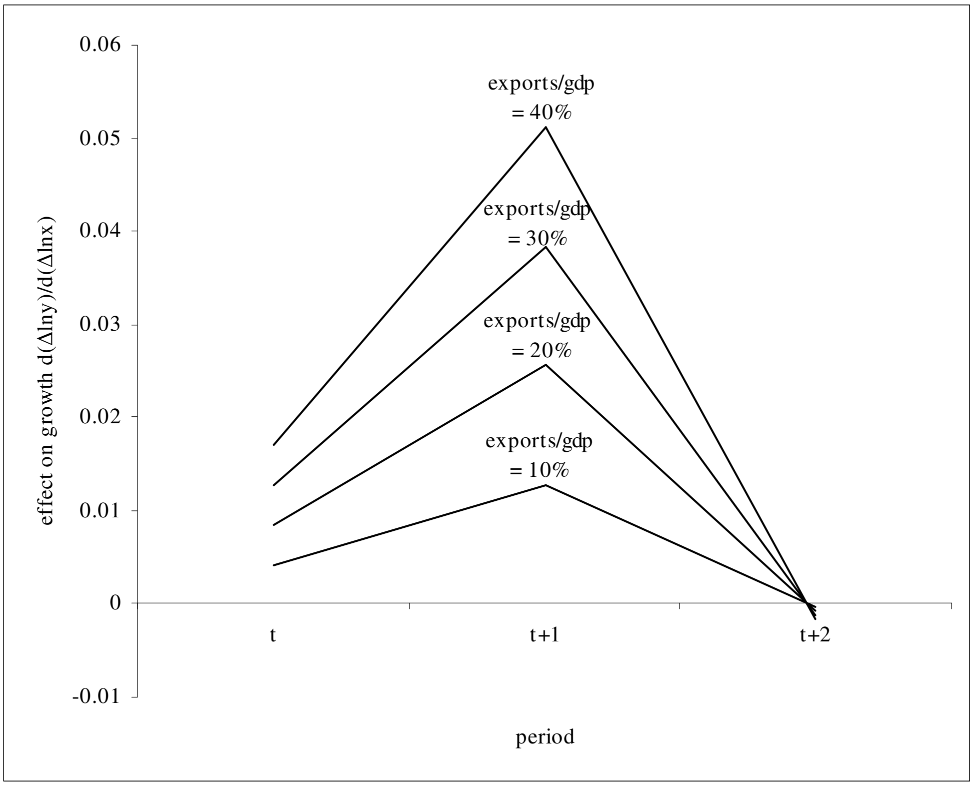
Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by country and are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

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	0.922***	1.002***	0.469***	0.502***
	(0.198)	(0.206)	(0.142)	(0.145)
Inflation (log)	-0.215*	-0.210*	-0.185**	-0.188**
	(0.123)	(0.121)	(0.076)	(0.077)
Reserves to GDP	0.767*	0.818*	0.571**	0.548*
	(0.403)	(0.424)	(0.277)	(0.282)
Commodity export price index	-2.051***		-1.261***	
	(0.426)		(0.342)	
Commodity export price index * good governance	1.757		1.689***	
	(1.510)		(0.635)	
Non-agricultural export price index		-2.261***		-1.369***
		(0.398)		(0.351)
Non-agricultural export price index * good governance		3.467***		2.124***
		(0.586)		(0.603)
Agricultural export price index		1.988		1.130
		(2.102)		(1.225)
Agricultural export price index * good governance		-9.003		-0.132
		(6.273)		(3.475)
Oil import price index	-0.170***	-0.177***	-0.127	-0.136
	(0.065)	(0.065)	(0.084)	(0.087)
	Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.039***	-0.040***	-0.065***	-0.065***
	(0.005)	(0.005)	(0.009)	(0.009)
Δ GDP per capita $(\log)_{t-1}$	0.165***	0.162***	0.155***	0.154***
	(0.034)	(0.034)	(0.031)	(0.031)
Δ Trade to GDP $_{t-1}$	0.004	0.003	0.004	0.003
	(0.016)	(0.016)	(0.017)	(0.017)
Δ Inflation $(\log)_{t-1}$	-0.006	-0.006	-0.002	-0.002
	(0.004)	(0.004)	(0.005)	(0.005)
Δ Reserves to GDP $_{t-1}$	0.107***	0.105***	0.054	0.054
	(0.039)	(0.039)	(0.043)	(0.043)
Δ Commodity export price index $_t$	0.086	0.085	0.029	0.032
	(0.058)	(0.058)	(0.063)	(0.064)
Δ Commodity export price index $_{t-1}$	0.145**	0.139**	0.203***	0.199***
	(0.062)	(0.062)	(0.072)	(0.072)
Δ Commodity export price index $_{t-2}$	0.073	0.068	0.062	0.058
	(0.051)	(0.051)	(0.062)	(0.062)
Δ Oil import price index $_t$	0.000	0.000	-0.003	-0.002
	(0.003)	(0.003)	(0.007)	(0.007)
Δ Oil import price index $_{t-1}$	-0.004	-0.004	0.011	0.010
	(0.003)	(0.003)	(0.008)	(0.008)
Δ Oil import price index $_{t-2}$	-0.006	-0.006*	-0.003	-0.004
	(0.004)	(0.004)	(0.008)	(0.008)
Coup $_t$	-0.028***	-0.028***	-0.028***	-0.028***
	(0.009)	(0.009)	(0.007)	(0.007)
Civil war $_t$	-0.019***	-0.019***	-0.020***	-0.021***
	(0.005)	(0.005)	(0.006)	(0.006)
Natural disaster $_t$	-0.004**	-0.004**	-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
Observations	3087	3087	3087	3087
R-squared within	0.14	0.15	0.28	0.29

Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by country 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)	(2)	(3)	(4)
	bad gov.	good gov.	bad gov.	good gov.
	Estimates of long-run coefficients			
Trade to GDP	0.680***	1.724***	0.390***	2.018**
	(0.159)	(0.435)	(0.128)	(0.843)
Inflation (log)	-0.164	-2.023***	-0.173**	-2.102*
	(0.109)	(0.648)	(0.075)	(1.067)
Reserves to GDP	1.225**	0.343	0.585	0.976
	(0.537)	(0.221)	(0.446)	(0.856)
Non-agricultural export price index	-1.779***	1.663**	-1.314***	1.465
	(0.351)	(0.603)	(0.386)	(1.038)
Oil import price index	-0.133**	-0.296**	-0.158	-0.344**
	(0.061)	(0.111)	(0.112)	(0.154)
	Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.048***	-0.031***	-0.072***	-0.028***
	(0.008)	(0.005)	(0.011)	(0.006)
Δ GDP per capita $(\log)_{t-1}$	0.161***	0.246***	0.147***	0.220***
	(0.036)	(0.032)	(0.034)	(0.030)
Δ Trade to GDP $_{t-1}$	0.005	0.013	-0.001	0.065**
	(0.018)	(0.024)	(0.019)	(0.024)
Δ Inflation (log) $_{t-1}$	-0.006	-0.077	-0.002	-0.053
	(0.004)	(0.053)	(0.005)	(0.044)
Δ Reserves to GDP $_{t-1}$	0.097**	0.092*	0.052	0.121**
	(0.047)	(0.053)	(0.052)	(0.050)
Δ Non-agricultural export price index $_{t}$	0.071	0.199**	0.009	0.133
	(0.066)	(0.087)	(0.077)	(0.082)
Δ Non-agricultural export price index $_{t-1}$	0.115*	0.061	0.165**	0.135**
	(0.069)	(0.054)	(0.080)	(0.055)
Δ Non-agricultural export price index $_{t-2}$	0.059	-0.006	0.050	0.002
	(0.056)	(0.047)	(0.072)	(0.053)
Δ Oil import price index $_{t}$	-0.000	0.001	-0.008	0.010
	(0.004)	(0.003)	(0.011)	(0.008)
Δ Oil import price index $_{t-1}$	0.001	-0.013***	0.008	0.001
	(0.004)	(0.004)	(0.011)	(0.007)
Δ Oil import price index $_{t-2}$	-0.004	-0.006	-0.005	-0.006
	(0.005)	(0.004)	(0.011)	(0.005)
Coup $_{t}$	-0.029***	-0.059***	-0.029***	-0.051***
	(0.009)	(0.004)	(0.007)	(0.006)
Civil war $_{t}$	-0.019***	-0.009***	-0.022***	0.002
	(0.006)	(0.002)	(0.006)	(0.005)
Natural disaster $_{t}$	-0.005**	-0.004	-0.004*	-0.004
	(0.002)	(0.003)	(0.002)	(0.003)
Country fixed effects	YES	YES	YES	YES
Regional time dummies	NO	NO	YES	YES
Observations	2290	797	2290	797
R-squared within	0.14	0.34	0.28	0.57

Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by country 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	0.978***	1.131***	0.492***	0.634***
	(0.198)	(0.153)	(0.142)	(0.112)
Inflation (log)	-0.215*	-0.093	-0.186**	-0.133**
	(0.121)	(0.088)	(0.076)	(0.057)
Reserves to GDP	0.806**	0.966***	0.557**	0.711***
	(0.405)	(0.350)	(0.276)	(0.234)
Non-agricultural export price index	-2.236***	-2.013***	-1.294***	-1.420**
	(0.406)	(0.629)	(0.349)	(0.635)
Non-agricultural export price index * good governance	3.365***	3.077	2.117***	2.017
	(0.664)	(2.128)	(0.606)	(1.359)
Oil import price index	-0.182***	-0.218***	-0.132	-0.183*
	(0.065)	(0.064)	(0.087)	(0.105)
	Estimates of short-run coefficients			
GDP per capita $(\log)_{t-1}$	-0.039***	-0.041***	-0.065***	-0.066***
	(0.005)	(0.004)	(0.009)	(0.006)
Δ GDP per capita $(\log)_{t-1}$	0.165***	0.141***	0.157***	0.140***
	(0.033)	(0.020)	(0.031)	(0.021)
Δ Trade to GDP $_{t-1}$	0.003	-0.003	0.003	-0.001
	(0.016)	(0.013)	(0.018)	(0.014)
Δ Inflation (log) $_{t-1}$	-0.006	-0.006	-0.002	-0.003
	(0.004)	(0.005)	(0.005)	(0.005)
Δ Reserves to GDP $_{t-1}$	0.109***	0.131***	0.059	0.068**
	(0.039)	(0.031)	(0.043)	(0.033)
Δ Non-agricultural export price index $_{t}$	0.073	0.130**	0.030	0.143
	(0.063)	(0.051)	(0.071)	(0.103)
Δ Non-agricultural export price index $_{t-1}$	0.115*	0.170***	0.158**	0.301***
	(0.068)	(0.052)	(0.076)	(0.102)
Δ Non-agricultural export price index $_{t-2}$	0.059	0.091*	0.046	0.249**
	(0.054)	(0.050)	(0.067)	(0.100)
Δ Oil import price index $_{t}$	0.000	-0.002	-0.003	0.004
	(0.003)	(0.004)	(0.008)	(0.014)
Δ Oil import price index $_{t-1}$	-0.004	-0.005	0.006	0.017
	(0.003)	(0.004)	(0.008)	(0.014)
Δ Oil import price index $_{t-2}$	-0.006	-0.007*	-0.005	0.025*
	(0.004)	(0.004)	(0.008)	(0.014)
Coup $_{t}$	-0.028***	-0.024***	-0.028***	-0.022***
	(0.009)	(0.005)	(0.007)	(0.005)
Civil war $_{t}$	-0.019***	-0.015***	-0.020***	-0.017***
	(0.005)	(0.004)	(0.006)	(0.004)
Natural disaster $_{t}$	-0.004**	-0.006***	-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
Method	OLS	2SLS	OLS	2SLS
Observations	3087	2634	3087	2634
R-squared within	0.14	0.16	0.28	0.30

Notes: The dependent variable is the first-differenced log of real GDP per capita. Columns (1) and (3) report OLS estimation 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. Standard errors are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 6 Testing the channels of the resource curse

	Dutch disease		Governance		Conflict		Excessive borrowing		Inequality		Volatility		Manufacturing	
Indicator	0.08		1.23***		0.01		-0.18**		-0.72		0.93		1.34	
	(0.07)		(0.44)		(0.02)		(0.08)		(1.01)		(1.43)		(0.95)	
Indicator	0.64		3.48		0.02		0.46		5.22		-0.99		-7.58*	
* Non-agri exports/GDP	(0.41)		(2.57)		(0.17)		(0.32)		(9.03)		(6.28)		(3.98)	
Δ Indicator	-0.01		0.27***		-0.02***		-0.07***		-0.12		-0.27**		0.07	
	(0.01)		(0.06)		(0.01)		(0.02)		(0.09)		(0.13)		(0.15)	
Δ Indicator	0.08		-0.00		-0.06		0.17*		1.97**		0.37		-0.46	
* Non-agri exports/GDP	(0.07)		(0.35)		(0.04)		(0.08)		(0.87)		(0.87)		(1.45)	
Non-agri export price index	-1.25***	-1.29***	-1.39***	-1.43***	-1.36***	-1.29***	-1.66***	-1.46***	-1.14**	-0.99*	-1.30***	-1.29***	-1.41***	-1.45***
	(0.38)	(0.35)	(0.34)	(0.34)	(0.37)	(0.35)	(0.47)	(0.40)	(0.50)	(0.52)	(0.36)	(0.35)	(0.36)	(0.37)
Non-agri export price index	2.30***	2.29***	2.12***	2.11***	2.14***	2.12***	-0.72	1.14	1.64***	1.69***	2.14***	2.12***	0.65	0.81
* good governance	(0.62)	(0.61)	(0.62)	(0.60)	(0.63)	(0.61)	(4.72)	(4.36)	(0.60)	(0.61)	(0.61)	(0.61)	(0.61)	(0.62)
Observations	2915	2915	2857	2857	3087	3087	1872	1872	1752	1752	3087	3087	2396	2396
R-squared within	0.29	0.28	0.32	0.29	0.28	0.28	0.32	0.29	0.37	0.37	0.28	0.28	0.31	0.30

	Services		De jure openness		De facto openness		Public consumption		Private consumption		Total investment		Dutch disease (2)	
Indicator	0.05		0.00		-0.10		-0.60		-1.14***		2.60***		0.10**	
	(0.50)		(0.08)		(0.07)		(0.74)		(0.40)		(0.58)		(0.05)	
Indicator	2.55		0.95		-0.46		-5.08		6.23***		-7.54***		-1.44**	
* Non-agri exports/GDP	(2.60)		(0.70)		(0.43)		(6.75)		(2.17)		(2.57)		(0.70)	
Δ Indicator	-0.13		0.01		0.01		-0.22		-0.18***		0.42***		-0.00	
	(0.09)		(0.01)		(0.02)		(0.15)		(0.05)		(0.07)		(0.02)	
Δ Indicator	-0.19		0.11		-0.04		-1.14		-0.11		-0.16		-0.13	
* Non-agri exports/GDP	(0.56)		(0.07)		(0.10)		(0.84)		(0.47)		(0.73)		(0.10)	
Non-agri export price index	-1.30***	-1.36***	-0.71**	-0.72**	-1.11**	-1.25***	-0.99**	-1.27***	-1.15***	-1.39***	-1.18***	-1.37***	-0.52	-1.86
	(0.37)	(0.38)	(0.33)	(0.31)	(0.46)	(0.42)	(0.43)	(0.36)	(0.38)	(0.37)	(0.35)	(0.37)	(1.28)	(1.11)
Non-agri export price index	1.44**	1.32*	1.52***	1.51***	1.94***	2.00***	1.89***	2.03***	1.97***	2.26***	2.10***	2.34***	-	-
* good governance	(0.71)	(0.68)	(0.49)	(0.48)	(0.63)	(0.63)	(0.61)	(0.61)	(0.52)	(0.58)	(0.78))	(0.58)		
Observations	2742	2742	1956	1956	2689	2689	3046	3046	2972	2972	2989	2989	856	856
R-squared within	0.29	0.29	0.35	0.34	0.31	0.31	0.30	0.28	0.30	0.28	0.35	0.28	0.37	0.36

Notes: The dependent variable is the first-differenced log of real GDP per capita. All regressions are based on the specification in Table 5, column (3), and include country-specific fixed effects and regional time dummies. We only report the coefficients and standard errors of the variables of interest. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. See section 5 for an explanation of the indicators.

Table 7 Testing the channels of the resource curse (continued)

	(1)	(2)	(3)	(4)
	Estimates of long-run coefficients			
Real effective exchange rate index	0.11***		0.09**	
	(0.03)		(0.04)	
Real effective exchange rate index * Non-agri exports/GDP	-2.13***		-1.59**	
	(0.53)		(0.78)	
Public consumption	-1.23		-0.81	
	(0.89)		(1.07)	
Public consumption * Non-agri exports/GDP	-1.20		-8.69	
	(6.97)		(8.17)	
Private consumption	-1.35*		-1.13	
	(0.78)		(0.88)	
Private consumption * Non-agri exports/GDP	9.51**		4.51	
	(4.02)		(3.79)	
External debt			-0.05	
			(0.11)	
External debt * Non-agri exports/GDP			0.67	
			(0.71)	
Manufacturing			2.39*	
			(1.22)	
Manufacturing * Non-agri exports/GDP			-5.15	
			(7.14)	
Services			0.59	
			(0.75)	
Services * Non-agri exports/GDP			-1.60	
			(3.99)	
Non-agricultural export price index	0.78	-1.96*	1.10	-2.60*
	(0.97)	(1.13)	(0.89)	(1.30)
	Estimates of short-run coefficients			
Δ Real effective exchange rate index	-0.00		-0.03**	
	(0.02)		(0.01)	
Δ Real effective exchange rate index * Non-agri exports/GDP	-0.11		-0.08	
	(0.11)		(0.09)	
Δ Public consumption	-0.22		-0.14	
	(0.15)		(0.14)	
Δ Public consumption * Non-agri exports/GDP	-0.37		-1.12	
	(0.84)		(0.94)	
Δ Private consumption	-0.21***		-0.17***	
	(0.05)		(0.05)	
Δ Private consumption * Non-agri exports/GDP	0.51		0.06	
	(0.44)		(0.37)	
Δ External debt			-0.05**	
			(0.02)	
Δ External debt * Non-agri exports/GDP			0.03	
			(0.09)	
Δ Manufacturing			0.09	
			(0.26)	
Δ Manufacturing * Non-agri exports/GDP			0.29	
			(1.21)	
Δ Services			-0.04	
			(0.11)	
Δ Services * Non-agri exports/GDP			0.96*	
			(0.56)	
Observations	849	849	708	708
R-squared within	0.40	0.36	0.49	0.42

Notes: The dependent variable is the first-differenced log of real GDP per capita. All regressions are based on the specification in Table 5, column (3), and include country-specific fixed effects and regional time dummies. We only report the coefficients and standard errors of the variables of interest. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Appendix A Data description and sources

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

Commodity export price index Commodity export values for 1990 from UNCTAD Commodity Yearbook 2000 and UN International Trade Statistics 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 exports in total exports < 10% or share of 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 commodity exports in GDP (GDP in current US dollars, WDI). The sub-indices for non-agricultural and agricultural commodities are constructed in the same way.[34] 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

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

Review 2002). Since these are expressed in 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 ($\geq$ 0.5% of pop. affected, or damage $\geq$ 0.5% of GDP, or $\geq$ 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.

Appendix B Panel unit root and panel cointegration tests

The long-run equilibrium equation underlying equation (1) can be written as follows:

$$y_{i,t} = \frac{1}{\lambda}(\gamma_i + \theta' \tau_{i,t} + \beta_1' x_{i,t} + \eta_{i,t}) \qquad (2)$$

where γ_i is a country-specific fixed effect and $\tau_{i,t}$ is an $N \times 1$ vector of N country-specific time trends. Note that both the constant and the coefficient on the time trend are allowed to vary across countries. This is due to the fact that we left the country-specific fixed effect in equation (1) unrestricted. Therefore, it captures a country-specific constant in both the levels and the differenced equation. The first is represented by γ_i in equation (2), while the second implies a country-specific linear time trend in the levels equation (2), which is captured by $\theta' \tau_{i,t}$.[35]

Equation (1) allows us to estimate the long-run relationship in equation (2) if $y_{i,t}$ and $x_{i,t}$ are cointegrated. This implies that the individual variables are integrated of order 1 and the residuals of a regression of $y_{i,t}$ on $x_{i,t}$ are stationary. To test this, we first performed panel unit root tests on both the levels and the differences of the individual variables in $y_{i,t}$ and $x_{i,t}$ and then performed a panel cointegration test. The results are reported in Table B.1. We use the parametric panel unit root test by Im, Pesaran and Shin (2003, IPS hereafter) and the non-parametric test by Maddala and Wu (1999, MW hereafter). Both tests are based on augmented Dickey-Fuller (ADF) tests for the individual series in the panel and allow the test statistic to vary across groups. Under the null, all groups have a unit root while under the alternative one or more groups do not have a unit root.[36] Also, while IPS is designed for balanced panels, MW can be used for both balanced and unbalanced panels. We therefore report the test results of IPS for balanced panels, MW for balanced panels, and MW for unbalanced panels.[37] For the differences, the tests always reject the null of non-stationarity at 1 percent significance, which confirms that the variables are stationary in differences. For the level variables, the IPS test does not reject the null of non-stationarity, except for inflation. However, the MW tests for balanced and unbalanced panels reject the null of non-stationarity in most of the cases. It is important to note that rejection of the null does not mean that all series in the panel are stationary, but that at least one of the series is stationary. It is therefore possible that the tests reject

[35] To limit the number of regressors in the cointegration test, we leave out the regional time dummies.
[36] The oil import price index equals either zero or the world oil price index, which does not vary across oil importing countries. Therefore, a panel unit root test is inappropriate and we use a Dickey-Fuller test on the world oil price index instead.
[37] The unbalanced sample is the sample for which all the long-run variables are available, while the balanced sample includes 41 countries for which the long-run variables are available for (a minimum of) 42 years.

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 (augmented) Dickey-Fuller (ADF) test for the individual countries. Next to the test statistics, Table B.1 reports the number of countries for which the individual ADF test rejects the null of stationarity at 5%, as a ratio of the total number of countries in the sample. The results show that for the vast majority of countries, the ADF tests do not reject non-stationarity in the levels variables, while rejecting non-stationarity in the differenced variables. It therefore seems justified to assume the variables are integrated of order 1. We next perform a panel cointegration test, as suggested by Pedroni (1999). We first run the following regression for each country separately:

$$y_t = \alpha_0 + \alpha_1 \rho + \alpha_2' x_t + \varepsilon_t \qquad (3)$$

where y_t is log real GDP per capita in year *t*, ρ is a time trend, and x_t is a vector of long-run GDP determinants (trade openness, inflation, international reserves, the commodity export price index, and the oil import price index). This allows for country-specific fixed effects, country-specific time trends and country-specific coefficients for the long-run GDP determinants. We collect the residuals from these regressions and run ADF regressions for each country. Following Pedroni (1999), we allow the lag order of the dependent variable to vary across countries by including the lags that enter significant at 10 percent. We then calculate the mean ADF t-statistic, derive the group t-statistic, and express it in the form of equation (2) on p. 665 in Pedroni (1999). Table B.1 reports this standard normally distributed group t-statistic. We strongly reject the null hypothesis of no cointegration and thus conclude that the variables are cointegrated.[38]

[38] Intuitively, a rejection of the null hypothesis implies that a sufficiently large number of individual cross-sections have statistics that differ substantially from the means predicted by theory were they to be generated under the null.

Table B.1 Panel unit root and panel cointegration tests

	Panel unit root tests											
	Im, Pesaran, Shin, balanced sample				Maddala and Wu, balanced sample				Maddala and Wu, full sample			
	Levels		*Differences*		*Levels*		*Differences*		*Levels*		*Differences*	
GDP per capita (log)	-1.16	7/41	-3.72***	34/41	121.7***	7/41	453.6***	34/41	261.1	11/131	1488.5***	91/131
Trade to GDP	-1.36	3/41	-6.01***	41/41	77.03	3/41	1415.7***	41/41	353.0***	12/131	3593.4***	117/131
Inflation (log)	-2.22***	3/41	-4.83***	40/41	120.3***	3/41	698.2***	40/41	642.2***	16/131	1962.4***	102/130
Reserves to GDP	-1.65	7/41	-4.75***	39/41	132.5***	7/41	864.0***	39/41	349.7***	15/131	1918.1***	101/131
Commodity export price index	-1.64	4/41	-4.96***	41/41	125.8***	4/41	1251.1***	41/41	478.9***	19/131	3064.9***	118/131
	Dickey-Fuller, balanced				Dickey-Fuller, unbalanced							
	Levels		*Differences*		*Levels*		*Differences*					
Oil import price index	-1.62		-6.88***		-1.52		-7.10***					
	Panel cointegration test											
	Pedroni, full sample											
Group t-Statistic, $N(0,1)$	-5.67***											

Notes: Table B.1 reports the results of the panel unit root and panel cointegration tests. For the panel unit root tests, we report both the test statistic and the ratio of the number of countries for which the individual (augmented) Dickey-Fuller test rejects the null of stationarity at 5% to the total number of countries in the sample. The test statistics correspond to the t-bar statistic in Im, Pesaran, and Shin (2003), the Fisher χ^2 test statistic in Maddala and Wu (1999), and the group t-statistic, expressed in the form of equation (2) on p. 665 in Pedroni (1999). We included a constant but no trend in the panel unit root tests. Since equation (2) includes a time trend, we also ran the panel unit root tests for GDP per capita with a trend and found similar results. The choice of lag order in the panel unit root tests was based on a pooled (augmented) Dickey-fuller regression with fixed effects, except for the oil import price index, for which we ran an ordinary (augmented) Dickey-Fuller regression. The number of lags is 1, 0, 2, 1, 1, and 0 for GDP per capita, trade to GDP, inflation, reserves to GDP, the commodity export price index, and the oil import price index, respectively. *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively.