CSAE WPS/2007-15

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

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August, 2007

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 a substantial part of it is explained by high public and private consumption, low or inefficient total investment, and an overvalued exchange rate. Our results fully account for the cross-section results in the seminal paper by Sachs and Warner (1995).

Keywords: commodity prices; natural resource curse; growth

JEL classification: O13, O47, Q33

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* We would like to thank Chris Adam, Kofi Adjepong-Boateng, Robin Burgess, Mardi Dungey, Klaus Schmidt-Hebbel, Ron Smith, Måns Söderbom, and participants at the UNU-WIDER Conference "Aid: Principles, Policies, and Performance", Helsinki, June 2006, the LSE and UCL Development and Growth seminar, London, February 2007, the CSAE Conference "Economic Development in Africa", Oxford, March 2007, the ESRC's World Economy and Finance Research Programme advisory meeting, Oxford, April 2007, the G-20 Workshop "Commodity Cycles and Financial Stability", Washington DC, May 2007, and the Treasury/DfID seminar "Economics of Africa", London, June 2007, for useful comments. We also acknowledge support of the UK Economic and Social Research Council (ESRC) under the project *Managing Macroeconomic Risks in Developing Countries: Policies and Institutions* and the UK Department for International Development under the project *Improving Institutions for Pro-Poor Growth*. All remaining errors are our own.

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1. Introduction

A large literature suggests that there is a 'resource curse': natural resource abundant countries tend to grow slower than resource scarce countries.[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 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

[4] This empirical finding is documented in Sachs and Warner (1995, 2001), Gylfason et al. (1999), Leite and Weidmann (1999), Auty (2001), Bravo-Ortega and De Gregorio (2001), and Sala-i-Martin and Subramanian (2003). Van der Ploeg (2007) provides a survey of the resource curse literature.

[5] Few studies use panel data. 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. Murshed (2004) uses a panel of 91 developing countries from 1970 till 2000 and finds that the presence of point-source resources harms growth through retarding democratic and institutional development.

different types of commodities. We investigate all the transmission channels of the resource curse proposed in the literature and we test the extent to which the results of the cross-sectional growth regressions in Sachs and Warner (1995) are explained by the mechanisms we identify in our estimations. Finally, 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. Our approach also enables us to investigate possible transmission channels in a systematic manner. We find that none of the transmission channels proposed in the literature individually accounts for the curse. However, a combination of public and private consumption, total investment, and exchange rate overvaluation explains a substantial part of it. We also find that the negative long run effects of commodity booms that we identify in our estimations fully account for the cross-section results in Sachs and Warner (1995). In fact, once we control for these long-run adverse effects, resource abundance has a positive effect on average cross-country growth rates.

The rest of this paper is structured as follows. Section 2 describes the empirical analysis, including the error-correction specification, the variables used in estimation, the cointegration tests, and a description of the data. Section 3 reports the estimation results and simulates the short and long run effects of higher commodity export prices on growth. Section four investigates whether the resource curse occurs conditional on governance. Section five deals with the endogeneity of resource dependence and governance. Section six tests the importance of the resource curse transmission channels, as proposed in the theoretical literature. Section seven investigates to what extent our results explain the findings in the empirical literature. Section eight 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:

$$\Delta Y_{i,t} = \alpha_i + \delta_t + \lambda Y_{i,t-1} + \beta_1 X_{i,t-1} + \beta_2 \Delta Y_{i,t-k} + \beta_3 \Delta X_{i,t-k} + \beta_4 S_{i,t} + u_{i,t} \qquad (1)$$

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 δ_t is a vector of regional time dummies.[6] $X_{i,t-1}$ is a vector of variables that is expected to affect GDP both in the short run and long run.

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; the log of inflation, measured as the annual percentage change in consumer prices; 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 measures of public, private, and total investment, indicators of institutional quality, government consumption, democracy, and capital account openness, the black market premium, and the number of assassinations. 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 a vector of 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 natural disasters (geological, climatic, and human disasters).

[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.

[7] In order to test the long run effects of commodity prices we need a panel with relatively large T. The time span of the observations used in estimation is 1963-2003. Many of the available control variables have missing data for the early years in our sample and would therefore severely limit the sample size. 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 effects.

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. Next, we discuss how the commodity price indices were constructed.

2.1 Constructing commodity price indices

The commodity export price index was constructed using the methodology of Deaton and Miller (1995) and Dehn (2000). In particular, we collected data on world commodity prices and commodity export values for as many commodities as data availability allowed. Table 1b lists the 58 commodities in our sample. For each of the countries, we calculate the total value of commodity exports in 1990. 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. Finally, to allow the effect of commodity export prices to be larger for countries with higher commodity exports, we weigh the log of the deflated index by the share of commodity exports in a country's GDP. The separate indices for non-agricultural and agricultural commodities were constructed in a similar 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

[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).

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 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 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.

[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. We found a similar result when using the specification of Table 2, column (4), discussed below. These additional results are available upon request.

[10] Recall that the commodity export price index is weighed 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.

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 statistically significant at 1 percent. The size of the coefficient suggests that the speed of adjustment to long-run equilibrium (i.e. the speed at which deviations from long-run equilibrium are corrected) 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 statistically 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, although these effects are not statistically 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 year of the coup. The negative impact of civil war is estimated to be similar at 2.2 percentage points for each year of the war, this being consistent with Collier (1999) who also estimates a growth loss of 2.2 percentage points. 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

[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.

[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.

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.

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

[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.

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 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 sustainably higher output. These findings support the hypothesis that the resource curse occurs conditional on bad governance.

The agricultural commodity export price index enters with a positive sign and is insignificant, while its interaction with the good governance dummy enters with a negative sign 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 and further support the hypothesis that the resource curse occurs conditional on bad governance. In column (3), the general commodity export price index again enters negative and is significant at 1 percent, while its interaction with the good governance dummy is again positive but is now significant at 1 percent. In column (4), the non-agricultural commodity export price 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 and available upon request. In particular, the results for the general commodity export price index and the sub-indices for non-agricultural and agricultural commodities 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 coefficient of the non-agricultural commodity export price index enters with a negative sign and is statistically 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 coefficient of the non-agricultural commodity export price index is now 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] All in all, these results provide 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.

[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. These additional results are available upon request.

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 and the dummy for good governance are endogenous, i.e. correlated with the error term in equation (1).[18] 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. However, as explained in section 2.1, we weigh the index by the ratio of commodity exports over GDP to allow the effect of commodity prices to be larger for more resource-dependent countries. As argued by Brunnschweiler and Bulte (2006), this ratio is potentially driven by variables that also affect growth, such as government policies or institutions, and thus might suffer from endogeneity.[19] As far as the endogeneity relates to omitted time-invariant variables, the inclusion of fixed effects in our estimations solves the problem. But if the endogeneity instead relates to omitted time-varying regressors or reverse causality, our estimated coefficients could be biased. In addition, the dummy for good governance also potentially suffers from endogeneity, which could lead to a biased coefficient of the interaction term.

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, and for the good governance dummy. For comparison, Table 5, columns (1) and (3), 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 results confirm the earlier long-run resource curse effect for countries with bad governance and the absence of a resource curse in countries with good governance. Both the non-agricultural commodity export price index and its interaction with the good governance dummy are statistically significant at 1 percent in both specifications. The short-run effect of higher non-agricultural commodity prices is positive and the effect is largest in the first year after the price increase, which is consistent with the results for the general commodity export price index in Table 3.

[18] Endogeneity is sometimes understood to mean reverse causality only. We use it more generally to describe a non-zero correlation between the regressor and the error term.

[19] Government policies and institutions are very likely to affect the ratio of commodity exports over GDP through the denominator (GDP).

We next instrument for the ratio of non-agricultural commodity exports over GDP and the good governance dummy. As an instrument for non-agricultural exports over GDP, we use estimates of the stock of sub-soil assets (minerals) in 2000 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. These 13 commodities closely resemble the 15 non-agricultural commodities in our index. In fact, 12 commodities are included in both groups. We can therefore use the value of sub-soil assets in current US dollars per capita as an instrument for the ratio of non-agricultural commodity exports over GDP. Because this value refers to the stock of resources in the ground, we believe it to be exogenous with respect to a country's growth rate. The ratio of non-agricultural commodity exports over GDP does 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 in the following way. We again construct the log of the deflated non-agricultural commodity export price index, as before, but we do not weight it by the ratio of non-agricultural commodity exports over GDP. Instead, we weight it by the 2000 value of sub-soil assets in current US dollars per capita. We then use the lagged level, difference, and two lagged differences of this newly constructed variable as instruments for the lagged level, difference, and two lagged differences of the non-agricultural export price index in Table 5, columns (1) and (3).

In addition, we need an instrument for the interaction term of the non-agricultural commodity export price index with the good governance dummy. The best instrument for governance or institutions 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 instruments, 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. Using these three instruments, we construct an instrument for the interaction term of the index with the good governance dummy in the following way. We first run a probit regression of the good governance dummy on the three

[20] The World Bank estimates of natural capital were earlier used by Brunnschweiler and Bulte (2006) and Arezki and Van der Ploeg (2007) to proxy resource abundance.

instruments for the sample in Table 5, column (1), and collect the fitted values. All three instruments enter with the expected positive signs and are statistically significant at 1 percent. The pseudo R-squared is 0.69. We interact the fitted values of the probit regression with the instrument for the non-agricultural commodity export price index discussed above.[21] This yields an additional instrument for the interaction term of the non-agricultural commodity export price index and the good governance dummy in Table 5, columns (1) and (3).

We next use all our constructed instruments to perform a two-stage-least-squares estimation procedure. The second-stage results of this procedure are reported in Table 5, columns (2) and (4).[22] 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.[23] The coefficients of the interaction of the index with the good governance dummy are also similar to the coefficients in Table 5, columns (1) and (2), although no longer statistically 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 for the interaction terms in columns (2) and (4), respectively. These results indicate that the OLS estimates in columns (1) and (3) are likely to be consistent and do not suffer from any substantial endogeneity bias.

The short-run coefficients of the non-agricultural commodity export price 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 commodity export price index in column (4).[24] This

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[21] Goderis and Ioannidou (forthcoming, 2007) perform a similar procedure to construct instruments, following Wooldridge (2002), p. 237.

[22] 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.

[23] The Davidson-MacKinnon test is similar to the Durbin-Wu-Hausman test of endogeneity.

[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.

evidence suggests that any endogeneity bias is likely to be small and if anything leads to a small underestimation of the positive short-run growth effect of higher non-agricultural commodity export prices.

All in all, 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 six candidate explanations for the resource curse effect: Dutch disease, governance, conflict, excessive borrowing, inequality and volatility. 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, 1995, 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 (Matsuyama, 1992; Michaels, 2006). To test for the importance of this channel, we add an index of real exchange rate overvaluation[25] to the specifications in Table 3, columns (3) and (4). As an overvalued exchange rate could potentially affect GDP both in the short run 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 exchange rate overvaluation is different for resource-abundant countries, we also include interaction terms of the level and differenced overvaluation index with the share of non-agricultural exports in GDP. If the negative long-run effect of non-agricultural commodity export prices works

[25] We use the index of real exchange rate overvaluation from the Global Development Network Growth Database of the NYU Development Research Institute.

through their impact on the real exchange rate, then the effect of the export price indices should disappear once we control for exchange rate overvaluation. The results are reported in the two columns in the top left corner of Table 6. In the first column, the level of the overvaluation index enters negative, although not significant, suggesting that, consistent with Dutch disease, an overvalued exchange rate indeed has a negative effect on long-run GDP per capita. The interaction of the index with the share of exports in GDP also enters negative, suggesting that the effect of an overvalued exchange rate is more severe in resource-abundant countries. The differenced index enters positive, but for resource-abundant countries this effect is smaller or even negative as the interaction term enters negative. All in all, these results seem to suggest that, at least in the long run, an overvalued exchange rate lowers GDP. In addition, adding the overvaluation index leads to a slightly smaller 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 overvaluation index. More specifically, the long-run coefficient changes from -1.25 to -1.11, which suggests that Dutch Disease explains around 11 percent of the long-run resource curse effect.[26]
Although countries with good governance do not suffer from a resource curse, their long-term gain from higher commodity export prices might be negatively affected by Dutch Disease.[27] This long-term gain is captured by the linear combination of the coefficients of the non-agricultural export price index and its interaction with the good governance dummy. This combination changes from 0.75 for the sample without the overvaluation index to 0.83 in the sample with the overvaluation index, indicating that Dutch Disease also harms long-run GDP in good governance countries.

We next explore whether the resource curse induces weak institutions of governance. The literature has proposed several such routes. Resource rents may invite non-productive lobbying and rent seeking, as in Lane and Tornell (1996), 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

[26] Aguirre and Calderón (2005) show that the effect of real exchange-rate misalignment is non-linear: large overvaluations and very large undervaluations reduce growth, while moderate undervaluation raises growth. We experimented with these non-linearities by allowing the effects of over- and undervaluation to be different and to be non-linear (we thank Klaus Schmidt-Hebbel for this suggestion). In both cases, adding these additional exchange rate indicators did not cause substantial changes in the model's goodness of fit or the significance of the Dutch Disease indicators. We also further tested the importance of Dutch Disease by controlling for the shares of manufacturing and services in GDP, sectors that would arguably suffer from Dutch Disease. The results are consistent with our finding that Dutch Disease explains a small part of the resource curse. We report these results in the bottom part of Table 6, when we investigate the routes through which governance drives the resource curse.
[27] 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.

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. Acemoglu et al. (2004) and 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). The protection of domestic industries as a possible explanation of the resource curse is also emphasized by Arezki and Van der Ploeg (2007). 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 parallel market exchange rate premium[28], civil liberties and political rights (Freedom House), two measures of political constraints (Henisz, 2000, 2002), democracy, autocracy, and a combined measure of democracy and autocracy (Polity IV), checks and balances (Database of Political Institutions 2004), and the Composite International Country Risk Guide (ICRG) risk rating (PRS Group). To save space, the third and fourth column of the top left corner of Table 4 ("Governance") only report the results for the parallel market exchange rate premium since adding any of these other indicators scarcely changes the long run results. The parallel market premium enters negative and is significant at 1 percent. This indicates that countries with high premiums on the parallel market exchange rate grow more slowly. For resource-abundant countries, this effect is smaller, as the interaction term of the parallel market premium with the share of exports in GDP enters positive and is significant at 5 percent. The short-run effect of the parallel market premium is also negative and is significant at 5 percent, while not significantly different for resource-abundant countries. Although these results indicate that this governance indicator is an important GDP determinant, it does not lead to a smaller resource curse effect. In fact, the long-run coefficient of commodity prices even increases in absolute size. This suggests that the deterioration of governance is not the central explanation of the

[28] We use the log of the black market premium from the Global Development Network Growth Database of the NYU Development Research Institute.

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 four other proposed channels for the resource curse. 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). Resource abundance can also tempt a government into excessive external borrowing, as in Mansoorian (1991), Manzano and Rigobon (2006), and Kuralbayeva and Vines (2006). It also exposes countries to commodity price volatility which could, for example, discourage investment (Sala-i-Martin and Subramanian, 2003). Finally, as suggested by Sokoloff and Engerman (2000), resource abundance can lead to increased inequality and this could harm growth through lower savings or investment rates, lower quality of institutions, credit-market imperfections, distorting redistribution, or socio-political unrest (Barro, 2000).

We investigate the importance of these channels through the same approach.[29] 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.[30]

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-offs from these activities is high in countries with grabber-friendly ("bad") institutions but low in countries with producer-friendly ("good") institutions. This leads entrepreneurs in countries with bad institutions away from productive activities into non-productive rent-seeking

[29] We use the following indicators for conflict, excessive borrowing, inequality and volatility, respectively: the cumulative number of civil war years; total external debt to gross national product, from the 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.

[30] The coefficient of the interaction term of the export price index with the good governance dummy is statistically insignificant for the specifications under "Excessive borrowing". This is due to a very low availability of the external debt variable for good governance countries.

activities, which in the long run slows down industrial development. We empirically test this theory by adding a measure of industrial development, the share of manufacturing in GDP, to our specification. If the resource curse works through the underdevelopment of the manufacturing sector, then the effect of the export price index should disappear once we control for the size of this sector. We also investigate the importance of the high-productivity services sector.[31] The results for manufacturing are reported in the two columns in the bottom left corner of Table 6. The level of manufacturing enters with a positive sign, indicating that a larger manufacturing sector leads to a higher level of long-run GDP. However, this effect is not statistically significant. Also, the negative coefficient on the interaction term with the share of exports suggests that the effect of manufacturing is smaller for resource-rich economies. Although we find some evidence that a larger manufacturing sector is good for long-run growth, this does not seem to explain much of the resource curse effect. The absolute size of the coefficient is only around 3 percent smaller when controlling for the share of manufacturing in GDP.[32] We can therefore conclude that the resource curse does not seem to work through a slower speed of industrial development. It also does not work through lower growth in the services sector, as can be seen from the results in the third and fourth column of the bottom part of Table 6.[33]

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" (Robinson et al., 2006). 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

[31] Manufacturing and services as a share of GDP were both taken from the World Development Indicators.
[32] The coefficient of the interaction term of the export price index with the good governance dummy is statistically insignificant for the specifications under "Manufacturing". This is due to a lower availability of the manufacturing indicator for good governance countries.
[33] As we will show below, we find some evidence of the importance of the services sector once we control for other important channel indicators, i.e. exchange rate overvaluation, public and private consumption, and total investment.

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 trade openness to our specification.[34] If the resource curse works through the provision of public sector jobs, then the effect of the export price index should disappear once we control for public consumption, which amongst other things includes civil servant salaries. Similarly, if the curse works through trade protection, then controlling for trade openness should make the resource curse effect disappear. The results for openness are reported in the fifth and sixth columns of the bottom part of Table 6 and indicate that more open countries grow faster, although the effects are not significant. However, the long-run resource curse effect hardly changes when controlling for this possible channel, suggesting that trade protection is not an important driver of the resource curse. The results for public consumption are reported in the next two columns of the bottom part of Table 6. The long-run and short-run coefficients are all negative, although not significant, suggesting that higher levels of public consumption negatively affect growth, both in the short run and in the long run. The negative effect of public consumption also seems to explain part of the resource curse effect. When controlling for government consumption, the absolute size of the coefficient falls by 22 percent. These results support the argument of Robinson et al. (2006) that commodity booms lead to inefficient public sector employment provision which then slows down economic development.

In both explanations of why governance is crucial for the occurrence of the curse, commodity booms lead workers or entrepreneurs away from productive activities into less productive rent-seeking or public sector activities. With this shift away from productive activities, 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. In addition to lowering investment levels, commodity booms may also lead to a lower quality of investment projects. Robinson and Torvik (2005) provide a theory in which "white elephants", investment projects

[34] As a measure of public consumption we use general government final consumption expenditure as a share of GDP, taken from the 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 and available at http://www.cid.harvard.edu/ciddata.

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 resource curse. In addition to public sector employment, trade protection, subsidies, and regulation, inefficient redistribution during and after commodity booms can also occur 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.[35] The results are reported in the last four columns of the bottom part of Table 6. We find that the share of private consumption in GDP has a negative effect on GDP growth, both in the short run and in the long run. Both effects are statistically significant at 1 percent. However, we also find that the long-run effect is significantly smaller for resource-rich economies, as the interaction term enters with a positive sign and is significant at 1 percent. Just as in the case of public consumption, controlling for private consumption takes away part of the long-run resource curse effect. The absolute size of the coefficient falls by 17 percent.

We next turn to the results for total investment. The long-run and short-run coefficients are positive and statistically significant at 1 percent. This indicates that higher levels of total (public and private) investment lead to higher GDP levels, as one would expect. The positive long-run effect of investment is significantly smaller in resource-rich economies. This is clearly consistent with the theory of inefficient redistribution through inefficient investment projects in Robinson and Torvik (2005). Controlling for the level of total investment and the lower long-run return on investment in resource-rich economies also explains part of the resource curse effect, as the absolute size of the long-run coefficient falls by 14 percent.

So far we have only considered indicators individually. We next investigate whether combinations of indicators can explain the resource curse effect in our estimations. We start by combining the four variables that are individually important in explaining the resource curse: exchange rate overvaluation, public consumption, private consumption, and total investment. The results are reported in the first two columns of Table 7. We again find that exchange rate overvaluation negatively affects GDP in the long run, while public consumption and private consumption negatively affect GDP both in the short run and in the long run. However, of these effects only

[35] 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.

the long-run effect of public consumption is statistically significant at 10 percent. In addition, we again find a positive and highly significant long-run effect of total investment on GDP, both in the short and in the long run. The interaction term of the level of total investment with the share of exports in GDP again enters negative, although the coefficient is now statistically insignificant. The combination of the four variables explains a substantial part of the resource curse effect. Once we control for these variables, the absolute size of the long-run resource curse coefficient falls by 47 percent, while its statistical significance disappears.

We next test whether the four variables all contribute to the lower long-run resource curse effect by eliminating each of them individually and observing the change in the long-run coefficient, while keeping the sample constant. We find that indeed all four variables are important in explaining the curse. 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 two exceptions are volatility and the share of services in GDP. If we control for each of these two additional variables, the absolute size of the long-run resource curse coefficient further decreases. Table 7, columns (3) and (4), report the results when we add both of these variables to the four variables that were included in the specifications of columns (1) and (2). While volatility has a positive but highly insignificant long-run effect on GDP, the large negative coefficient of the interaction of volatility with the share of exports in GDP suggests that the long-run effect of volatility is negative for countries with substantial commodity exports. The long-run effect of a growing services sector is positive, although insignificant. Its short-run effect is negative, but the positive and significant interaction with the share of exports in GDP suggests that this effect is positive for countries with substantial commodity exports. These results provide some evidence that volatility and growth in the services sector affect GDP in resource-rich economies. The combination of the six variables explains almost the entire long-run resource curse effect. Once we control for these variables, the absolute size of the long-run resource curse coefficient falls by 95 percent and becomes highly insignificant.

We conclude that a substantial part of the resource curse effect can be explained by exchange rate overvaluation, public and private consumption, total investment, and to a lesser extent volatility and services. This supports recent theory that points at the importance of inefficient redistribution through public sector employment provision or inefficient investment projects (white elephants). It also supports the more general idea that commodity booms lead the entrepreneurs in an economy away from productive activities and into non-productive rent-seeking, lobbying, or public sector activities. Finally, it lends some support to the large literature that stresses the importance of Dutch disease in resource-rich economies.

7. Explaining the results in the empirical literature

Finally, we investigate to what extent our empirical findings can explain the cross-section results in the empirical resource curse literature. We first collect the data from the seminal paper by Sachs and Warner (1995) (S&W hereafter)[36] and replicate their main results. Table 8, columns (1) to (5), report the results of the specifications in Table I of the revised 1997 version of S&W. The dependent variable is the 1970-1990 average annual growth in real GDP divided by the economically active population. The regressors are the 1970 share of primary exports in GNP (sxp), the 1970 log of real GDP divided by the economically-active population (lgdpea70), the fraction of years between 1970 and 1990 in which the country is rated as an open economy (sopen), the average 1970-1989 log of the ratio of real gross domestic investment to real GDP (linv7089), a rule of law index (rl), and the 1970-1990 average annual growth in the log of the external terms of trade.

The results in Table 8, columns (1) to (5), are very similar to the results in S&W. In particular, the coefficients of the share of primary exports in GNP (sxp) are always negative and statistically significant at 1 percent, as in S&W.[37] This is the familiar resource curse effect. To explore how much of this effect can be explained by the long-run adverse effect of commodity prices in our estimations, we run two regressions. The first regression repeats the specification in Table 5, column (1), but excluding the level of the non-agricultural commodity export price index. The second

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[36] We use the data and estimations from the revised 1997 version of the paper, available at http://www.cid.harvard.edu/ciddata.
[37] The coefficients of the other regressors are also very similar. We do not exclude outliers and hence our sample is slightly larger than the sample in S&W. However, all our results go through when we exclude outliers, as in S&W. These results are available upon request.

regression too repeats the specification in Table 5, column (1), but this time with the level of the index included. Hence, the only difference between the two regressions is that the first does not control for our long-run resource curse effect, while the second does. We collect the country-specific fixed effects from both regressions and use each set of fixed effects as a dependent variable in the S&W regressions. Table 8, columns (6) to (10), reports the results of the S&W regressions when replacing the dependent variable of S&W by the fixed effects from the regression *without* the level of the non-agricultural commodity export price index. Our results are very similar to the S&W results. In particular, the coefficients of the share of primary exports in GNP (sxp) are again always negative and statistically significant at 1 percent. In other words, when we use the fixed effects from the regression in which we do not control for the negative long-run effect of non-agricultural commodity prices, the S&W resource curse effect remains.

We next repeat the same procedure but now using the fixed effects from the regression *with* the level of the commodity export price index. The results are reported in Table 8, columns (11) to (15). The coefficients of the share of primary exports in GNP (sxp) are now positive in two out of the five specifications and equal to zero or very close to zero in the three other specifications. All five coefficients are highly insignificant. These results indicate that the S&W resource curse effect disappears once we control for our long-run negative effect of commodity prices.

We test the robustness of these results by replacing the share of primary exports in GNP (sxp) by the 1971 share of mineral production in GNP (snr), an alternative measure of resource intensity also taken from S&W. This measure is particularly suited as it nicely squares with our finding that the resource curse is confined to non-agricultural commodities (minerals). The estimation results are reported in Table 8, columns (16) to (30). Again, our long-run negative effect of non-agricultural commodity prices fully explains the resource curse effect in S&W. In fact, the coefficients of the share of mineral production in GNP (snr) in the specifications of Table 8, columns (26) to (30), are positive and statistically significant in four out of the five specifications. So once we control for the adverse long-run effects of higher mineral prices, minerals are a blessing, not a curse.

These results provide robust evidence that the S&W cross-sectional resource curse effect disappears once we control for the long-run negative effect of non-agricultural commodity prices. We therefore believe that our panel data analysis on the short- and long-run effects of commodity booms *within* countries also provides a comprehensive explanation of why the empirical resource curse literature so often finds a negative effect of resource abundance on growth in cross-section regressions.

8. 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 an overvalued exchange rate, high public and private consumption, low or inefficient investment, and to a lesser extent commodity price volatility and slow growth in the services sector explain a substantial part of the curse. These findings are consistent with recent theory that points at inefficient redistribution in return for political support as the root of the curse but also lend some support to the large Dutch disease literature. In addition, the results support the more general idea that commodity booms lead countries away from productive activities and provide incentives for non-productive activities, such as rent-seeking, lobbying, or public sector employment. We find that the negative long-run effects of commodity booms that we identify in our estimations fully account for the cross-section results in the seminal paper by Sachs and Warner (1995). Once we control for these long-run adverse effects, non-agricultural resource abundance has a positive effect on average cross-country growth rates.

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Table 1a Summary Statistics

	Obs.	Mean	St. Dev.	Min.	Max.
GDP per capita (log)	3608	7.56	1.56	4.31	10.55
Trade to GDP	3583	0.65	0.36	0.06	2.29
Inflation (log)	3601	0.13	0.29	-0.24	5.48
Reserves to GDP	3602	0.10	0.10	0.00	1.24
Commodity export price index	3608	0.39	0.42	0.00	2.67
Commodity exports to GDP	3608	0.08	0.09	0.00	0.45
Non-agricultural commodity export price index	3608	0.21	0.40	0.00	2.67
Non-agricultural commodity exports to GDP	3608	0.04	0.08	0.00	0.40
Agricultural commodity export price index	3608	0.18	0.22	0.00	1.13
Agricultural commodity exports to GDP	3608	0.04	0.05	0.00	0.24
Dummy good governance	3608	0.22	0.41	0	1
Oil import price index	3608	3.13	1.86	0.00	4.96
Δ GDP per capita (log)	3608	0.02	0.05	-0.36	0.30
Δ Trade to GDP	3583	0.01	0.08	-0.88	1.21
Δ Inflation (log)	3601	0.00	0.18	-3.62	2.52
Δ Reserves to GDP	3602	0.00	0.03	-0.37	0.31
Δ Commodity export price index	3608	0.00	0.02	-0.27	0.41
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

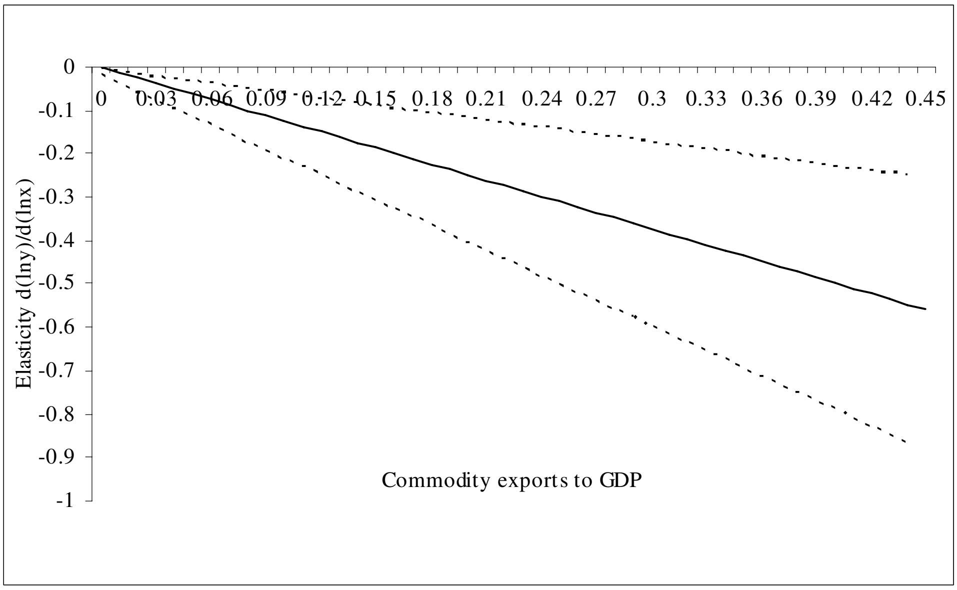
		Non-agricultural		
Aluminum	Gasoline	Natural gas	Phosphatrock	Uranium
Coal	Ironore	Nickel	Silver	Urea
Copper	Lead	Oil	Tin	Zinc
		Agricultural		
Bananas	Fish	Oliveoil	Rice	Sunfloweroil
Barley	Fishmeal	Oranges	Rubber	Swinemeat
Beef	Groundnuts	Palmkerneloil	Shrimp	Tea
Butter	Groundnutoil	Palmoil	Sisal	Timber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

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

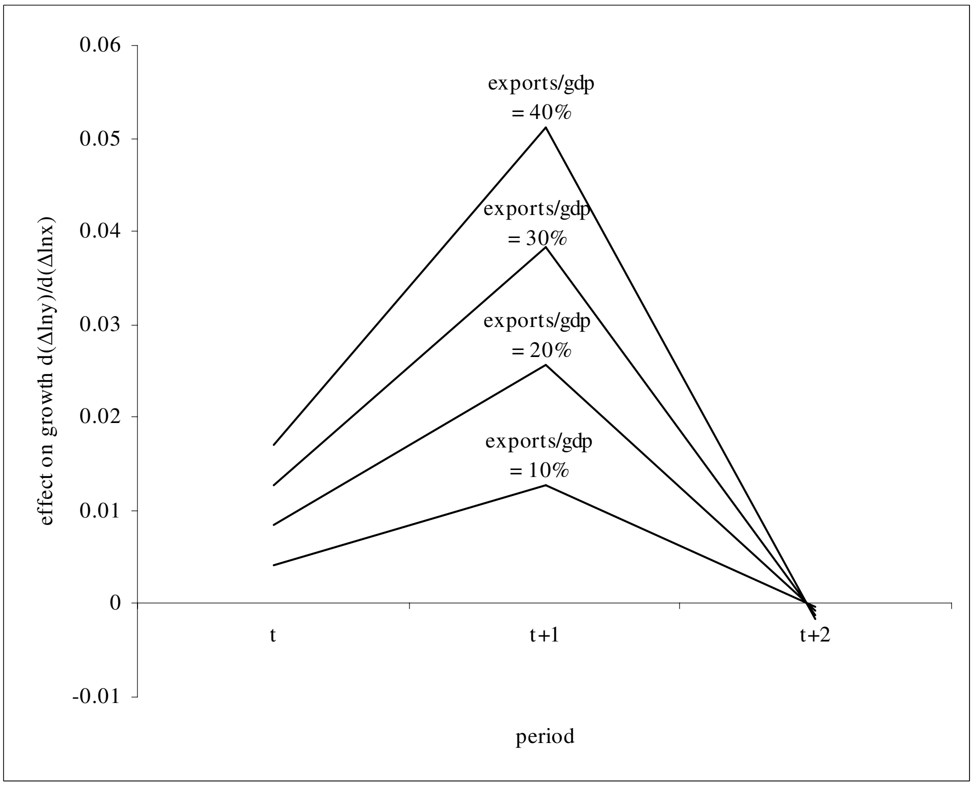
Notes: The dependent variable is the first-differenced log of real GDP per capita. Robust standard errors are clustered by country and are reported in parenthesis. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

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***	
	(1.509)		(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 parenthesis. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively.

Table 4 Estimation results: subsamples good and bad governance

	(1) bad gov.	(2) good gov.	(3) bad gov.	(4) good gov.
	Estimates of long-run coefficients			
Trade to GDP	0.680***	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 parenthesis. ***, **, 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 parenthesis. ***, **, 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	
Indicator	-0.10		-0.12***		0.01		-0.18**		-0.72		0.93	
	(0.07)		(0.04)		(0.02)		(0.08)		(1.01)		(1.43)	
Indicator * Non-agri exports/GDP	-0.46		0.61**		0.02		0.46		5.22		-0.99	
	(0.43)		(0.24)		(0.17)		(0.32)		(9.03)		(6.28)	
Δ Indicator	0.01		-0.01**		-0.02***		-0.07***		-0.12		-0.27**	
	(0.02)		(0.00)		(0.01)		(0.02)		(0.09)		(0.13)	
Δ Indicator * Non-agri exports/GDP	-0.04		0.03		-0.06		0.17*		1.97**		0.37	
	(0.10)		(0.03)		(0.04)		(0.08)		(0.87)		(0.87)	
Non-agri export price index	-1.11**	-1.25***	-1.16***	-1.05***	-1.36***	-1.29***	-1.66***	-1.46***	-1.14**	-0.99*	-1.30***	-1.29***
	(0.46)	(0.42)	(0.30)	(0.27)	(0.37)	(0.35)	(0.47)	(0.40)	(0.50)	(0.52)	(0.36)	(0.35)
Non-agri export price index * good governance	1.94***	2.00***	2.07***	2.03***	2.14***	2.12***	-0.72	1.14	1.64***	1.69***	2.14***	2.12***
	(0.63)	(0.63)	(0.57)	(0.53)	(0.63)	(0.61)	(4.72)	(4.36)	(0.60)	(0.61)	(0.61)	(0.61)
Observations	2689	2689	2317	2317	3087	3087	1872	1872	1752	1752	3087	3087
R-squared within	0.31	0.31	0.33	0.32	0.28	0.28	0.32	0.29	0.37	0.37	0.28	0.28
	Manufacturing		**Services**		**De jure openness**		**Public consumption**		**Private consumption**		**Total investment**	
Indicator	1.34		0.05		0.00		-0.60		-1.14***		2.60***	
	(0.95)		(0.50)		(0.08)		(0.74)		(0.40)		(0.58)	
Indicator * Non-agri exports/GDP	-7.58*		2.55		0.95		-5.08		6.23***		-7.54***	
	(3.98)		(2.60)		(0.70)		(6.75)		(2.17)		(2.57)	
Δ Indicator	0.07		-0.13		0.01		-0.22		-0.18***		0.42***	
	(0.15)		(0.09)		(0.01)		(0.15)		(0.05)		(0.07)	
Δ Indicator * Non-agri exports/GDP	-0.46		-0.19		0.11		-1.14		-0.11		-0.16	
	(1.45)		(0.56)		(0.07)		(0.84)		(0.47)		(0.73)	
Non-agri export price index	-1.41***	-1.45***	-1.30***	-1.36***	-0.71**	-0.72**	-0.99**	-1.27***	-1.15***	-1.39***	-1.18***	-1.37***
	(0.36)	(0.37)	(0.37)	(0.38)	(0.33)	(0.31)	(0.43)	(0.36)	(0.38)	(0.37)	(0.35)	(0.37)
Non-agri export price index * good governance	0.65	0.81	1.44**	1.32*	1.52***	1.51***	1.89***	2.03***	1.97***	2.26***	2.10***	2.34***
	(0.61)	(0.62)	(0.71)	(0.68)	(0.49)	(0.48)	(0.61)	(0.61)	(0.52)	(0.58)	(0.78))	(0.58)
Observations	2396	2396	2742	2742	1956	1956	3046	3046	2972	2972	2989	2989
R-squared within	0.31	0.30	0.29	0.29	0.35	0.34	0.30	0.28	0.30	0.28	0.35	0.28

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 exchange rate overvaluation	-0.07		-0.01	
	(0.06)		(0.07)	
Real exchange rate overvaluation * Non-agri exports/GDP	-0.46		-1.35**	
	(0.49)		(0.57)	
Public consumption	-1.75*		-1.69*	
	(0.91)		(0.90)	
Public consumption * Non-agri exports/GDP	5.81		4.46	
	(8.76)		(8.49)	
Private consumption	-0.89		-0.99*	
	(0.58)		(0.57)	
Private consumption * Non-agri exports/GDP	3.86		4.60	
	(3.12)		(3.03)	
Total investment	2.20***		2.23***	
	(0.71)		(0.68)	
Total investment * Non-agri exports/GDP	-5.15		-5.34	
	(3.94)		(3.82)	
Volatility			0.96	
			(1.24)	
Volatility * Non-agri exports/GDP			-10.35	
			(7.16)	
Services			0.37	
			(0.41)	
Services * Non-agri exports/GDP			-0.40	
			(2.56)	
Non-agricultural export price index	-0.78	-1.46***	-0.07	-1.32***
	(0.48)	(0.45)	(0.45)	(0.42)
Non-agricultural export price index * good governance	1.75**	2.23***	0.18	1.26*
	(0.73)	(0.63)	(1.05)	(0.66)
	Estimates of short-run coefficients			
Δ Real exchange rate overvaluation	0.01		0.02	
	(0.02)		(0.02)	
Δ Real exchange rate overvaluation * Non-agri exports/GDP	0.04		0.01	
	(0.07)		(0.07)	
Δ Public consumption	-0.26		-0.22	
	(0.16)		(0.16)	
Δ Public consumption * Non-agri exports/GDP	-0.70		-1.01	
	(0.85)		(0.98)	
Δ Private consumption	-0.06		-0.04	
	(0.05)		(0.05)	
Δ Private consumption * Non-agri exports/GDP	-0.66*		-0.70*	
	(0.34)		(0.39)	
Δ Total investment	0.38***		0.40***	
	(0.08)		(0.08	
Δ Total investment * Non-agri exports/GDP	0.21		0.22	
	(0.53)		(0.49)	
Δ Volatility			-0.19	
			(0.13)	
Δ Volatility * Non-agri exports/GDP			1.11	
			(0.79)	
Δ Services			-0.12	
			(0.09)	
Δ Services * Non-agri exports/GDP			0.73*	
			(0.41)	
Observations	2577	2577	2367	2367
R-squared within	0.42	0.31	0.43	0.31

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.

Table 8 Explaining the results in Sachs and Warner (1995)

	Dependent variable: average gdp growth 1970-1990 (Sachs and Warner, 1995)					Dependent variable: fixed effects from Table 5, column (1), excluding the long-run effect of commodity prices					Dependent variable: fixed effects from Table 5, column (1), including the long-run effect of commodity prices				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
sxp	-6.71***	-5.90***	-6.99***	-8.48***	-8.24***	-0.17***	-0.16***	-0.17***	-0.18***	-0.18***	-0.00	-0.00	-0.02	0.00	0.01
	(1.93)	(1.41)	(1.43)	(1.59)	(1.46)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)	(0.06)	(0.06)	(0.05)	(0.06)	(0.06)
lgdpea70	0.09	-0.84***	-1.32***	-1.71***	-1.75***	0.06***	0.05***	0.04***	0.04***	0.04***	0.06***	0.05***	0.04***	0.04***	0.04***
	(0.18)	(0.18)	(0.16)	(0.26)	(0.27)	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
sopen		3.10***	2.43***	1.54***	1.59***		0.04***	0.03***	0.02*	0.02**		0.02*	-0.00	-0.00	-0.00
		(0.40)	(0.40)	(0.39)	(0.37)		(0.01)	(0.01)	(0.01)	(0.01)		(0.01)	(0.01)	(0.01)	(0.01)
linv7089			1.35***	1.12***	0.82***			0.02***	0.01	0.02**			0.03***	0.03***	0.03***
			(0.20)	(0.27)	(0.29)			(0.01)	(0.01)	(0.01)			(0.01)	(0.01)	(0.01)
rl				0.33**	0.39***				0.01**	0.00				0.00	0.00
				(0.13)	(0.13)				(0.00)	(0.00)				(0.00)	(0.00)
dtt7090					0.11**					-0.00***					0.00
					(0.06)					(0.00)					(0.00)
Obs.	93	89	89	72	72	85	83	83	71	71	85	83	83	71	71
R-sq.	0.16	0.52	0.67	0.71	0.74	0.77	0.83	0.85	0.87	0.90	0.68	0.69	0.77	0.75	0.77
	(16)	(17)	(18)	(19)	(20)	(21)	(22)	(23)	(24)	(25)	(26)	(27)	(28)	(29)	(30)
snr	-6.02***	-2.81*	-3.55**	-4.29*	-6.44***	-0.16***	-0.12***	-0.12***	-0.15***	-0.13***	0.09	0.11*	0.11**	0.12***	0.11**
	(1.49)	(1.49)	(1.44)	(2.42)	(1.92)	(0.03)	(0.03)	(0.03)	(0.04)	(0.05)	(0.06)	(0.06)	(0.04)	(0.04)	(0.05)
lgdpea70	0.21	-0.68***	-1.08***	-1.45***	-1.51***	0.06***	0.05***	0.05***	0.04***	0.04***	0.06***	0.05***	0.04***	0.04***	0.04***
	(0.16)	(0.19)	(0.18)	(0.31)	(0.31)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)
sopen		3.20***	2.51***	1.66***	1.54***		0.04***	0.03***	0.02	0.02		0.03***	0.01	0.01	0.01
		(0.45)	(0.42)	(0.39)	(0.36)		(0.01)	(0.01)	(0.01)	(0.01)		(0.01)	(0.01)	(0.01)	(0.01)
linv7089			1.25***	1.11***	0.72*			0.01**	0.01	0.02**			0.02***	0.03**	0.03**
			(0.26)	(0.41)	(0.41)			(0.01)	(0.01)	(0.01)			(0.01)	(0.01)	(0.01)
rl				0.33*	0.42**				0.01*	0.00				0.00	0.00
				(0.18)	(0.16)				(0.00)	(0.00)				(0.00)	(0.00)
dtt7090					0.22***					-0.00					0.00
					(0.08)					(0.00)					(0.00)
Obs.	99	93	93	72	72	90	88	88	71	71	90	88	88	71	71
R-sq.	0.13	0.45	0.58	0.57	0.64	0.77	0.81	0.82	0.86	0.86	0.68	0.70	0.74	0.79	0.80

Notes: All estimations by OLS. See section 7 and the revised 1997 version of Sachs and Warner (1995) for explanation of variables. Robust standard errors in parenthesis. ***, **, 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 dollars (World Development Indicators (WDI))

Commodity export price index Commodity export values for 1990 from the UNCTAD Commodity Yearbook 2000 and the UN International Trade Statistics 1993 and 1994. Quarterly world commodity price indices are from the 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 (EIA, 2005) (Column (1) in Tables 5.24 and 6.7). Four price series (coal, plywood, silver, and sorghum) had short gaps in the early sample 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 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.[38] The oil import price index was constructed by interacting the log of the annual average deflated oil price index with a dummy variable that equals 1 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, International Energy Annual 2002). Since these components are expressed in thousands of barrels per day, we multiplied them by 365 times the 2001 average weekly world oil price per barrel, also from EIA.

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

Inflation the log of the annual percentage change in consumer prices, taken from the WDI.

International reserves over GDP International reserves taken from the IFS (series 1..SZF and AA.ZF). GDP in current US dollars, taken from the WDI.

Civil war 1 for civil war, 0 otherwise (Gleditsch,, 2004), weber.ucsd.edu/~kgledits/expwar.html.

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, taken from Banks' Cross-National Time-Series Data Archive (www.scc.rutgers.edu/cnts/about.cfm). Unsuccessful coups are not counted.

Natural disasters number of large geological, climatic, or human disasters (≥ 0.5% of population affected, or damage ≥ 0.5% of GDP, or ≥ 1 death per 10000, IMF (2003)). Data are from the WHO Collaborating Centre for Research on the Epidemiology of Disasters (CRED), available at www.em-dat.net. Geological disasters: earthquakes, landslides, volcano eruptions, tidal waves; Climatic disasters: floods, droughts, extreme temperatures, wind storms; Human disasters: famines, epidemics.

[38] 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.

Appendix B Panel unit root and panel cointegration tests

Using equation (1), the long-run equilibrium equation of log real GDP per capita can be written as follows:

$$Y_{i,t} = -\frac{1}{\lambda}(\gamma_i + \theta_1 t + \beta_1 X_{i,t} + \eta_{i,t}) \qquad (2)$$

where γ_i is a country-specific fixed effect and t is a time trend. Note that both the constant and the coefficient on the time trend are allowed to vary across countries. This follows from the fact that we left the country-specific fixed effect α_i in equation (1) unrestricted. Therefore, it captures a country-specific constant in both the levels and the differenced equation. The constant in the levels equation is represented by γ_i in equation (2). The country-specific constant in the differenced equation implies a country-specific linear time trend in the levels equation (2), which is captured by $\theta_1 t$.[39]

Equation (1) allows us to estimate the long-run relationship in equation (2) if $Y_{i,t}$ and $X_{i,t}$ are cointegrated, i.e. if $Y_{i,t}$ and $X_{i,t}$ have a common stochastic trend which is cancelled out by the linear combination. This implies that i) the individual variables are integrated of order 1, i.e. non-stationary in levels and stationary in first differences, and ii) 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 panel unit root tests by Im, Pesaran and Shin (2003, IPS hereafter) and Maddala and Wu (1999, MW hereafter). Both tests are based on augmented Dickey-Fuller (ADF) tests for the individual series in the panel. This ensures that the ADF test statistic is allowed to vary across groups, unlike for example in the panel unit root test by Levin, Lin and Chu (2002). The null hypothesis is that all groups have a unit root while under the alternative one or more groups do not have a unit root. The IPS test and the MW test differ in that the first test is parametric and is based on the t-statistics of the individual unit root tests, while the second is non-parametric and is based on the p-values of the individual unit root tests.[40] 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.[41] 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 entirely

[39] To limit the number of regressors in the cointegration tests, we leave out the regional time dummies.
[40] 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.
[41] 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.

possible that the tests reject non-stationarity while most of the series are in fact non-stationary. To determine the proportion of countries for which non-stationarity is rejected, we performed (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 (augmented) Dickey-Fuller 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:

$$Y_t = \alpha_0 + \alpha_1 t + \alpha_2 X_t + \varepsilon_{i,t} \qquad (3)$$

where Y_t is log real GDP per capita in year t and X_t includes the long-run GDP determinants (trade openness, log inflation, international reserves over GDP, 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 number of lags of the dependent variable in the ADF regressions to vary across countries by including the lags that enter statistically significant at 10 percent. We then calculate the average of the ADF t-statistics, 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.[42]

[42] Intuitively, "rejection of the null hypothesis means that enough of the individual cross-sections have statistics 'far away' from the means predicted by theory were they to be generated under the null" (Baltagi and Kao, 2000).

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.