Growth, Development and Natural Resources: New Evidence Using a Heterogeneous Panel Analysis*

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Abstract

This paper explores whether natural resource abundance is a curse or a blessing. To do so, we firstly develop a theory consistent econometric model, in which we show that there is a long run relationship between real income, the investment rate, and the real value of oil production. Secondly, we investigate the long-run (level) impacts of natural resource abundance on domestic output as well as the short-run (growth) effects. Thirdly, we explicitly recognize that there is a substantial cross-sectional dependence and cross-country heterogeneity in our sample, which covers 53 oil exporting and importing countries with very different historical and institutional backgrounds, and adopt the non-stationary panel methodologies developed by Pesaran (2006) and Pedroni (2000) for estimation. Our results, using the real value of oil production, rent or reserves as a proxy for resource endowment, reveal that oil abundance has a positive effect on both income levels and economic growth. While we accept that oil rich countries could benefit more from their natural wealth by adopting growth and welfare enhancing policies and institutions, we challenge the common view that oil abundance affects economic growth negatively.

JEL Classifications: C23, O13, O40, Q32.

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1 Introduction

The aim of this paper is to investigate the following questions: Is an abundance of natural resources, in particular oil, a curse or a blessing? What are the effects of natural resource abundance on levels of per capita output and economic growth in general? Following the influential work by Sachs and Warner (1995), a growing empirical literature on and interest in the resource curse paradox was generated. According to this paradox, resource rich countries perform poorly when compared to countries which are not endowed with oil, natural gas, minerals and other non-renewable resources. Therefore, resource abundance is believed to be an important determinant of economic failure, which implies that oil abundance is a curse and not a blessing.

There are different explanations for why resource rich economies might be subject to this curse. Dutch disease (see Corden and Neary (1982), Neary and van Wijnbergen (1986), and Krugman (1987)) is one of the channels through which the resource curse makes itself felt: an increase in natural resource revenue leads to an appreciation of the real exchange rate, which raises the cost (in foreign currency) of exports of the products of other industries, making them less competitive with possible negative effects on economic activity. Economic growth might also be adversely affected by the resulting re-allocation of resources from the high-tech and high-skill manufacturing sector to the low-tech and low-skill natural resource sector. Another explanation for the resource curse paradox is based on rent-seeking theories, which argue that natural resource abundance generates an incentive for economic agents to engage in non-productive activities and for the state to provide fewer public goods than the optimum. See for instance, Lane and Tornell (1996), Leite and Weidmann (1999), Tornell and Lane (1999), and Collier and Hoeffler (2004). Finally, Mehlum et al. (2006) have attempted to show that the impact of natural resources on growth and development depends primarily on institutions, while Boschini et al. (2007) have argued that the type of natural resources possessed is also an important factor. It is not our goal to discuss these theories in detail, or to determine their validity. We refer to Sachs and Warner (1995), Rosser (2006), and Caselli and Cunningham (2009) for an extensive examination of these prominent accounts of the natural resource curse paradox, as well as van der Ploeg and Venables (2009) for a more recent survey.

The empirical evidence on the resource curse paradox is rather mixed. Most papers in the literature tend to follow Sachs and Warner's cross-sectional specification, introducing new explanatory variables for resource dependence/abundance, while others derive theoretical models that are loosely related to their empirical specification. Some of them confirm Sachs and Warner's results (see Rodriguez and Sachs (1999), Gylfason et al. (1999), and Bulte et al. (2005) among others); others tend to shed doubt on the validity of the resource curse paradox (see Alexeev and Conrad (2009), Arezki and van der Ploeg (2007), Cavalcanti et al. (2011), and van der Ploeg and Poelhekke (2010)).

An important drawback in most of these studies is their measure of resource abundance. Sachs and Warner (1995), for instance, use the ratio of primary-product exports to GDP in the initial period as a measure of resource abundance. This ratio, as clearly pointed out by Brunnschweiler and Bulte (2008), measures resource dependence rather than abundance. The latter should be introduced in the growth regressions as the stock or the flow of natural resources. Moreover, a cross sectional growth regression augmented with this regressor clearly

suffers from endogeneity and omitted variable problems. Brunnschweiler and Bulte (2008) argue that the so-called resource curse does not exist, and that while resource dependence, when instrumented in growth regressions, does not affect growth, resource abundance in fact positively affects economic growth. The positive effect of resource abundance on development and growth is also supported by Esfahani et al. (2009), who develop a long run growth model for three major oil exporting economies and derive conditions under which oil revenues are likely to have a lasting impact. However, their approach contrasts with the standard literature on "Dutch disease" and the "resource curse", which primarily focuses on the short run implications of a temporary resource discovery. On the other hand, Stijns (2005), using different measures for resource abundance, indicates that the effect of this variable on growth is ambiguous.

Another branch of the literature investigates the channels through which natural resource abundance affects economic growth negatively. Gylfason (2001), for instance, shows that natural resource abundance appears to crowd out human capital investment with negative effects on the pace of economic activity, while Bravo-Ortega et al. (2005) show that higher educational attainments can in fact offset the negative effects of resource abundance. Therefore, it can be seen that the empirical findings on the resource curse paradox are still not conclusive.

There are a number of grounds on which the econometric evidence of the effects of resource abundance on growth may be questioned. Firstly, the literature relies primarily on a cross-sectional approach to test the resource curse hypothesis, and as such does not take into account the time dimension of the data. As noted above, the cross-sectional approach is also subject to endogeneity problems, and this is perhaps the most important reason for being skeptical about the findings of such econometric studies suggesting a positive or negative association between resource abundance and growth. Secondly, the vast majority of existing studies focus on the effects of resource abundance on the rate of economic growth, even though most growth models in the Solow/Ramsey tradition suggest that the effects on growth should be transitory, though could be permanent for the level of per capita income.[1]

In addition, even when panel data techniques are employed, most studies make use of homogeneous panel data approaches, such as the traditional fixed and random effects estimators, which often apply the instrumental variable (IV) technique proposed by Anderson and Hsiao (1981) and Anderson and Hsiao (1982), and the generalized methods of moments (GMM) estimators of Arellano and Bond (1991), Arellano and Bover (1995), among others.[2] While homogeneous panel data models allow the intercepts to differ across countries all other parameters are constrained to be the same. Therefore, a high degree of homogeneity is still imposed.

As discussed in Pesaran and Smith (1995), the problem with these dynamic panel data techniques, when applied to testing growth effects, is that they can produce inconsistent and

[1] This is also consistent with the empirical evidence provided by Klenow and Rodriguez-Clare (1997). In a different setting, Henry (2007) calls into question the usefulness of the cross-county approach to testing the relationship between capital account liberalization and growth. He argues that the capital deepening channel of gain from financial integration should imply only a temporary, rather than a permanent, increase in growth, but most of the cross-sectional studies that have been conducted do not test this.

[2] For a comprehensive survey of the econometric methods employed in the growth literature, and some of their shortcomings, see Durlauf et al. (2005) and Durlauf et al. (2009).

potentially very misleading estimates of the average values of the parameters, since growth models typically exhibit substantial cross-sectional heterogeneity. In fact Lee et al. (1997), using a panel of data on 102 countries, illustrate that there is pervasive heterogeneity in speeds of convergence and in growth rates across countries and show that the conventional method of imposing homogeneity are subject to substantial biases. In addition, Lee et al. (1998) test the null hypothesis of homogeneity in growth rates as well as the null of common speed of convergence and find that these hypotheses are rejected for 102 non-oil countries. The same test outcomes are obtained for 61 intermediate group of countries, while for 22 OECD countries the null of a common speed of convergence is not rejected.

More recently Pedroni (2007) shows that there are significant differences in the aggregate production functions across countries. By taking into account these differences he argues that it is possible to explain the observed patterns of per capita income divergence across countries. Finally, the current econometric evidence does not address the problem of cross-sectional dependence arising from common unobserved factors or shocks. Thus estimations and inference based on models that do not take into account cross-country heterogeneity and dependence, such as the cross-sectional specifications widely used in the literature, can yield biased and misleading results.[3]

In this paper we adopt a different approach in order to test the resource curse hypothesis. We explicitly recognize that there is a substantial degree of heterogeneity in the growth experience of different resource abundant countries and therefore use a heterogeneous panel data approach, which allows for different dynamics across countries. We also account for cross-country error dependencies that potentially arise from the presence of multiple unobserved common factors, and allow the individual responses to these factors to differ across countries. A possible source of cross-sectional dependency might be due to world-wide common shocks that affect all cross-sectional units, but with different intensities. Changes in technology and in the price of oil provide examples of such common shocks that may affect real GDP per capita, but to different degrees across countries.[4] To address the issues raised above we adopt the common correlated effects estimator of Pesaran (2006), a sufficiently general and flexible econometric approach, which is applicable under both cross-section dependence and cross-country heterogeneity. Moreover, we investigate the long-run (level) effects of natural resource abundance on domestic output as well as the short-run (growth) effects through an Error Correction Model (ECM). We also check the robustness of our long-run results by using the fully modified OLS method of Pedroni (2000). An advantage of our non-stationary panel data approach is that the fixed effects and the heterogeneous trends capture country-specific unobserved factors, such as social and human capital, which are very difficult to measure or observe accurately. In addition, omitted variables that are either constant or evolve smoothly over time are absorbed into the country specific deterministic trend.

Furthermore, we develop a standard theoretical growth model that requires the use of natural resources as an input in the production of the consumption good. We view natural

[3] In fact when we apply a standard cross-sectional estimation technique to our sample of countries, we observe that resource abundance has a negative effect on growth, whereas we find no evidence of an adverse income (growth) effect in any of our regressions when we account for cross-sectional dependencies and heterogeneity, among other things.

[4] Different forms of cross section dependence are discussed and formally defined in Pesaran and Tosetti (2010).

resources as a proxy for energy and power. We assume that agents can extract natural resources at a rate which is optimally determined, and rent them out to firms for production. In contrast to the literature on exhaustible natural resources and economic growth, for instance Dasgupta and Heal (1974), we assume that costly investment will enable new reserves to be found and old fields to be developed. It is important to emphasize that while we do not believe that natural resources are limitless, oil production and more importantly viable reserves do seem to increase over the horizon we are investigating empirically.[5] Finally, our theoretical model suggests a long run relationship between per capita income, the investment rate and the real value of oil production per capita which we use as the basis of our empirical investigation.

In contrast to most studies in the literature, our results, using the real value of oil production, rent and reserves as a proxy for resource abundance, indicate that oil abundance is in fact a blessing and not a curse. Estimating a panel error correction model, we also show that oil abundance has a significantly positive growth effect in the short run.

The plan of this paper is as follows. Section 2 develops the growth theory underlying our econometric model. Section 3 provides a brief review of our panel data model and the estimation methods employed. Section 4 reports the estimation results, and Section 5 provides some concluding remarks.

2 The Theoretical Model

One of the major drawbacks of the empirical literature on growth and natural resource abundance is the lack of an explicit theoretical model. Either an *ad hoc* approach is used, in which output growth is regressed on a number of variables that are arbitrary chosen, or a theoretical model is developed but when it comes to estimation the econometric model is not connected to the theory derived restrictions. To address this drawback, this paper develops a robust, theoretically informed econometric model, in order to test the resource curse paradox.

We suppose that the economy under consideration is populated by a continuum of identical firms of measure one. The representative firm uses physical capital, $K(t)$, labor, $L(t)$, and natural resources, $O(t)$, to produce the consumption good, $Y(t)$, according to the following production function:

$$Y(t) = K(t)^{\alpha_1} O(t)^{\alpha_2} (A(t)L(t))^{1-\alpha_1-\alpha_2}, \ \alpha_1, \alpha_2 > 0, \ \alpha_1 + \alpha_2 < 1, \tag{1}$$

where $A(t) = A(0)e^{gt}$ is the labor augmenting technical progress, and $A(0)$ is an economy-specific initial endowment of technology. The production function exhibits constant returns to scale. We could also assume, as in Romer (1990), that the economy-wide capital stock, $\mathbf{K}(t)$, embodies technology, such that $A(t) = G(\mathbf{K}(t))$. In this case, and at the aggregate level, the economy would exhibit increasing returns to scale. All results derived in this section would hold as long as the externality of the capital stock is not very large.[6] Otherwise, the

[5] If natural resources in our model represent power and energy used in production, we can also view the increase in reserves as discovery of new energy sources.

[6] For instance, if $G(\mathbf{K}(t)) = \mathbf{K}(t)^{\frac{\mu}{1-\alpha_1-\alpha_2}}$, then all results would be the same for $\mu < 1 - \alpha_1 - \alpha_2$.

growth rates would not be stationary in the long run and would grow over time. This, however, is not consistent with the empirical observation or the evidence on conditional convergence (e.g., Klenow and Rodriguez-Clare (1997)).

We could also include human capital as an input in the production process but we abstract from this to simplify the analysis. In fact, in Section 3, we argue that our non-stationary panel approach allows us to capture both human capital, in the form of education, and social capital, in the form of social and political institutions, as these effects are absorbed by the fixed effects and the heterogeneous trends in our specification, as will be set out in equation (15).

Let $r(t)$, $p(t)$, and $w(t)$ denote the rental price of capital, natural resources, and labor, respectively. Competition implies that factors are remunerated according to their marginal productivity, such that:

$$r(t) = \alpha_1 \frac{Y(t)}{K(t)},\ p(t) = \alpha_2 \frac{Y(t)}{O(t)},\ w(t) = (1 - \alpha_1 - \alpha_2)\frac{Y(t)}{L(t)}.$$

Since the production function exhibits constant returns to scale, profits are zero, and firm ownership is not important.

The economy is inhabited by a continuum of identical infinitely lived households with measure one. We can therefore work with a representative household, which grows at the exogenous rate, n, such that: $N(t) = N(0)e^{nt}$, and $N(0)$ is the initial endowment of labour. Each household member is endowed with a unit of productive time as well as $K(0)$, the initial level of capital stock. Capital is rented to firms for production of the consumption good and depreciates at rate δ. Let $I^K(t)$ be investment in physical capital. Therefore:

$$\dot{K}(t) = I^K(t) - \delta K(t).$$

Let $c(t)$ denote real per capita household consumption with preferences defined as

$$\int_0^\infty e^{-\rho t} N(t) u(c(t)) dt,$$

where $\rho > 0$ is the subjective discount rate. The instantaneous utility function is given by:

$$u(c) = \frac{c^{1-\theta} - 1}{1 - \theta},\ \theta > 0,$$

where θ is the coefficient of relative risk aversion. In addition to capital, households are also endowed with a stock of natural resources, $S(t)$, which can be extracted at rate $\gamma(t)$ and rented out to firms for the production of the consumption good. New reserves can be found and old fields can be developed, but this requires investment, $I^S(t)$, such that for each unit of investment households have to pay $\phi\left(\frac{I^S(t)}{S(t)}\right)$ units of output. We assume that this cost is convex, such that $\phi'(\cdot) \geq 0$ and $\phi''(\cdot) \geq 0$, and that $\phi(0) = \phi'(0) = 0$. The evolution of the stock of natural resources is then given by

$$\dot{S}(t) = -\gamma(t)S(t) + I^S(t),$$

where the initial stock of natural resources, $S(0)$, is given. For the sake of notation, we will abstract from the time indicator variable from here on. The problem faced by the representative household is to choose the path of consumption, c, natural resource extraction rate, γ, investment in natural resources, I^S, capital, K, and the stock of natural resources, S, so as to maximize

$$\max \int_0^\infty e^{-\rho t} N u(c) dt \quad \text{subject to} \tag{2}$$

$$\begin{aligned} Nc + \dot{K} + I^S \left(1 + \phi\left(\frac{I^S}{S}\right)\right) &= wN + rK - \delta K + p\gamma S, && (3)\\ \dot{S} &= -\gamma S + I^S, && (4)\\ & c \geq 0,\ \gamma \geq 0,\ I^S > 0. && (5) \end{aligned}$$

Equation (3) corresponds to the household's budget constraint, and (5) states the constraint on the choice variables.

The current value Hamiltonian for this problem is then given by

$$\begin{aligned} H(t; c, I^S, \gamma; K, S; \lambda, \mu) &= e^{-\rho t} N u(c) + \mu[I^S - \gamma S] \\ &\quad + \lambda \left[wN + rK + p\gamma S - \delta K - Nc - I^S \left(1 + \phi\left(\frac{I^S}{S}\right)\right)\right], \end{aligned}$$

where λ and μ are the costate variables for physical capital and the stock of natural resources, respectively.

In equilibrium, we have that $O = \gamma S$, $L = N$, and:

$$\begin{aligned} \lambda = e^{-\rho t} c^{-\theta} \quad &\Rightarrow \quad \frac{\dot{\lambda}}{\lambda} = -\left(\rho + \theta \frac{\dot{c}}{c}\right), && (6)\\ \rho + \theta \frac{\dot{c}}{c} &= \alpha_1 \frac{Y}{K} - \delta, && (7)\\ \lambda \left[1 + \phi\left(\frac{I^S}{S}\right) + \phi'\left(\frac{I^S}{S}\right)\frac{I^S}{S}\right] = \mu \quad &\Rightarrow \quad \lambda q^S = \mu && (8)\\ \lambda p = \mu \quad &\Rightarrow \quad \lambda \alpha_2 \frac{Y}{\gamma S} = \mu, && (9)\\ -\dot{\mu} &= \phi'\left(\frac{I^S}{S}\right)\left(\frac{I^S}{S}\right)^2 \lambda, && (10)\\ Nc + \dot{K} + I^S \left[1 + \phi\left(\frac{I^S}{S}\right)\right] &= Y - \delta K, && (11)\\ \dot{S} &= -\gamma S + I^S, && (12)\\ \lim_{t\to\infty} \lambda K = 0, \quad &\text{and} \quad \lim_{t\to\infty} \mu S = 0. && (13) \end{aligned}$$

Equations (6) and (7) correspond to the change in the shadow price of capital and the traditional Euler equation, respectively. Condition (8) states that the marginal cost of investment in natural resources is equal to its shadow value. Equation (9) is the condition

on the rate of extraction, γ, and equation (10) defines the change in the shadow price of S. Equations (11) and (12) are the resource constraint and the evolution of the stock of natural resources, respectively. Finally, equation (13) defines the transversality conditions: one for physical capital and another for the stock of natural resources.

For any variable X, let us define variable $\tilde{x} = \frac{X}{e^{\left(g_X^*\right)t}}$, in which g_X^* is the growth rate of variable X along the balanced growth path, or the long-run equilibrium.

Proposition 1 *Consider the above growth model and let $\phi'(\cdot) > 0$, $\phi''(\cdot) \geq 0$, and $\phi(0) = \phi'(0) = 0$.*

1. *Then, there exists a non-zero balanced growth path equilibrium with $\tilde{k}^* > 0$, $\tilde{y}^* > 0$, $\tilde{o}^* > 0$, $g_\gamma^* = 0$, $g_K^* = g_Y^* = g_S^* = (n+g)$, and $g_c^* = g$.*

2. *In addition, if p is determined in the world market,[7] then this balanced growth path is saddle-path stable.*

Proof. See Appendix A ■

Part one of proposition 1 suggests that along the balanced growth path, the rate of natural resource extraction is constant,[8] with the growth rate of both output and natural resources per capita equal to the rate of technological progress, g. This implies that contrary to other models with exhaustible natural resources, our model does not imply a long run degenerated level of output. If oil can be imported or exported a similar balanced growth path equilibrium would hold.

Part two of proposition 1 states that if the price of the natural resource is determined in the international market, being exogenously given, then the non-zero balanced growth path equilibrium is saddle-path stable, which implies that there exists a one dimensional stable manifold converging to this long run equilibrium.[9]

Writing the production function (1) in terms of the steady state values of the variables in our model denoted by *:

$$\widetilde{y_t}^* = \left(\tilde{k}_t^*\right)^{\alpha_1} \left(\widetilde{o}_t^*\right)^{\alpha_2},$$

or equivalently:

$$\left(\widetilde{y_t}^*\right)^{1-\alpha_1} = \left(\frac{\tilde{k}_t^*}{\widetilde{y_t}^*}\right)^{\alpha_1} \left(\widetilde{o}_t^*\right)^{\alpha_2},$$

we can use the equation of motion of capital to write the above as:

$$\left(\widetilde{y_t}^*\right)^{1-\alpha_1} = \left(\frac{\left(\frac{I_t^K}{Y_t}\right)^*}{g+n+\delta}\right)^{\alpha_1} \left(\widetilde{o}_t^*\right)^{\alpha_2},$$

[7] Which in fact is quite a plausible assumption.

[8] In a "cake-eating" problem of the optimal depletion of exhaustible resources, Dasgupta and Heal (1974) show that the optimal depletion rate is constant. See also Stiglitz (1974a) and Stiglitz (1974b).

[9] If the price of the natural resource is determined in the domestic market, then it cannot be guaranteed that the balanced growth path equilibrium is saddle path stable.

and taking the natural logarithm we have:

$$(1-\alpha_1)\ln \widetilde{y}_t^{\,*} = \alpha_1 \ln\left(\frac{I_t^K}{Y_t}\right)^* - \alpha_1 \ln(g+n+\delta) + \alpha_2 \ln \widetilde{o}_t^{\,*},$$

or equivalently in per capita terms

$$\begin{aligned}\ln y_t &= \frac{(1-\alpha_1-\alpha_2)}{(1-\alpha_1)}\ln A_0 - \frac{\alpha_1}{(1-\alpha_1)}\ln(g+n+\delta) \\ &\quad + \frac{(1-\alpha_1-\alpha_2)}{(1-\alpha_1)} gt + \frac{\alpha_1}{(1-\alpha_1)}\ln\left(\frac{I_t^K}{Y_t}\right) + \frac{\alpha_2}{(1-\alpha_1)}\ln o_t, \qquad (14)\end{aligned}$$

where lower case letters denote variables in per capita terms. Equation (14) is thus the key empirical reduced form equation, which states that there is an equilibrium relationship between real gross domestic product (GDP) per capita, the share of capital investment in real GDP, and the real value of natural resource (oil) production per capita. In fact this equilibrium relation is consistent with any long-run model in which oil production to income ratio is strictly positive. A similar relationship is also derived in Esfahani et al. (2009) in which they distinguish between the two cases where the growth of oil income, g^0, is less than the natural growth rate (the sum of the population growth, n, and the growth of technical progress, g) and when $g^0 \geq g+n$. Under the former, the effects of oil income on the economy's steady growth rate will vanish eventually, whilst under the latter, oil income enters the long run output equation as it does in our model in equation (14).

3 The Econometric Model and Methodology

Our theoretical model, derived in Section 2, suggested a long run relationship between real gross domestic product (GDP), the investment share of GDP, and oil production. Using equation (14) we can write a panel model for oil producing countries, both net exporters and importers, as an equilibrium relationship compatible with the long-run theory developed:

$$\ln y_{jt} = a_j + d_j t + \beta_{j1}\ln(I/Y)_{jt} + \beta_{j2}\ln o_{jt} + u_{jt} \qquad (15)$$

where $\ln y_{jt}$ is the logarithm of real GDP per capita for countries $j = 1, ..., J$ and time periods $t = 1, ..., T$. Likewise $\ln o_{jt}$ is the logarithm of real value of oil production per capita and $\ln(I/Y)_{jt}$ is log of the investment share of GDP over the same countries and time periods, with a_j denoting country specific fixed effects and $d_j t$ representing heterogeneous country specific deterministic trends. Also remember that the slope coefficients are directly related to shares of capital and oil in output given that: $\beta_{j1} = \frac{\alpha_{j1}}{1-\alpha_{j1}}$ and $\beta_{j2} = \frac{\alpha_{j2}}{1-\alpha_{j1}}$.

Two features of the above long-run relation are worth noting; while in the augmented Solow and Ramsey models the parameters α_1 and α_2 in equation (1) are traditionally taken to be common across all countries, we do not impose this restriction as this is a feature of the model that we wish to investigate. This is also clear from our econometric specification in which the parameter vector of the slope coefficients, $\boldsymbol{\beta}_j = \left(\beta_{j1}, \beta_{j2}\right)'$, is allowed to be heterogeneous across countries. Similarly, we do not impose homogeneity of the depreciation

rate, δ_j, or the growth rates of labour, n_j, and technology, g_j, which is accommodated through the fixed effects and the deterministic trends.

While equation (15) defines the theory derived long-run relation for oil producing countries, the short-run dynamics and their adjustment to the long-run across countries are accommodated through the error term, u_{jt}, which is assumed to have the following multi-factor error structure:

$$u_{jt} = \boldsymbol{\tau}_j' \mathbf{f}_t + \varepsilon_{jt} \tag{16}$$

where $\mathbf{f}_t$ is a vector of unobserved common shocks, which can be stationary or nonstationary (see Kapetanios et al. (2011)) and are allowed to be serially correlated and possibly correlated with the logarithm of the investment share, $\ln(I/Y)_{jt}$, as well as oil production, $\ln(o_{jt})$. The individual-specific errors, ε_{jt}, are allowed to be serially correlated over time and weakly dependent across countries, but are assumed to be distributed independently of both the regressors and the unobserved common factors.

Following Pesaran (2006), assuming a random coefficient model, $\boldsymbol{\beta}_j = \boldsymbol{\beta} + \boldsymbol{\varpi}_j$, where $\boldsymbol{\varpi}_j \sim IID(\mathbf{0}, \mathbf{V}_{\varpi})$, we focus on the estimation of the average value of $\boldsymbol{\beta}_j$, namely $\boldsymbol{\beta}$. To eliminate cross sectional dependence (CD) asymptotically, arising from both strong factors (like oil price shocks or the global financial crisis) and weak factors (such as local spillover effects), we make use of the Common Correlated Effects (CCE) type estimators developed by Pesaran (2006). These estimators augment the OLS regression in equation (15) with the cross-sectional averages of the dependent variable and the regressors – which act as proxies for unobserved common factors. Note that the CCE estimation method allows countries to respond differently to the unobserved shocks. One of the estimators pools observations over the cross sectional units and is called the CCE pooled (CCEP) estimator. If the share of capital in output, α_{j1}, and the share of oil in output, α_{j2}, are the same across countries, thus implying that the individual country slope coefficients, $\boldsymbol{\beta}_j$, are the same, then efficiency gains from pooling observations can be achieved. The other estimator, CCE mean group (CCEMG), additionally allows coefficients of interest to vary across countries and is defined as a simple average of the individual country CCE estimators given by:

$$\widehat{\boldsymbol{\beta}}_{CCEMG} = J^{-1} \sum_{j=1}^{J} \widehat{\boldsymbol{\beta}}_j. \tag{17}$$

Although our econometric specification is simple it is also very general. For instance, as opposed to the traditional cross-sectional and/or homogenous panel approaches in which one needs to find quantifiable variables that can act as proxies for unobserved factors, in our non-stationary panel approach the country specific deterministics, a_j and $d_j t$, capture a broad class of those variables. In addition, the unobserved common components of u_{jt} absorb a number of different factors that drive real output but are at the same time difficult to measure accurately. Changes in technology and in the price of oil as well as local spillover effects provide examples of such (world-wide) common shocks that may affect real GDP per capita, but to different degrees across countries.[10]

[10] Another way of dealing with cross-country dependence is to apply the cross-sectionally augmented Pooled Mean Group (CPMG) approach, see Cavalcanti et al. (2011) for more details. This estimator, just like CCEMG and CCEP, augments the original regression equation with the cross-sectional averages of the regressors and the dependent variable, however, it uses the ML rather than OLS method for estimation.

Moreover, any omitted variables that are either constant or evolve smoothly over time are also absorbed into the country specific fixed effects and the heterogeneous trend components. Furthermore, although our theoretical model does not include human capital, in the form of education, or social capital, in the form of social and political institutions, these unobserved and difficult to measure factors are in fact captured by a_j and $d_j t$ in our cointegrated panel specification, see Pedroni (2007). Finally, another advantage of our non-stationary panel approach is that we explicitly estimate the long-run (low-frequency) relationships among the variables, using annual data rather than trying to take five-year averages to filter out business cycle fluctuations common in the growth literature. This is in contrast to the traditional stationary dynamic and static panel approaches which might inadvertently uncover high-frequency relationships, see Durlauf and Quah (1999). The estimators are also consistent under cointegration and are robust to the omission of variables that are not part of the equilibrium relation defined in equation (14).

4 Empirical Results

To empirically test our theoretical model we obtain annual data from 1980 to 2006 on the logarithm of the real gross domestic product per capita, $\ln y_{jt}$, the logarithm of the investment share of real GDP, $\ln (I/Y)_{jt}$, and the logarithm of the real value of oil production per capita, $\ln o_{jt}$. As we assume, but also prove, in Section 2 that oil production is a constant fraction, γ^*, of the total stock of oil reserves available, we are able to perform a robustness check by estimating our econometric model, given by equation (15), with the real value of oil reserves per capita, $\ln s_{jt}$, instead of $\ln o_{jt}$. As we also have access to data on oil rent for different countries a further robustness check is performed replacing $\ln o_{jt}$ with the real value of oil rent per capita, $\ln or_{jt}$. Our data set covers 53 countries, see Table 9, the number for which oil production and rent data is available. As we do not have oil reserve data for Hungary, we estimate the equation with $\ln s_{jt}$ (also later referred to as specification (c)) using the remaining 52 countries. Out of the 53 countries included in our sample 10 belong to the Organization of the Petroleum Exporting Countries (OPEC), but our sample also includes 17 out of the 30 OECD countries. As such there is a large degree of heterogeneity across countries. These 53 countries together cover 85 percent of world GDP, 77 percent of world oil production per day, and 81 percent of world proven oil reserves. Thus our sample is very comprehensive. A more detailed description of our data and their sources are provided in Appendix B.

4.1 Cross-sectional Estimation

Before we proceed to test our theory-derived econometric model, following Sachs and Warner (1995) we employ a standard cross-sectional estimation technique to investigate the growth effects of resource abundance.[11] Accordingly, we estimate the following equation:

$$g_{y,j} = \varsigma_0 + \varsigma_1 \ln y_{80,j} + \varsigma_2 \overline{I/Y}_j + \varsigma_3 \ln \overline{o}_j + \xi_j, \tag{18}$$

[11] For a recent review of the cross-country growth literature see Eberhardt and Teal (2011).

where $g_{y,j}$ is the average growth rate of the Gross Domestic Product per capita between 1980 and 2006 for country $j = 1..., J$, and $\ln y_{80,j}$ is the logarithm of the initial GDP per capita (in 1980). $\overline{I/Y}$ is the average investment share of GDP and $\ln \overline{o}_j$ is the average of the logarithm of the real value of oil production per capita between 1980 and 2006. We also estimate equation (18) by replacing $\ln \overline{o}_j$, with the average of real oil rent per capita, $\ln \overline{or}_j$, and the average real value of oil reserves per capita, $\ln \overline{s}_j$, between 1980 and 2006. For a detailed description of our data and their sources see Appendix B.

Table 1: Cross-sectional Estimation Results, 1980-2006

	(a)	(b)	(c)
$\ln y_{80,j}$	-0.07	-0.06	-0.06
	(0.149)	(0.148)	(0.148)
$\overline{I/Y}_j$	0.23***	0.23***	0.23***
	(0.049)	(0.048)	(0.049)
$\ln \overline{o}_j$	-0.18**	–	–
	(0.083)		
$\ln \overline{or}_j$	–	-0.18**	–
		(0.083)	
$\ln \overline{s}_j$	–	–	-0.18**
			(0.075)
Observations	53	53	52
$\bar{R}^2$	0.34	0.34	0.35

Notes: Method of estimation is Ordinary Least Squares. The dependent variable is the average growth rate of GDP per capita between 1980 and 2006, $g_{y,j}$. Constants are included in all regressions but not reported. Standard errors are given in parenthesis. Symbols ***, **, and * denote significance at 1%, 5%, and at 10% respectively.

The estimation results of the three cross-sectional specifications are reported in Table 1.[12] While the coefficients of the investment share are significantly positive, the estimated values for the three measures of resource abundance are all statistically significant and negative, thus suggesting that the resource curse is present for the countries in our sample. However, as discussed in the introduction, these cross-sectional estimates as well as the homogenous panel results in the literature are subject to a number of important problems. For example, the regressions assume that the slope coefficients are homogeneous across countries, that the errors are cross-sectionally independent, and that there are no omitted variables. Given these observations, we employ the estimation procedure developed by Pesaran (2006) and applied in Holly et al. (2010) to see whether the resource curse is in fact present in our sample or whether the evidence of the resource curse is due to the limitations of the cross-sectional regressions used. Moreover, the insignificance of the coefficient of $\ln y_{80,j}$ can be associated to the large cross country heterogeneity that exists in our sample of 53 countries. We provide strong evidence for conditional convergence to country-specific steady states in our sample of oil producing countries when we account for slope heterogeneity in the ECM estimations.

[12] We also estimated (18) using data on the levels of the three oil abundance variables and obtained similar results as reported in Table 1, with the coefficient of the oil variables in each specification being significantly negative.

4.2 Panel Unit Root Test Results

Before we proceed with estimation of our model we need to test for cross sectional dependence of the errors and to consider the unit root properties of the variables in our model. It is important to make sure that we do not work with a mixture of $I(1)$ and $I(2)$ variables so that we can make sensible interpretation of the long-run relationships. We start by looking at the CD (Cross-section Dependence) test of Pesaran (2004), which tends to a normal distribution as the number of countries tend to infinity, and is based on the average of the pair-wise correlations of the OLS residuals from the individual regressions of (15) in the panel. Table 2 reports the cross sectional dependence of the residuals from the ADF(p) regressions of the logs of real GDP per capita, investment share of GDP, as well as the real value of oil production, rent, and reserves and their lagged differences, over the period 1980 to 2006 across all of the 53 countries. For each $p = 0, 1, 2,$ and 3, the reported CD statistics are highly significant, with the three oil related variables displaying very large test statistics. The presence of the cross sectional dependence implies that the use of standard panel unit root tests, such as the test proposed by Im et al. (2003), from now on the IPS test, are not valid.

Table 2: CD Test Statistics of ADF(p) Regressions

Variable	ADF(0)	ADF(1)	ADF(2)	ADF(3)
(a) With an Intercept				
$\ln y_{jt}$	12.32	12.11	10.25	8.93
$\ln(I/Y)_{jt}$	3.48	3.16	3.17	2.86
$\ln o_{jt}$	131.72	131.23	129.80	129.10
$\ln or_{jt}$	140.50	140.06	138.87	138.28
$\ln s_{jt}$	100.90	100	99.21	96.90
$\Delta \ln y_{jt}$	10.49	9.73	8.12	8.37
$\Delta \ln(I/Y)_{jt}$	3.89	3.62	3.21	3.60
$\Delta \ln o_{jt}$	137.49	135.20	133.76	132.10
$\Delta \ln or_{jt}$	146.59	144.33	143.12	141.30
$\Delta \ln s_{jt}$	103.68	103.49	101.14	100.79
(b) With an Intercept and a Linear Trend				
$\ln y_{jt}$	10.52	9.85	8.63	8.36
$\ln(I/Y)_{jt}$	3.47	3.32	3.39	3.06
$\ln o_{jt}$	135.42	133.64	129.79	128.34
$\ln or_{jt}$	144.70	143.10	139.68	138.54
$\ln s_{jt}$	101.92	100.36	95.89	93.61

Notes: p^{th}-order Augmented Dickey-Fuller test statistics, ADF(p), for $\ln y_{jt}$, $\ln(I/Y)_{jt}$, $\ln o_{jt}$, $\ln or_{jt}$ and $\ln s_{jt}$ are computed for each cross section unit separately in two cases (a) with an intercept only and (b) with an intercept and a linear time trend. $CD = \sqrt{2T/J(J-1)} \sum_{j=1}^{J-1} \sum_{k=j+1}^{J} \hat{\rho}_{jk}$, with $\hat{\rho}_{jk}$ being the correlation coefficient of the ADF(p) regression residuals between j^{th} and k^{th} cross section units, tends to $N(0,1)$ under the null hypothesis of no error cross section dependence. For more details see Pesaran (2004).

Given the above results, to perform panel unit root tests we will make use of the cross-sectionally augmented IPS (CIPS) test proposed by Pesaran (2007). This test follows the CCE approach and filters out the cross section dependence by augmenting the ADF regressions carried out separately for each country with cross section averages.[13] The cross sectional augmented ADF (CADF) statistics are reported in Table 3 for different lag orders, from which it is clear that for all of our variables including investment shares,[14] at the log-level and with or without a trend, the unit root hypothesis cannot be rejected at the 5 percent level, and for most variables not even at the 10 percent level. On the other hand the unit root hypothesis is clearly rejected when applied to the first differences of these variables. Thus we can safely regard all the variables as $I(1)$, remaining confident that there is not a mixture of $I(1)$ and $I(2)$ variables in our model.

Table 3: Pesaran's CIPS Panel Unit Root Test Results

Variable	CADF(0)	CADF(1)	CADF(2)	CADF(3)
(a) With an Intercept				
$\ln y_{jt}$	-1.808	-2.109*	-1.969	-1.914
$\ln(I/Y)_{jt}$	-1.668	-2.042*	-1.876	-1.686
$\ln o_{jt}$	-1.549	-1.640	-1.624	-1.580
$\ln or_{jt}$	-1.532	-1.587	-1.576	-1.517
$\ln s_{jt}$	-1.905	-1.837	-1.542	-1.576
$\Delta \ln y_{jt}$	-3.426***	-2.808***	-2.39***	-2.493***
$\Delta \ln(I/Y)_{jt}$	-3.925***	-3.226***	-2.828***	-2.551***
$\Delta \ln o_{jt}$	-4.231***	-2.999***	-2.700***	-2.125**
$\Delta \ln or_{jt}$	-4.282***	-3.020***	-2.726***	-2.151**
$\Delta \ln s_{jt}$	-4.808***	-3.424***	-2.512***	-2.022
(b) With an Intercept and a Linear Trend				
$\ln y_{jt}$	-1.949	-2.561*	-2.514	-2.581*
$\ln(I/Y)_{jt}$	-1.827	-2.332	-2.215	-2.071
$\ln o_{jt}$	-2.127	-1.994	-2.093	-2.028
$\ln or_{jt}$	-2.159	-1.991	-2.108	-2.021
$\ln s_{jt}$	-2.528	-2.281	-1.932	-1.791

Notes: The reported values are CIPS(p) statistics, which are cross section averages of Cross-sectionally Augmented Dickey-Fuller (CADF(p)) test statistics, for more details see Pesaran (2007). The relevant lower 1, 5, and 10 percent critical values for the CIPS statistics are –2.23, -2.11, and -2.04 with an intercept case, and –2.73, -2.61, and -2.54 with an intercept and a linear trend case, respectively. Symbols denote *10%, **5%, ***1% rejections.

[13] While the CIPS test allows for heterogeneous unit root processes, it assumes one unobserved factor.

[14] As Pedroni (2007) notes the investment to income ratio can only be locally nonstationary. It is naturally bounded as a ratio between zero and one, but for purposes of estimation in a dynamic cointegrating panel, local nonstationarity is a helpful property.

4.3 Panel Level Effects

Having established that all of our variables are $I(1)$, we proceed by estimating the following equation:

$$\ln y_{jt} = a_j + d_j t + \beta_{j1} \ln(I/Y)_{jt} + \beta_{j2} \ln o_{jt} + u_{jt} \tag{19}$$

which we label as (a). But as previously discussed we also estimate the above equation by replacing the per capita real value of oil production, $\ln o_{jt}$, with (b) the per capita value of oil rent, $\ln or_{jt}$, and (c) the value of oil reserves per capita, $\ln s_{jt}$. The results for the three specifications, (a) to (c), are shown in Table 4. It is clear that the coefficient of oil in all of our specifications is significantly positive and thus in line with our theoretical model, implying that oil abundance leads to a positive level effect. The first three columns report the mean group (MG) estimates, which assume that the errors are cross-sectionally independent. While the estimates suggest similar coefficients on the investment share, those of the real value of oil production and rent are considerably smaller compared to the common correlated effects mean group (CCEMG) estimates, although not far from the common correlated effects pooled (CCEP) estimates. Not surprisingly there is also evidence of cross sectional dependence for the MG estimation errors. For the CCEP and CCEMG estimations we augment equation (19) with the simple cross sectional averages of all of our regressors and the dependant variable. From the CD test statistics it is clear that this augmentation has led to reduction of cross sectional dependence, to such extent that we cannot reject at the 10 percent level the null of no cross sectional dependence for either of the two CCE type estimators. The last three columns report the CCEP estimates which have smaller coefficients on all of the variables as compared to the MG estimates. Finally, the CCEMG estimates have significantly larger coefficients on both the investment share of output and the oil variables as compared to the CCEP estimates. We argue that while the countries in our data set all produce oil, there are substantial heterogeneity among them: some countries are net exporters while others are net importers of oil; some are developed others are developing; in addition, they have different geographical locations. Given this level of heterogeneity across countries we focus on the results of the CCEMG estimates which are reported in the middle panel of Table 4.[15]

The theoretical model in equation (14) provided two expressions for the estimated share of capital in output, $\widehat{\alpha}_1 = \frac{\widehat{\beta}_1}{(1+\widehat{\beta}_1)}$, and the share of oil in output, $\widehat{\alpha}_2 = \widehat{\beta}_2 (1 - \widehat{\alpha}_1)$. As expected, these shares vary depending on which oil variable we use in our analysis, however, it is clear from Table 5 that for the full sample in all cases $\widehat{\alpha}_1 > \widehat{\alpha}_2$, and their sum is about one-third. While the estimated shares of capital and oil are very similar using oil production, (a), and oil rent, (b), the values obtained using oil reserves, (c), are smaller for both α_1 and α_2. However, as previously mentioned no matter how the oil variable is measured the results indicate that the effect of oil on GDP is significantly positive.

To make sure that these results are not driven by a few countries with large coefficients on the oil variables, we look at the individual country CCEMG estimates for each of the three specifications considered.[16] For the full sample, the coefficients of the investment share and the three oil variables all lie within an expected range. There are only eight countries

[15] We also estimated all the regressions without the country specific trends and obtained very similar results to those reported in Tables 4 and 8. These are not reported here but are available upon request.

[16] The results for the individual countries are not reported in the paper, but they are available upon request.

Table 4: Estimation Results, 1980-2006

	MG			CCEMG			CCEP		
	(a)	(b)	(c)	(a)	(b)	(c)	(a)	(b)	(c)
$\ln(I/Y)_{jt}$	0.21^{***} (0.024)	0.22^{***} (0.024)	0.22^{***} (0.025)	0.21^{***} (0.023)	0.21^{***} (0.023)	0.19^{***} (0.024)	0.15^{***} (0.022)	0.15^{***} (0.021)	0.15^{***} (0.023)
$\ln o_{jt}$	0.06^{***} (0.015)	–	–	0.15^{***} (0.031)	–	–	0.06^{***} (0.014)	–	–
$\ln or_{jt}$	–	0.05^{***} (0.013)	–		0.14^{***} (0.033)	–		0.06^{***} (0.014)	–
$\ln s_{jt}$	–	–	0.05^{***} (0.014)		–	0.04^{*} (0.021)		–	0.01^{*} (0.007)
CD Test Statistic	3.23^{***}	3.35^{***}	2.43^{***}	1.59	1.31	1.31	-1.70	-1.65	-1.67

Notes: MG stands for Mean Group estimates and CCEMG and CCEP denote the Common Correlated Effects Mean Group and Pooled estimates respectively. The dependent variable is the logarithm of output per capita, $\ln y_{jt}$. Standard errors are given in parenthesis; for more details see Pesaran (2006). Symbols ***, **, and * denote significance at 1%, 5%, and at 10% respectively.

for which oil production has a negative effect on real output and nine countries for which oil rent has the same effect. However, this effect is only significantly negative for five countries (Chile, France, Netherlands, New Zealand, and Thailand) all of whom are net importers of oil. Thus even at the individual country level there is no evidence that oil abundance, as measured by oil production and rent values, stunts development. On the other hand using oil reserves we find a significant negative effect of oil abundance on 11 out of the 53 countries in our sample;[17] of which only five are net exporters of oil. However, as only five out of the 30 oil exporters in our sample have a negative coefficient on oil reserves in the CCEMG estimations, overall the results of the individual country estimates suggest that oil abundance does not stunt development, thus echoing the results obtained using oil production and rent values.

We split the sample into three subsets: net oil exporting countries (EX), countries that are members of the Organization of the Petroleum Exporting Countries (OPEC), and countries belonging to the Organization for Economic Cooperation and Development (OECD).[18] Re-estimating specifications (a) to (c), using these subsets, we report in Table 5 the CCEMG estimates of the shares of oil and capital in output. For all three subsets, just like in the full sample, the estimates for $\widehat{\alpha}_1$ and $\widehat{\alpha}_2$, have the correct signs and are very similar when considering the specifications with oil rent and oil production. In addition, while $\widehat{\alpha}_1$ is significantly positive for all countries in the subsets, $\widehat{\alpha}_2$ is positive but only significant for the oil exporters and the OPEC countries.

The estimates using oil reserve data, (c), are however, significantly smaller for both the

[17] Bahrain, China, Democratic Republic of the Congo, Denmark, Egypt, France, India, Iran, Italy, Japan, and Thailand.

[18] Some countries belong to more than one subset.

Table 5: Estimates of Capital and Oil Shares based on CCEMG Regressions

	(a)				(b)				(c)			
	All	EX	OPEC	OECD	All	EX	OPEC	OECD	All	EX	OPEC	OECD
$\widehat{\alpha}_1$	0.172	0.129	0.138	0.214	0.171	0.128	0.134	0.213	0.162	0.115	0.089†	0.211
$\widehat{\alpha}_2$	0.123	0.183	0.210	0.000†	0.116	0.184	0.218	0.002†	0.031	0.052	0.037†	-0.01†
$\widehat{\alpha}_1 + \widehat{\alpha}_2$	0.295	0.312	0.348	0.214	0.287	0.312	0.352	0.215	0.193	0.167	0.127†	0.201

Notes: † Symbol denotes that the coefficient is not significant in CCEMG regressions. (a) is estimated by augmenting (4) with the simple cross sectional averages of the regressors using $\ln o_{jt}$, whereas (b) and (c) are estimated in the same way but by using $\ln or_{jt}$ and $\ln s_{jt}$ respectively. EX refers to the 30 net oil exporting countries in our sample, OPEC to the Organization of the Petroleum Exporting countries, and OECD to the Organization for Economic Cooperation and Development, for further details see Table 9.

shares of capital and oil. While both $\widehat{\alpha}_1$ and $\widehat{\alpha}_2$ are still significantly positive for the oil exporting countries, they are positive but insignificant for the OPEC subset. As oil reserves are defined as "quantities of oil that geological and engineering information indicate with reasonable certainty can be recovered in the future from known reservoirs under existing economic and operating conditions" (see British Petroleum (2010)) they are subject to large margin of errors and so could be unreliable. Moreover, we argue that the flow measures are better indicators of abundance as they portrait a country's ability to extract its stock and make use of the proceeds. Therefore, we will mainly depend on the results using (a) and (b).

Focusing on the oil rent and production specifications, we can see that while for the OECD countries $\widehat{\alpha}_1 > \widehat{\alpha}_2$, this is not the case for the OPEC and EX countries for which $\widehat{\alpha}_1 < \widehat{\alpha}_2$, with the value of the shares being very similar when comparing the two subsets. These results are perhaps expected for these countries since the share of oil in output is quite significant as oil production dominates economic activity for these countries. Notice also that the sum of $\widehat{\alpha}_1$ and $\widehat{\alpha}_2$ are reasonable for these two subsets being around one-third. Our estimates for the share of capital in output falls in the estimated range of 0.12 to 0.18 for the non-OECD countries and 0.17 to 0.22 for the OECD countries as provided by Pedroni (2007). In addition, $\widehat{\alpha}_1$ for the oil exporting countries is close to the estimates for Iran, Norway and Saudi Arabia as provided in Esfahani et al. (2009). It is also interesting to note that for the OECD countries as a group the share of oil in output is not significantly different from zero. Again we would have expected $\widehat{\alpha}_2$ to be very small for these countries given that they tend to be non-oil exporters.

To make sure that the results are not driven by a few outliers in the sub-samples, we look at country-specific estimations for the OPEC and EX countries. Overall the coefficients of the oil variables and the share of investment in output are reasonable. The United Arab Emirates is the only country in OPEC for which oil rent and production have a negative effect on income, but this is in fact insignificant. However, using oil reserves in our estimation, we find a significant negative effect of oil abundance on income for Iran and Nigeria. Turning to the results for the EX countries we find that there are no countries for which the coefficient of oil rent and production is significantly negative. On the other hand, using oil reserves we

find that there are five countries (Bahrain, Democratic Republic of the Congo, Egypt, Iran and Nigeria) for which oil abundance has a negative effect on income. But as this is only five countries out of the 30 countries considered, the results do not seem to indicate that resource abundance harms development. Thus we can conclude that the estimates for the OPEC and the oil exporting countries do not seem to be affected by outliers and suggest that, in the long run, oil abundance is in fact a blessing and not a curse.

4.3.1 Robustness Check with FMOLS Approach

To check the robustness of our results we also estimate our model using Pedroni's group mean fully modified OLS (FMOLS) estimator. This methodology addresses the problem of simultaneity bias in a non-stationary static panel setting. In particular, a semi-parametric correction is made to the OLS estimator to eliminate the second order bias caused by endogenous regressors. Since our data are non-stationary, this means that by using Pedroni's FMOLS for heterogeneous cointegrated panels, we can potentially exploit the superconsistency properties of cointegrating parameters to address biases coming from endogeneity and omitted variables while at the same time partly allowing for cross-sectional dependence through common time effects.[19] The FMOLS group mean estimates and the corresponding standard errors are reported in Table 6. All of the coefficients are correctly signed with those of the real value of oil production and oil rent per capita statistically significant and positive. Overall, Table 6 confirms the robustness of our previous results and provides evidence of the positive level effects of oil abundance on real output.

Table 6: Group Mean FMOLS Estimation Results, 1980-2006

	(a)	(b)	(c)
$\ln(I/Y)_{jt}$	0.33^{***} (0.020)	0.32^{***} (0.019)	0.28^{***} (0.020)
$\ln o_{jt}$	0.12^{***} (0.014)	–	–
$\ln or_{jt}$	–	0.11^{***} (0.015)	–
$\ln s_{jt}$	–	–	0.02 (0.500)
$\widehat{\alpha}_1$	0.248	0.242	0.219
$\widehat{\alpha}_2$	0.090	0.083	0.016
$\widehat{\alpha}_1 + \widehat{\alpha}_2$	0.338	0.325	0.235

Notes: The estimates are based on the Pedroni (2000) group mean FMOLS estimator. The dependent variable is the logarithm of output per capita, $\ln y_{jt}$. Standard errors are given in parenthesis. Symbols ***, **, and * denote significance at 1%, 5%, and at 10% respectively.

The implied estimates for the shares of capital and oil in output are also reported in Table

[19] The FMOLS approach is based on subtracting the common factors prior to estimation rather than explicitly including them as additional covariates in the model. Note that the FMOLS estimator can deal with cross-section dependence only if the common factors are $I(0)$. For further details see Pedroni (2000).

6. As before the share of capital is greater than that of oil, and the total shares, $\widehat{\alpha}_1 + \widehat{\alpha}_2$, are estimated to be around one-third. Comparing these results with the shares computed using the CCEMG estimator in Table 5, we observe that generally the share of capital is larger and the share of oil is slightly smaller using the group mean FMOLS estimator.

4.4 Panel Cointegration Test Results

We use the residuals, $\widehat{e}_{jt}$, obtained from the CCEMG estimation of

$$\ln y_{jt} = a_j + d_j t + \beta_{j1} \ln(I/Y)_{jt} + \beta_{j2} \ln o_{jt} + b_{j0} \ln \overline{y_t} + b_{j1} \ln(\overline{I/Y})_t + b_{j2} \ln \overline{o}_t + e_{jt} \quad (20)$$

to test the null of no cointegration between real GDP per capita, the investment share of GDP, and each of the three different measures for oil, (a) to (c). The CCEMG estimation procedure applied in equation (20) asymptotically eliminates both weak as well as strong forms of cross sectional dependence in large panels. Thus our cointegration test is based on the IPS test procedure, as our goal is to determine whether the residuals, $\widehat{e}_{jt}$, contain a unit root or not. The panel cointegration test results are displayed in Table 7a and suggest rejection of the null hypothesis of no cointegration for all three specifications even at the one percent level and for all the augmentation orders, $p = 0, 1, 2,$ and 3.

Table 7: Panel Cointegration Test Results

(a) IPS test on residuals of CCEMG estimations

	ADF(0)	ADF(1)	ADF(2)	ADF(3)
(a)	-3.179***	-3.063***	-2.790***	-2.688***
(b)	-3.130***	-3.034***	-2.742***	-2.659***
(c)	-3.193***	-3.222***	-2.966***	-2.737***

(b) Pedroni's test on residuals of MG estimations

	Raw data			Demeaned data		
	ρ-statistic	PP	ADF	ρ-statistic	PP	ADF
(a)	2.94***	-2.05**	-4.65***	2.94***	-1.90*	-2.82***
(b)	2.79***	-2.34**	-5.10***	2.93***	-1.74*	-2.82***
(c)	2.95***	-1.94**	-3.80***	3.49***	-0.98	-3.04***

Notes: ρ-statistic, PP, and ADF columns report the Pedroni (1999, 2004) group mean tests for null of no cointegration. Fixed effects and heterogeneous trends have been included in all cases. Symbols denote *10%, **5%, ***1% rejections.

To check the robustness of our results we also apply Pedroni's (1999, 2004) tests for the null hypothesis of no cointegration to the residuals, $\widehat{u}_{jt}$ from the hypothesized cointegrating relationship in equation (19), but also using (b) oil rent and (c) reserves measures. The three panel cointegration test statistics based on a group mean approach are reported in

Table 7b. The first one is analogous to the Phillips and Perron ρ-statistic, and the other two are analogous to the Phillips and Perron t-statistic (non parametric) and the augmented Dickey-Fuller t-statistic (parametric).[20] Looking at the results for the raw data, as reported in the first three columns, we see that all tests reject the null of no cointegration at the 5 percent level and in most cases even at the one percent level. When using data that have been demeaned with respect to the cross-sectional dimension for each time period,[21] we reject the null at the one percent level using the ADF and ρ-statistics. The PP test statistics on the other hand suggest rejection of the null at the 10 percent level using (a) oil production and (b) rent data, but cannot reject the null using (c) oil reserves data. Thus in general the results, based on both the IPS and the Pedroni tests, suggest the existence of cointegrating relationships among the variables in our model for all three specifications, (a) to (c).

4.5 Panel Error Correction Specifications

Having established panel cointegration between real GDP per capita, the share of investment in real GDP, and the real value of oil production per capita (as well as real value of oil reserves and real oil rent per capita), we now estimate the following panel error correction model

$$\begin{aligned} \Delta \ln y_{jt} &= e_j + \psi_j \left[\ln y_{j,t-1} - \chi_{j1} \ln (I/Y)_{j,t-1} - \chi_{j2} \ln o_{j,t-1} \right] \\ &\quad + \kappa_{j1} \Delta \ln (I/Y)_{jt} + \kappa_{j2} \Delta \ln o_{jt} + \kappa_{j3} \Delta \ln y_{jt-1} + \upsilon_{jt} \end{aligned} \tag{21}$$

to determine the short-run and the long-run effects of oil on real GDP per capita. To investigate these effects, firstly we need to consider whether ψ_j, the coefficients on $\ln y_{j,t-1}$, are statistically different from zero. If this is not the case, the cointegration results would not be reliable. Secondly, we need to test the null hypothesis that the parameter vector of the short-run response coefficients, $\boldsymbol{\kappa}_j = (\kappa_{j1}, \kappa_{2j}, \kappa_{j3})'$, are equal to zero. If the null cannot be rejected, then there would be no evidence for short run dynamics. Table 8 displays the results from estimating the above equation. As before there is considerable cross sectional dependence in the MG regressions, see the first three columns. To address this issue, we computed CCEMG and CCEP estimates by augmenting equation (21) with simple cross sectional averages of the regressors. To check the robustness of our estimates we also estimated equation (21) by replacing $\ln o_{jt}$ with $\ln or_{jt}$ and $\ln s_{jt}$. The coefficients on $\ln y_{j,t-1}$, in all specifications are statistically significant and different from zero indicating that the system reverts to the long-run values following a shock. All other estimated coefficients are correctly signed with the coefficients of the real value of oil production and rent statistically significant and positive in both the short-run and the long-run, indicating that oil abundance has both positive level and growth effects. These findings shed doubts on the results based on the pure cross-sectional specifications usually employed in the literature, as reported in Section 4.1, and suggest that the cross-sectional estimates could be misleading.

When estimating equation (21) using the real value of oil reserves instead of production, while the MG and the CCEMG estimates show statistically significant and positive coefficients for $\ln s_{jt-1}$ and $\Delta \ln s_{jt}$, the CCEP estimates although being positive are insignificant.

[20] For a discussion and mathematical expositions of these statistics see Pedroni (1999) and Pedroni (2004).

[21] The demeaned version serves to extract common time effects from the data and the results can be interpreted as accounting for certain forms of cross-sectional dependency.

Table 8: Panel Error Correction Estimates, 1980-2006

	MG			CCEMG			CCEP		
	(a)	(b)	(c)	(a)	(b)	(c)	(a)	(b)	(c)
$\ln y_{j,t-1}$	-0.39*** (0.025)	-0.39*** (0.025)	-0.37*** (0.026)	-0.61*** (0.047)	-0.59*** (0.044)	-0.55*** (0.053)	-0.32*** (0.026)	-0.32*** (0.026)	-0.35*** (0.033)
$\ln(I/Y)_{j,t-1}$	0.07*** (0.012)	0.08*** (0.013)	0.08*** (0.015)	0.13*** (0.021)	0.13*** (0.021)	0.09*** (0.026)	0.07*** (0.013)	0.07*** (0.013)	0.07*** (0.015)
$\ln o_{j,t-1}$	0.04*** (0.010)	–	–	0.10*** (0.027)	–	–	0.03*** (0.005)	–	–
$\Delta \ln o_{jt}$	0.04*** (0.011)	–	–	0.11*** (0.024)	–	–	0.03** (0.015)	–	–
$\ln or_{j,t-1}$	–	0.03*** (0.008)	–	–	0.08*** (0.025)	–	–	0.03*** (0.005)	–
$\Delta \ln or_{jt}$	–	0.03*** (0.010)	–	–	0.11*** (0.025)	–	–	0.03** (0.015)	–
$\ln s_{j,t-1}$	–	–	0.02*** (0.006)	–	–	0.04** (0.017)	–	–	0.002 (0.004)
$\Delta \ln s_{jt}$	–	–	0.02*** (0.006)	–	–	0.02* (0.012)	–	–	0.002 (0.003)
$\Delta \ln(I/Y)_{jt}$	0.13*** (0.019)	0.13*** (0.019)	0.14*** (0.022)	0.15*** (0.021)	0.15*** (0.021)	0.12*** (0.025)	0.07*** (0.011)	0.07*** (0.011)	0.08*** (0.012)
$\Delta \ln y_{j,t-1}$	0.17*** (0.032)	0.17*** (0.032)	0.13*** (0.035)	0.09** (0.038)	0.08** (0.038)	0.12*** (0.044)	0.10*** (0.050)	0.11** (0.051)	0.11** (0.052)
CD Test Statistics	4.97***	5.09***	4.04***	-0.84	-0.90	-1.51	-1.40	-1.39	-0.49
$\widehat{\alpha}_1$	0.158	0.164	0.181	0.177	0.177	0.147	0.178	0.178	0.173
$\widehat{\alpha}_2$	0.083	0.069	0.052	0.132	0.115	0.055	0.069	0.071	0.004
$\widehat{\alpha}_1 + \widehat{\alpha}_2$	0.241	0.233	0.233	0.309	0.292	0.202	0.247	0.249	0.177

Notes: The country specific intercepts are estimated but not reported. MG stands for Mean Group estimates while CCEMG and CCEP denote the Common Correlated Effects Mean Group and Pooled estimates respectively. The dependent variable is the change in the logarithm of output per capita, $\Delta \ln y_{jt}$. Standard errors are given in parenthesis. Symbols ***, **, and * denote significance at 1%, 5%, and at 10% respectively.

As we believe that there is considerable heterogeneity and cross sectional dependence across countries in our sample, we focus on the CCEMG estimates, which generally has significantly larger coefficients on both the investment share of output and the oil variables considered as compared to the MG and CCEP estimates. The estimated share of capital in output, $\widehat{\alpha}_1$, and the share of oil in output, $\widehat{\alpha}_2$, vary depending on which oil variable we use in our analysis, but are very close and in line with those reported in Table 5. As before the share of capital in output is larger than the share of oil in output and they sum to less than one-third. The coefficients of the short-run parameters suggest an elasticity of real income with respect to both production and rent per capita of around 11 percent, with the reserve elasticity of income at two percent. Our results then seem to confirm that oil abundance has both a positive level (long-run) as well as growth (short-run) effects.

To check the robustness of our results to the choice of natural resource considered, we performed the same estimations with the three measures of (a) real value of production, (b) rent, and (c) reserves per capita but by using natural gas as well as combining natural gas

and oil data and obtained very similar results to the ones reported in Tables 4 and 8. For the sake of space, these results are not reported but are available upon request.

5 Concluding Remarks

This paper has re-visited the resource curse paradox in a panel made up of 53 countries over 27 years. The sample covered 85 percent of world GDP, 77 percent of world oil production, and 81 percent of world proven oil reserves. Not surprisingly, there is substantial degree of cross country heterogeneity. We started off by developing a theory consistent econometric model, which suggested a long run relationship between the variables in our model. We then employed the Common Correlated Effects type estimators developed in Pesaran (2006) to test whether natural resource abundance is a curse or a blessing, and also contrasted these results with those of the FMOLS approach of Pedroni (2000).

Using the Cross-section Dependence (CD) test of Pesaran (2004) we were able to establish that the errors of the variables in our model in fact exhibit a considerable degree of cross sectional dependence. Due to the presence of this cross sectional dependence, we employed Pesaran's CIPS test to determine whether these variables are non-stationary or not. Having established that they are, we tested the effects of the real value of: (a) oil production, (b) rent, and (c) reserve on real income. We mainly relied on the CCEMG estimates as they explicitly take into account both cross sectional dependence and cross country heterogeneity among the countries in our sample. The results suggested that the effect of oil abundance on real income is significantly positive. In addition, the CCEMG estimates suggested that the share of capital in output, $\widehat{\alpha}_1$, is larger than the share of oil in output, $\widehat{\alpha}_2$, with their sum being about one-third. To ensure that the results are not affected by outliers we looked at the individual country estimates and were able to confirm that the coefficients of the investment share and the three oil variables are all sensible and in line with the full sample estimations. As a further robustness check we estimated our model using Pedroni's group mean fully modified OLS (FMOLS) estimator and obtained very similar results. In addition, we performed cointegration tests, using the IPS methodology as well as Pedroni's cointegration tests, and were able to reject the null of no cointegration for all the specifications considered, providing empirical support for the theory derived long-run relationship between the variables in our model.

We also estimated separate models for oil exporters as a whole, the OPEC member countries, as well as the OECD countries. The results confirmed that in all three sets of country groupings, oil abundance has a positive and statistically significant effect on real income. It is interesting to note that while $\widehat{\alpha}_2 \approx 0$ for the OECD countries, $\widehat{\alpha}_2 > 0$ and statistically significant for the oil exporting countries. We would expect this result for the oil exporting countries (both the EX and the OPEC subsets), as the share of oil in real GDP is very large because oil production dominates economic activity for these countries.

To determine the short-run effects of oil abundance, we estimated a panel error correction model. All of the estimated oil variable coefficients, using (a) to (c), are positively signed and statistically significant, thus indicating that oil abundance has short-run growth enhancing effects. Moreover, the shares of oil and capital in output are in line with what is expected and the elasticity of oil in income is around 11 percent when using oil production and rent

as a proxy for resource abundance, while only being around two percent when using oil reserves. In general, the estimates using oil reserves, although statistically significant and positive, are weaker that the ones using oil production and rent. We argue that the flow measures are better indicators of abundance as they portrait a country's ability to extract its stock and make use of the proceeds. Our results suggest that oil abundance has both short-run growth enhancing as well as positive level effects on real income. Therefore, oil abundance by itself does not seem to be a curse. Thus, in light of our results, we believe that the question should not be whether having a large endowment of natural resources is bad or good for an economy, but instead focus should be placed on how these resource abundant economies could be made better off by adopting growth and welfare enhancing policies and institutions.

Appendix A: Proof of Proposition 1

We consider a particular equilibrium path in which the control, state and costate variables grow at a constant rate, and we will show that this equilibrium exists. Using the Euler equation (7), and given that along the balanced growth path the growth rate of consumption is constant, it must be that the growth rate of output is equal to the growth rate of the capital stock: $g_Y^* = g_K^*$. Let's guess that γ^* is constant, such that $g_\gamma^* = 0$, but given this and (12):

$$g_S^* = -\gamma^* + \frac{I^S}{S},$$

$\frac{I^S}{S}$ must then be constant. This implies that $q^S = 1 + \phi\left(\frac{I^S}{S}\right) + \phi'\left(\frac{I^S}{S}\right)\frac{I^S}{S}$ is constant and from (8) and (6), we have that $g_\lambda^* = g_\mu^* = -\left(\rho + \theta g_c^*\right)$. Next, taking the log derivatives of (9) we get:

$$g_\lambda^* + g_Y^* - g_\gamma^* - g_S^* = g_\mu^* \;\Rightarrow\; g_Y^* - g_\gamma^* - g_S^* = 0,$$

but since $g_\gamma^* = 0$, the above implies that $g_Y^* = g_S^*$. Therefore, from the production function, (1), we have that:

$$g_K^* = g_Y^* = g_S^* = n + g. \tag{22}$$

Rewriting the resource constraint, (11), as:

$$\frac{Nc}{Y} + \frac{\dot{K}}{K}\frac{K}{Y} + \frac{I^S}{S}\frac{S}{Y}\left[1 + \phi\left(\frac{I^S}{S}\right)\right] = 1 - \delta\frac{K}{Y},$$

and given that $\frac{K}{Y}$, $\frac{S}{Y}$, $\frac{I^S}{S}$, and $\frac{\dot{K}}{K}$ are all constants, we have that: $g_c^* = g$.

Now it remains to show that γ^* is constant and positive. From (6), (8), (10) and given that $g_\lambda^* = g_\mu^*$ we have:

$$\left[1 + \phi\left(\frac{I^S}{S}\right) + \phi'\left(\frac{I^S}{S}\right)\frac{I^S}{S}\right](\rho + \theta g) = \phi'\left(\frac{I^S}{S}\right)\left(\frac{I^S}{S}\right)^2.$$

From the above equation we can define $\frac{I^S}{S} = H\left(\rho + \theta g\right)$ as an implicit function, $H(\cdot)$, of $\rho + \theta g$. It has a unique positive solution with $\frac{I^S}{S} > (\rho + \theta g)$. In order to see this, observe that the right hand side of the above equation is an increasing function of $\frac{I^S}{S}$ that crosses the y-axis at the origin. The left hand side is also an increasing function of $\frac{I^S}{S}$, but it crosses the y-axis at $\rho + \theta g$. Moreover, at $\frac{I^S}{S} = (\rho + \theta g)$ the left hand side is larger than the right hand side. Finally, the slope of the left hand side is smaller than the slope of the right hand side for any $\frac{I^S}{S} > (\rho + \theta g)$. Therefore, the left hand side crosses (from above) the right hand side in only one point. Since the transversality condition requires that:

$$\rho + \theta g > n + g,$$

we have that

$$\gamma^* = H(\rho + \theta g) - (n + g) > 0,$$

which verifies our guess. For instance, if we assume that $\phi\left(\frac{I^S}{S}\right) = \left(\frac{I^S}{S}\right)$, then

$$\left(\frac{I^S}{S}\right)^* = (\rho + \theta g) \pm \sqrt{(\rho + \theta g)^2 + (\rho + \theta g)},$$

with a positive and a negative root. The positive root, $\left(\frac{I^S}{S}\right)^*_+$ is larger than $(\rho + \theta g)$. Therefore,

$$\gamma^* = \left(\frac{I^S}{S}\right)^*_+ - (n + g),$$

is a positive constant, as required.

Given that $\phi'(\cdot) > 0$, $\phi''(\cdot) \geq 0$, $\left(\frac{I^S}{S}\right)^*_+ > 0$, and $\phi\left(\left(\frac{I^S}{S}\right)^*_+\right) > 0$, q^S must be positive. Rewriting (9) as:

$$\lambda\alpha_2 \frac{Y^*}{O^*} = \mu \;\Rightarrow \frac{\widetilde{o}^*}{\widetilde{y}^*} = \alpha_2 \frac{\lambda}{\mu} = \frac{\alpha_2}{q^S} > 0, \tag{23}$$

where $\tilde{x}$ is the intensive from of X and defined by $\tilde{x} = \frac{X}{e^{g^*_X t}}$, in which g^*_X is the growth rate of variable X in the balanced growth path, and (7) as:

$$\rho + \theta \frac{\dot{c}}{c} = \alpha_1 \frac{Y^*}{K^*} - \delta \;\Rightarrow \frac{\widetilde{k}^*}{\widetilde{y}^*} = \frac{\alpha_1}{\rho + \theta g + \delta} > 0, \tag{24}$$

and writing the production function (1) in intensive form:

$$\widetilde{y}^* = \left(\frac{\widetilde{k}^*}{\widetilde{y}^*}\right)^{\frac{\alpha_1}{1-\alpha_1-\alpha_2}} \left(\frac{\widetilde{o}^*}{\widetilde{y}^*}\right)^{\frac{\alpha_2}{1-\alpha_1-\alpha_2}},$$

it is obvious that $\widetilde{y}^* > 0$. But this implies that both $\widetilde{o}^*$ and $\widetilde{k}^*$ are positive, and since γ^* is positive, it in turn means that $\widetilde{o}^* = \gamma^* \widetilde{s}^* > 0$. This proves part 1 of proposition 1.

The system of equations that describes the equilibrium dynamics is given by:

$$\dot{\tilde{c}} = \frac{1}{\theta}\left[\alpha_1 \tilde{k}^{\alpha_1 - 1}(\gamma \tilde{s})^{\alpha_2} - (\delta + \rho + \theta g)\right]\tilde{c}, \tag{25}$$

$$\dot{\tilde{k}} = \tilde{k}^{\alpha_1}(\gamma \tilde{s})^{\alpha_2} - (\delta + n + g)\tilde{k} - \tilde{c} - \frac{\tilde{\imath}^S}{\tilde{s}}\left(1 + \phi\left(\frac{\tilde{\imath}^S}{\tilde{s}}\right)\right)\tilde{s}, \tag{26}$$

$$\dot{\tilde{s}} = \left[\frac{\tilde{\imath}^S}{\tilde{s}} - (\gamma + n + g)\right]\tilde{s}, \tag{27}$$

where for any variable $\tilde{x}$, $\dot{\tilde{x}} = \frac{d\tilde{x}}{dt}$. Moreover, $p = q$, which implies that

$$1 + \phi\left(\frac{\tilde{\imath}^S}{\tilde{s}}\right) + \phi'\left(\frac{\tilde{\imath}^S}{\tilde{s}}\right)\frac{\tilde{\imath}^S}{\tilde{s}} = p.$$

When p is determined in the international market, then variable $\tilde{z} = \frac{\tilde{\imath}^S}{\tilde{s}}$ is independent of $\tilde{c}$, $\tilde{k}$, and $\tilde{s}$. Moreover, $\gamma\tilde{s} = \left(\frac{\alpha_2}{p}\right)^{\frac{1}{1-\alpha_2}} \tilde{k}^{\frac{\alpha_1}{1-\alpha_2}}$. Using $\tilde{z}$ and $\gamma\tilde{s}$ in this system and taking the first-order Taylor approximation around the steady-state, we have:

$$\begin{pmatrix} \dot{\tilde{c}} \\ \dot{\tilde{k}} \\ \dot{\tilde{s}} \end{pmatrix} \approx \begin{pmatrix} 0 & -(1-\alpha_1)\frac{(\delta+\rho+\theta g)\tilde{c}^*}{\theta\tilde{k}^*} & 0 \\ -1 & (\delta+\rho+\theta g)-(\delta+n+g) & -\tilde{z}(1+\phi(\tilde{z})) \\ 0 & \frac{\alpha_1}{1-\alpha_2}\frac{\gamma^*\tilde{s}^*}{\tilde{k}^*} & \gamma^* \end{pmatrix} \begin{pmatrix} \widetilde{c}-\widetilde{c}^* \\ \widetilde{k}-\widetilde{k}^* \\ \widetilde{s}-\widetilde{s}^* \end{pmatrix}.$$

The determinant of the matrix of partial derivative is $-(1-\alpha_1)\frac{(\delta+\rho+\theta g)\tilde{c}^*}{\theta\tilde{k}^*}\gamma^* < 0$. The trace of this matrix is $(\rho+\theta g)-(n+g)+\gamma^*$, and it is positive to satisfy the transversality conditions. These two results imply that one eigenvalue is negative, while the other two are positive, which completes the proof.

Appendix B: Sources and Construction of the Data

The panel data set used in this study is balanced and contains annual data from the World Bank (2010b) on the values of oil production and oil rent for all of the 53 countries reported in Table 9, over the period 1980-2006.

Oil reserve[22] data is available for all of the countries in our data set from the Energy Information Administration, see United States Department of Energy (2010). For Hungary reserve data is available only from 1992, as such it is excluded from the estimations with $\ln s_{jt}$. The data on GDP for all countries is obtained from the World Bank (2010a), but is not available for Bahrain (2006), for which we obtain the 2006 GDP figure by applying the growth rate of GDP in 2006 from the Central Bank of Bahrain. GDP data is also missing for Kuwait (1990-1994) so we obtain the missing data by splicing the GDP series from the Penn World Table Version 6.3, see Heston et al. (2009).[23] Finally the World Bank does not have GDP or gross capital formation for Qatar until 2000, so we use the Penn World Table data until then and obtain data for the later years by splicing the World Bank data. Our main source for data on gross fixed capital formation is the World Bank, but data is missing for Argentina (1980-1992), Colombia (2000-2006), and Oman (1983-1989) but available from the International Monetary Fund (2010) International Financial Statistics database and so we make use of that instead. Gross fixed capital formation data is also missing for Brunei (1980-1988) and Romania (1980, 1983-1989), but as this is not available from the IMF, it is obtained by splicing the data from the Penn World Table Version 6.3, see Heston et al. (2009). The nominal values were converted into real terms using the US GDP deflator from World Bank (2010a).

[22] Quantities of oil that geological and engineering information indicate with reasonable certainty can be recovered in the future from known reservoirs under existing economic and operating conditions.

[23] The missing data is retrieved by applying the growth rate of the PWT series to the World Bank data.

Table 9: Countries Included in the Sample

Oil Exporters		OECD	Other Countries
Algeria*	Kuwait*	Australia	Brazil
Argentina	Malaysia	Austria	Chile
Bahrain	Mexico	Canada	China
Bolivia	Nigeria*	Denmark	Cote d'Ivoire
State of Brunei Darussalam	Norway	France	India
Cameroon	Oman	Germany	Israel
Canada	Papua New Guinea	Greece	Morocco
Colombia	Qatar*	Hungary	Peru
Democratic Republic of the Congo	Saudi Arabia*	Italy	Romania
Republic of the Congo	Syria	Japan	Thailand
Ecuador*	Trinidad and Tobago	Mexico	
Egypt	Tunisia	Netherlands	
Gabon	United Arab Emirates*	New Zealand	
Indonesia*	United Kingdom	Norway	
Iran*	Venezuela*	Turkey	
		United Kingdom	
		United States	

Notes: * indicates that the country is a member of the Organization of the Petroleum Exporting Countries (OPEC). OECD refers to the Organization for Economic Cooperation and Development

Table 10: List of Variables and their Sources

Variables and definitions		Sources
y_{jt}	Gross Domestic Product per capita, in constant 2000 US dollars.	Authors' calculation using data from the World Bank (2010a) World Development Indicators (WDI).
$(I/Y)_{jt}$	Ratio of gross fixed capital formation to Gross Domestic Product.	
$d_{US,t}$	US GDP Deflator, indexed at 2000=100.	
or_{jt}	Oil rent per capita, in constant 2000 US dollars.	Authors' construction using data from the World Bank (2010b) Adjusted Net Saving Database.
o_{jt}	Value of oil production per capita, in constant 2000 US dollars.	
s_{jt}	Value of oil reserves per capita, in constant 2000 US dollars.	Authors' construction using data from Energy Information Administration of the United States Department of Energy (2010) and British Petroleum (2010).
p_t^o	Oil prices (Brent), in constant 2000 US dollars.	Authors' calculation using data from British Petroleum (2010) Statistical Review of World Energy.

Notes: Annual data between 1980 and 2006 ($T = 27$) for 53 Countries ($N = 53$), except for oil reserves for which we only have data for 52 countries (Hungary is excluded from this sample).

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