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Natural Resource Booms and Inequality: Theory and Evidence*

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Abstract

Surprisingly little is known about the impact of resource booms on income inequality in resource rich countries (Ross, 2007). This paper develops a simple theory, in the context of a two sector growth model in which learning-by-doing drives growth, to explain the time path of inequality following a resource boom. Under plausible conditions, we find that income inequality will fall in the short run immediately after a boom, and will then increase steadily over time as the economy grows, until the initial impact of the boom on inequality disappears. Using panel cointegration methodology for a sample of 90 countries between 1965 and 1999, we test the predictions of the model empirically. We find strong evidence in support of the theory. Resource booms, especially mineral booms, lower inequality in the year of the boom. This effect then gradually diminishes over time until inequality returns to its pre-boom level in the long run.

Keywords: Resource Booms; Inequality; Dutch Disease
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1 Introduction

The large windfalls from the current commodity price booms pose important questions for policy makers in natural resource-rich economies. The question that has received most attention is how the windfalls can be used to promote *economic growth*. However, surprisingly little is known about the impact of booms on *income inequality* (Ross, 2007). The persistence of high poverty rates in many resource-rich economies makes the question of how booms affect the distribution of income between the rich and the poor particularly relevant.

A traditional explanation for why many resource-rich countries grow slower than resource-scarce countries is Dutch Disease. The resource revenues lead to an appreciation of the real exchange rate, which harms the competitiveness of the non-resource exports sector and hampers economic growth if there are positive externalities to production in this sector (Corden and Neary, 1982; Van Wijnbergen, 1984; Sachs and Warner, 1995, 1999; Torvik, 2001). The theory of Dutch Disease provides a useful starting point for analyzing the effect of resource booms on inequality. It shows how booms affect the returns to capital and labor, which are primary determinants of the personal distribution of income (Daudey and García-Peñalosa, 2007).

This paper develops a simple theory, in the context of a two sector growth model in which learning-by-doing drives growth, to explain the time path of inequality following a resource boom. Under the plausible conditions of low elasticity of substitution between nontraded and traded goods in consumption, a relatively labor intensive nontraded sector, and balanced growth, we find that income inequality will fall in the short run immediately after a boom, and will then increase steadily over time as the economy grows, until the initial impact of the boom on inequality disappears.

The paper then tests the predictions of the model empirically. In particular, we adopt panel cointegration methodology to analyze global data for 90 countries between 1965 and 1999 to disentangle the short and long run effects of commodity export prices on income inequality. We find strong evidence in support of the theory. Resource booms, especially mineral booms, lower inequality in the year of the boom. This effect then gradually diminishes over time until inequality returns to its pre-boom level in the long run.

This paper is loosely related to the literature on natural resource endowments and inequality. Leamer et al. (1999) argue that, since resource exploitation does not require much human capital, the labor force in resource rich economies is unprepared for the emergence of human-capital-intensive manufacturing. As a result, these economies may experience higher income inequality for longer periods than resource-scarce economies. This could explain why inequality is higher in resource-rich Latin America than in resource-scarce East Asia. Sokoloff and Engerman (2000) instead point at how natural resource endowments affect inequality through the evolution of institutions. In colonies where economies of scale led to vastly unequal land ownership, like the ones in Latin America, the inequality was sustained by political institutions that favored the rich and excluded the poor. By contrast, in more temperate colonies such as the United States and Canada, the absence of economies of scale led to a much more equal land distribution, which was subsequently sustained by institutions that emphasized equal opportunities. Gylfason and Zoega (2003) argue that natural resource dependence leads to both lower growth and increased inequality, and could therefore explain the inverse relationship between growth and inequality in cross-country data. Finally, this paper is related to Ross (2007), who looks at possible mechanisms through which mineral wealth can affect inequality and discusses appropriate government responses.

The rest of this paper is organized as follows. In section 2 we present the model and state the pre-boom equilibrium. In section 3 we analyze the effects of a resource boom on inequality in the short run when productivity is assumed constant. Section 4 then extends the model to analyze the effects of a boom in the long run when productivity is assumed to be driven by learning-by-doing. In section 5 we turn to the empirical analysis and describe the data and methodology, including the error-correction specification and the panel unit root and panel cointegration tests. Section 6 reports the estimation results and presents impulse response analysis that shows the estimated time path of inequality following a resource boom. Section 7 concludes.

2 The Model

Our model of the dependent economy has a nontraded sector, which we call the N sector, and a non-resource traded sector, which we call the T sector. The resource sector of the model is represented as an exogenous gift of resource income R. There is a labor force of size L and a capital stock of size K. Both the N sector and the T sector use both capital and labor, which are the only factors of production, and the stocks of capital and labor are fixed. Letting L_N and L_T represent the labor force in the N and T sectors, respectively, the labor market clearing condition can be written as

$$L_N + L_T = L. \tag{1}$$

The short run version of our model is equivalent to the model in section III of Corden and Neary (1982), but under the restriction that the resource sector uses neither labor nor capital, and only generates income, which itself is distributed equally to all individuals.

While the issue of unequal distribution of resource income is certainly important in reality, as noted for example by Ross (2007), our purpose here is to focus on a distinct yet important channel through which resource spending affects inequality: the shift of the factors of production to the relatively labor-intensive nontraded sector as a consequence of the spending effect during resource booms. As documented in Gelb (1988) for oil rich countries, and emphasized in Neary and Purvis (1982), the spending effects of a boom have generally been acknowledged as the more important force driving adjustment. While it would be straightforward to modify our theoretical results on the change in overall income inequality to take into account asymmetry of resource income between capitalists and workers, we opt to abstract from this issue in order to focus on the channel mentioned above.

As in Corden and Neary (1982), sections II and III, the N and T sectors in our model use capital and labor. Writing K_N and K_T as the N and T sector capital stocks at any point in time, the capital market clearing condition reads

$$K_N + K_T = K. \tag{2}$$

In the case where capital is sector specific, it is assumed that K_N and K_T are fixed at the levels that obtain in the pre-boom equilibrium.

For simplicity, we will assume Cobb Douglas production functions with a unit elasticity of substitution between factors and constant returns to scale. Denoting output and productivity in the sectors by X_j and A_j, respectively, we have

$$X_N = A_N K_N^{\theta_{KN}} L_N^{\theta_{LN}} \quad \text{and} \quad X_T = A_T K_T^{\theta_{KT}} L_T^{\theta_{LT}}. \tag{3}$$

Constant returns to scale implies that $\theta_{Kj} + \theta_{Lj} = 1$, $j = N, T$. We make the assumption henceforth that the nontraded sector is relatively labor intensive: $\theta_{LN} > \theta_{LT}$. Additionally, we follow the notation of e.g. Jones (1965) in defining λ_{ij} as the proportion of the supply of factor i used by sector j in equilibrium. Market clearing requires that $\lambda_{iN} + \lambda_{iT} = 1$, $i = L, K$.

Profit maximization requires that labor, and sector-specific capital, earns its marginal product at all times. Since labor is perfectly mobile, this implies equalization of the marginal product between sectors, so that the wage is given by

$$w = p_N X_N'(L_N) = X_T'(L_T) \tag{4}$$

Here p_N is the relative price of nontraded goods in terms of traded goods. The returns to sector specific capital are given by

$$r_N = p_N X_N'(K_N) \quad \text{and} \quad r_T = X_T'(K_T) \tag{5}$$

in the two sectors, respectively. In the case where capital is perfectly mobile between sectors, we must have $r_N = r_T$, but when capital is sector specific, or transitions slowly in response to return differentials, the returns to capital in the two sectors can differ if a shock forces the capital market into temporary disequilibrium.

The profits to specific capital in each sector are equal to the value shares of capital in the output of each sector, so that

$$\pi_N = \theta_{KN} p_N X_N \quad \text{and} \quad \pi_T = \theta_{KT} X_T. \tag{6}$$

The return to sector specific capital in a sector multiplied by the sector's capital stock must equal profits, so that $r_N K_N = \pi_N$ and $r_T K_T = \pi_T$. The total profit to capital in the economy, π, is equal to the sum of sector specific capital profits: $\pi = \pi_N + \pi_T$.

Now we introduce the concept, following Mussa (1974), of the return to capital *as a whole* under capital specificity. The return r to the capital stock is defined according to $\pi = rK$. Thus, the overall return to specific capital can be expressed as a weighted average of the returns to capital in both sectors,

$$r = \lambda_{KN} r_N + \lambda_{KT} r_T, \tag{7}$$

where the weights $\lambda_{KN} = K_N/K$ and $\lambda_{KT} = K_T/K$ are the proportions of the capital stock currently in the N and T sectors, respectively.

With the factor returns in hand, we can write down a measure of income inequality. It is assumed that workers earn only labor income, and capitalists earn only capital income. There is no within-class inequality, only between-class inequality. Between-class inequality is naturally measured by the ratio of the income of one capitalist to the income of one worker. The population of workers is equal to the labor force, L. For simplicity, assume that the population of capitalists is also equal to L. This assumption will not affect our results, and is natural given we are only interested in the inequality of capital vis-à-vis labor. The natural measure of inequality is given by the ratio of the income of a capitalist to the income of a worker:

$$I = \frac{rK}{wL}. \tag{8}$$

Thus inequality is measured by the total income of capital divided by the total income of labor. This structural measure of inequality captures the value share of capital in producing output as a multiple of the value share of labor in producing output.

All agents are assumed to have identical preferences, and maximize a CES aggregator of N and T good consumption, which is given by

$$U = \left[(1-\gamma)^{1/\sigma} C_T^{\frac{\sigma-1}{\sigma}} + (\gamma)^{1/\sigma} C_N^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}} \tag{9}$$

Here σ is the elasticity of substitution between traded and nontraded goods. Since agents have identical preferences, we can treat aggregate consumption of the two goods as being determined by the decision of a representative agent who chooses C_N and C_T to maximize U subject to the budget constraint of the entire economy,

$$p_N C_N + C_T = Y. \tag{10}$$

Here $Y = p_N X_N + X_T + A_T R$ is aggregate income. First, it is important to note that resource income will be measured in the productivity units A_T of the traded sector, as in Torvik (2001). A boom will be considered as an increase in R. This detail is irrelevant in the short run, when productivity in both sectors is constant, but will be highly relevant in the transition to the long run, when productivity in both sectors is growing. We will discuss this further in section 4. Second, the assumption that capitalists and workers have identical tastes rules out the possibility for demand composition effects to play any role in the economy's response to booms. As in Obstfeld and Rogoff (1996), the overall price index in the economy, P, for the consumption basket that is the solution to the consumer's problem is given by

$$P = \left[1-\gamma+\gamma p_N^{1-\sigma}\right]^{\frac{1}{1-\sigma}} \tag{11}$$

The demand functions for traded and nontraded goods respectively are given by

$$C_T = (1-\gamma)P^{\sigma-1}Y \quad \text{and} \quad C_N = \gamma P^{\sigma-1} p_N^{-\sigma} Y. \tag{12}$$

The model is closed by the requirement that the market for nontraded goods must clear, which is to say, that the output produced and consumed must be equal:

$$X_N = C_N. \tag{13}$$

By Walras' Law, we can omit the market equilibrium for traded goods.

In what follows, we proceed in three steps. First, we discuss the short run effects of a boom on inequality when productivity in both sectors is constant. In that exercise, we consider both the case of sector specific capital and mobile capital. Second, we consider the case where productivity growth in both the N and T sectors is driven by learning-by-doing. That exercise allows us to analyze the long run effects of a permanent resource boom on inequality.

3 The Short Run Effects of a Resource Boom on Inequality: The Cases of Sector Specific and Mobile Capital

Having completed the description of the model, let us first state the static effects of a boom on inequality when the productivity in the two sectors, A_N and A_T, is constant. In the short run, the only means that the economy has to respond to a resource boom, is via a change in the price of nontraded goods, p_N, and a movement of labor, and capital if it is mobile, into the nontraded sector in order to increase production of nontraded goods in response to increased demand. It is known from Corden and Neary (1982), section II, that the spending effect of a resource boom, under capital specificity, will cause the price of nontraded goods p_N to appreciate and labor to move from the traded to the nontraded sector. The wage w will increase, the return r_N to specific capital in the N sector will increase, and the return r_T to specific capital in the T sector will decrease. All of these structural effects are well known.

Denote log changes in variables by $\hat{I} = d\ln I = dI/I$. The proportional change in inequality is given by the proportional increase in the overall return to capital minus the proportional increase in the wage:

$$\hat{I} = \hat{r} - \hat{w}. \tag{14}$$

There are three cases worth considering in light of the literature: the case where capital is sector specific and immobile, the case where capital is initially immobile but moves slowly in response to return differentials created by the resource boom, and the case where capital, like labor, is perfectly mobile. Corden and Neary (1982), section II, consider the case where capital is sector specific, but do not work out the sign of the change in inequality in this case. The first theorem of the paper, stated below, provides an answer to that question.

Theorem 1 *When capital is sector specific, inequality will rise when the nontraded sector is relatively capital intensive ($\theta_{LN} < \theta_{LT}$), and will fall in the short run when the nontraded sector is* *relatively labor intensive ($\theta_{LN} > \theta_{LT}$).*

Proof. See Appendix 1. ■

The inequality result of Corden and Neary (1982), in other words, also holds in the case of capital specificity. The next natural question is what happens to inequality in the "medium run", when capital transitions slowly between sectors in response to the fact that returns to sector specific capital in the N sector are higher than those in the T sector.

For the dynamics of capital, we follow Neary (1978), who examines adjustment in a model with two traded goods, in assuming that the rate of capital movement between sectors is an increasing function of the return differential:

$$\dot{K}_N = \nu(r_N/r_T) = -\dot{K}_T. \tag{15}$$

Here ν is an increasing function of the ratio of sector specific returns at any moment in time, with $\nu(1) = 0$ and $\nu'(r_N/r_T) > 0$.

In Corden and Neary (1982), the models of sections II and III with only spending effects correspond to the short run and post-transition versions of this model. Corden and Neary (1982) compare the equilibria of those two models to the pre-boom equilibrium, but not to each other. We fill in that comparison by adding dynamics to the capital stock. The result is provided by our second theorem:

Theorem 2 *If capital transitions slowly between sectors in response to return differentials after the shock, the sign of the change in inequality will take the sign of the product $(\theta_{LN} - \theta_{LT})(\sigma^* - \sigma)$, where $\sigma^* < 1$ is an endogenously defined threshold for the elasticity:*

$$\sigma^* = \frac{\chi\theta_{KT} + \theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}}{\theta_{KT} + \theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}} \tag{16}$$

and $\chi = X_T/(X_T + A_T R)$.

Proof. See Appendix 1. ■

Inequality will only increase during the transition of capital when the elasticity of substitution in consumption between N and T goods is sufficiently low, because in that case labor will need to move in the opposite direction of capital, to offset declining production in the T sector as capital moves out of that sector in response to higher returns in the N sector. Regardless of the direction of the change in inequality during the period in which capital transitions, the net result must be the same as the result for the change in inequality in the case when both labor and capital are perfectly mobile. The case of perfect labor and capital mobility is discussed in Corden and Neary (1982), section III. The net result on our measure of inequality follows directly from Corden and Neary (1982):

Theorem 3 *When both capital and labor are perfectly mobile, inequality will rise when the non-traded sector is relatively capital intensive, and will fall in the short run when the nontraded sector is relatively labor intensive.*

The change in inequality under perfect capital mobility will be of the same sign as the change in inequality under capital specificity, although the magnitude of the change may be greater than or less than that under capital specificity.

The salient point about the change in inequality, holding the total factor productivity of each sector constant, can be summarized succinctly: when spending effects are the primary driver of short run adjustment to resource booms, inequality is likely to fall because in most countries, the nontraded sector, usually identified with services, real estate, and sometimes agriculture, is likely to be more labor intensive than the traded sector, which is usually identified with manufacturing. This result can generally be expected to hold regardless of the nature of capital adjustment.

4 The Long Run: Productivity Growth and the Path of Inequality in the Dependent Economy

To analyze the transition of the economy to the long run, we will henceforth assume perfect labor and capital mobility in response to a resource boom, and relax the assumption of constant total factor productivity levels in the N and T sectors. That is, we will focus only on the case where labor and capital markets are always in equilibrium, as opposed to the cases of sector specific and sluggishly mobile capital considered in the previous section.

The rate of productivity growth in each sector will be endogenously determined, and we will consider the effects of learning-by-doing (LBD) in both sectors with the possibility of knowledge spillovers. The model in this section is a generalization of the model of Torvik (2001), who considers productivity growth in a specific factors model without an explicit consideration of the role of capital. Including capital makes it possible to study the level of inequality, both in the balanced growth steady state, and in response to resource booms.

This section proceeds as follows. First, we describe our assumptions about the nature of productivity growth in both sectors. Second, we discuss the condition for dynamic stability, which requires that the elasticity of substitution $\sigma \leq 1$. Third, we discuss the relationship between inequality and growth in the economy with balanced growth, and analyze the impact of resource booms.

In the specification of Torvik (2001), the evolution of productivity in each sector is given by the following pair of differential equations:

$$\frac{\dot{A}_N}{A_N} = u\lambda_{LN} + v\delta_T\lambda_{LT} \tag{17}$$

$$\frac{\dot{A}_T}{A_T} = u\delta_N\lambda_{LN} + v\lambda_{LT} \tag{18}$$

These deserve comment. First, productivity growth is driven by LBD in both sectors, and the amount of LBD is determined by sector employment levels. Second, the constants δ_T and δ_N measure, respectively, the size of spillover effects of LBD in the T sector on N sector productivity growth, and of LBD in the N sector on T sector productivity growth. Third, the terms $u\lambda_{LN}$ and $v\lambda_{LT}$ measure, respectively, the direct effects of LBD in the N sector and of LBD in the T sector. In this general specification, both sectors have the capacity to generate productivity growth, as well as to benefit from productivity growth in the other sector.

Important previous papers in the literature on the dynamic effects of the Dutch Disease assume dynamic specifications for productivity growth that can be nested within the above specification. In these equations, as Torvik (2001) notes, Van Wijnbergen (1984) and Krugman (1987) represent the case where $u = \delta_T = \delta_N = 0$, while Sachs and Warner (1995) represent the case where $u = \delta_N = 0$ and $\delta_T = 1$. The first of these subcases gives rise to unbalanced growth, as the traded sector generates and benefits from productivity growth, while the nontraded sector does neither. As shown by Torvik (2001), unbalanced growth in favor of either sector leads to the eventual specialization of the economy in the production of the favored good. In this case of Van Wijnbergen (1984) and Krugman (1987), productivity growth only occurs in and benefits the T sector, so the economy will eventually specialize in the production of T goods.

The second of these cases, studied in Sachs and Warner (1995), gives rise to balanced growth, as the traded sector exclusively drives productivity growth, but both sectors benefit equally from it. However, in this case, the productivity of both sectors must be equal.

In the more general specification that we employ, it is possible to have balanced growth, with the *ratio* of productivity levels in the two sectors constant, but not necessarily equal to one. Define this ratio by

$$\phi = A_T/A_N. \tag{19}$$

Then the rate of change of the ratio over time is governed by

$$\frac{\dot{\phi}}{\phi} = \frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} = -u(1-\delta_N)\lambda_{LN} + v(1-\delta_T)\lambda_{LT}. \tag{20}$$

In order to determine the dynamic stability of the system, it is necessary to see for what range of ϕ the growth rate of the productivity differential, $\dot{\phi}/\phi$, is positive and in what range it is negative. Let ϕ^* denote the value at which $\dot{\phi}/\phi = 0$. To achieve such a dynamic equilibrium, we must have that the fraction of labor in the N sector is equal to:

$$\lambda^*_{LN} = \frac{v(1-\delta_T)}{u(1-\delta_N) + v(1-\delta_T)} \tag{21}$$

As we shall see shortly, the level of ϕ^* corresponding to λ^*_{LN} depends on the resource income R. If $\dot{\phi}/\phi > 0$ for $\phi > \phi^*$, the system will be unstable, and will exhibit unbalanced growth leading to complete specialization in one sector or the other. Conversely, if $\dot{\phi}/\phi < 0$ for $\phi > \phi^*$, the system will be stable, and will exhibit balanced growth in equilibrium, with the ratio of sector productivities given by ϕ^*.

To derive the condition for dynamic stability, we must determine the sign of the rate of change of relative productivity growth. This is found by computing

$$\frac{d(\dot{\phi}/\phi)}{dt} = -[u(1-\delta_N)\lambda_{LN} + v(1-\delta_T)]\frac{d\lambda_{LN}}{d\phi}\frac{d\phi}{dt} \tag{22}$$

For stability, we need $\frac{d(\dot{\phi}/\phi)}{dt}$ to take the opposite sign from $\frac{d\phi}{dt}$. This condition is satisfied if and only if the static effect of increased ϕ on N sector employment is positive: $\frac{d\lambda_{LN}}{d\phi} > 0$. In the model with mobile capital, this condition amounts to the same thing as in the model with fixed capital: dynamic stability requires that $\sigma \leq 1$.

To derive this comparative static result is algebraically more complicated, but has exactly the same intuition, as the case with specific capital. We relegate the formal comparative static proof to Appendix 2, and here give the intuition. An increase in the relative productivity of the traded sector has two offsetting effects. First, it creates an incentive for capital and labor to move into the T sector, as their marginal returns are higher there when ϕ rises. This decreases the relative supply of N goods, which will raise p_N. Since p_N is higher, there will be a consumption response that depends on σ. If $\sigma > 1$, consumers will raise the proportion of their incomes they spend on T goods, and this will add further incentives for labor and capital to move into the T sector.

Higher labor and capital in the sector will sustain $\dot{\phi} > 0$, and the result will be destabilizing, with ϕ growing without bound until the economy eventually specializes in T goods.

On the other hand, if $\sigma < 1$, the rise in the price p_N will be accompanied by an increase in the proportion of income that consumers spend on N goods, and a reduction in the proportion of income they spend on T goods. This will more than offset the incentive of factors to move to the T sector caused by the rise in ϕ, and factors will move into the N sector in response to an increase in ϕ. As a result, the rise in ϕ will eventually come to a halt, as seen from the equation for the dynamics of ϕ, and in a stable equilibrium λ_{LN} will be unchanged.

For $\sigma = 1$, we have the case of Cobb-Douglas preferences. In this case, we will have $\frac{d\lambda_{LN}}{d\phi} = 0$. Changes in ϕ do not affect employment in this case, so we will concentrate on the case $\sigma < 1$.

Before proceeding, note that if we have $\sigma > 1$, corresponding to unbalanced growth, then an economy that starts at its equilibrium, and experiences a resource boom, will be pushed onto a path in which it specializes in N goods in the limit as $t \to \infty$. Conversely, an economy that begins at the unstable equilibrium value for ϕ^* and experiences a resource bust will be pushed onto a path in which it specializes in T goods in the limit as $t \to \infty$. Torvik (2001) notes this in the case of the model with specific capital. Given that this is the case, then we can add simply that the level of inequality that obtains in either of those limits will be the value for I that corresponds to the ratio of the value product of capital to labor in the N sector, in the first case, or the T sector, in the second case. The former value is the lower bound for I in the economy, as we have already pointed out in section 3, and the latter value is the upper bound for I.

Now let us examine the value of inequality that obtains when the economy is at its long run balanced growth equilibrium. In equilibrium, the values of λ_{LN} and λ_{KN} are fixed by two conditions. The first condition is the requirement that $\lambda_{LN} = \lambda^*_{LN}$, which corresponds to dynamic stability. The second requirement is that both labor and capital markets be in equilibrium. Labor and capital market equilibrium is characterized by factor price equalization between the sectors and profit maximization:

$$w = \partial X_T / \partial L_T = p_N \partial X_N / \partial L_N \tag{23}$$

and

$$r = \partial X_T / \partial K_T = p_N \partial X_N / \partial K_N \tag{24}$$

These joint requirements define the *contract curve*, which traces out the locus of pairs (L_N, K_N), or equivalently $(\lambda_{LN}, \lambda_{KN})$, in the Edgeworth-Bowley box for the economy. The contract curve is defined by the relationship

$$\frac{\partial X_T/\partial K_T}{\partial X_T/\partial L_T} = \frac{\partial X_N/\partial K_N}{\partial X_N/\partial L_N}, \tag{25}$$

and is independent of ϕ. For the Cobb-Douglas production functions we have assumed, the contract curve relationship above states that we must have

$$\frac{\theta_{KT}}{\theta_{LT}}\frac{\lambda_{LT}}{\lambda_{KT}} = \frac{\theta_{KN}}{\theta_{LN}}\frac{\lambda_{LN}}{\lambda_{KN}} \tag{26}$$

in equilibrium. Note that both sides of this equation are equal to our measure of inequality I. On the contract curve, the level of inequality can be determined solely as a constant times the ratio of labor to capital in the sector. Making the substitutions $\lambda_{LN} = 1 - \lambda_{LT}$, and $\lambda_{KN} = 1 - \lambda_{KT}$, which must obtain according to full utilization of labor and capital, and using the fact that we must have $\lambda_{LN} = \lambda^*_{LN}$ in a dynamic equilibrium, it follows that the equilibrium fraction of capital in the N sector, which we will denote by λ^*_{KN}, is determined by

$$\frac{\theta_{KT}}{\theta_{LT}}\frac{\lambda^*_{LT}}{\lambda^*_{KT}} = \frac{\theta_{KN}}{\theta_{LN}}\frac{1-\lambda^*_{LT}}{1-\lambda^*_{KT}}. \tag{27}$$

Figure 1 illustrates this point in the Edgeworth-Bowley box, with capital on the vertical and labor on the horizontal axis. The nontraded sector labor force and nontraded sector capital stock are measured as distances from the point O_N in the bottom left hand corner, and the traded sector labor force and capital stock are measured as distances from the top right hand corner. The curve is drawn in $(\lambda_{LN}, \lambda_{KN})$ space, or equivalently, for $K = L = 1$. We have drawn the contract curve so that it lies everywhere below the diagonal 45 degree line $O_N O_T$, which corresponds to the assumption that the N sector is relatively labor intensive. The vertical line corresponds to $\lambda^*_{LN} = \frac{v(1-\delta_T)}{u(1-\delta_N)+v(1-\delta_T)}$, the steady state equilibrium fraction of labor in the nontraded sector.

As discussed in our analysis of the short run change in inequality, more unequal economies will be those in which more activity is concentrated in the relatively capital intensive sector. If this is the T sector, that means that inequality is decreasing as we move right on the horizontal axis from the point O_N: higher values for λ^*_{LN} correspond to lower equilibrium inequality in a balanced growth economy, when the N sector is relatively labor intensive. The reason is simply that the ratio of capital to labor income in the economy will vary between two extremes, which are represented by the ratios θ_{KN}/θ_{LN} and θ_{KT}/θ_{LT}. These are the ratios of the value products of capital and labor in the N sector, and the T sector, respectively.

Let us now relate the structure of LBD effects to equilibrium inequality in the balanced growth economy. First, from the above discussion, a structure for LBD effects that implies a high value for nontraded sector employment λ_{LN}^* will imply a low level of long run equilibrium inequality, when the N sector is relatively labor intensive, and the opposite will be true if the N sector is relatively capital intensive. With that understanding, we will only discuss the case where the N sector is relatively labor intensive. From the formula for λ_{LN}^*, it is apparent that we will thus have lower long run equilibrium inequality when the strength v of LBD in the T sector is high, when sector T to sector N learning spillovers δ_T are low, when LBD in the N sector u is low, and sector N to sector T spillovers δ_N are high.

The intuition behind these results is as follows. The stronger are direct LBD effects v in the T sector, other things equal, the more quickly ϕ will tend to rise for a given T sector employment level. In order to have ϕ constant, as required by a balanced growth equilibrium, however, we must then have *more* labor in the N sector, in equilibrium, to balance out the tendency of labor to generate a lot of T sector productivity growth via LBD when it is in that sector. This is just a consequence of demand inelasticity. On one hand, the profit motive drives capital and labor into the T sector when productivity growth there is relatively high. On the other hand, demand inelasticity works in the opposite direction, since higher ϕ exerts upward pressure on p_N, which causes individuals to increase the fraction of N goods in their consumption basket. When $\phi < 1$, the demand elasticity effect is stronger.

If v is higher, therefore, the result is that more labor and capital have to stay in the N sector in order for the dynamics of the system to be stable. When the T to N spillovers δ_T are high, other things equal, strong direct LBD effects in the T sector do not require the same degree of factor movement into the N sector for the reasons just discussed, so more labor and capital will be in the relatively capital intensive T sector. This will correspond to higher long run inequality. When LBD effects u in the N sector are low, more labor and capital must reside there to produce for the inelastic demands of consumers for N goods, so inequality will be lower. Finally, when δ_N is high, we use a parallel, but reversed, logic as for the role of δ_T to see that more labor can be in the N sector, and thus inequality will be lower.

4.1 The effect of a resource boom

Now let us examine the effect of a resource boom, meaning a one-off increase in R, on the path taken by inequality. From the comparative statics of the model for constant levels of A_T and A_N, we saw that N sector employment will increase, and when the N sector is relatively labor intensive, inequality I will fall when capital is sector specific. When capital is mobile, as in this model, it is well known that r must fall and w must rise due to the spending effect of higher R. Thus, as in the case of specific capital, inequality must fall upon impact of higher R. This is generally consistent with what we have seen in the previous two sections when the movement of capital is sluggish.

Also, due to increased demand for N goods, both capital and labor will move into that sector. Thus, for given values of A_N and A_T, a resource boom will increase the growth rate of A_N and decrease the growth rate of A_T upon impact, relative to the (common) rates of growth in the sectors that obtained before the boom. With $\lambda_{LN} > \lambda_{LN}^*$, this will initially set

$$\frac{\dot{\phi}}{\phi} = \frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} < 0.$$

The relative productivity ϕ will continue to fall, and λ_{LN} decrease, until $\lambda_{LN} = \lambda_{LN}^*$ is again reached, and a lower, equilibrium level of ϕ, is obtained. These dynamics are discussed in the specific capital version of the model by Torvik (2001). One additional point worth noting is that, as a consequence of this feature of the balanced growth economy, the value of ϕ^* in equilibrium will in general reflect the history of the shocks to R that have been experienced by the economy. As λ_{LN} decreases in the transition back to its lower, equilibrium value, inequality will be increasing.

Finally, it is worthwhile to take note of the connection between inequality and growth following a resource boom. If direct LBD effects in the T sector, measured by v, are strong relative to direct LBD effects in the N sector, measured by u, then the initial movement of labor out of that sector, ceteris paribus, may lead to a reduction in the overall growth rate. If we take the overall growth rate to be a weighted average of the rates of growth of A_N and A_T, then the growth rate will fall as long as the weight on the growth of A_N is not too high. In this case, which seems relevant in light of the strong assumptions on T sector productivity growth made in past literature, the dependent economy is likely to exhibit an *inequality-growth tradeoff* following resource booms.

To summarize, our theoretical results indicate that under plausible conditions, most economies will experience a short run fall in inequality following resource booms, followed by a period of rising inequality, with the net result that the long run value of inequality will be the same as its pre-boom value under conditions of balanced growth, even when the size of the resource income relative to the size of the economy is permanently larger. We next test the predictions of our model empirically.

5 Methodology and Data

In this section, we describe the econometric methodology and the variables used in estimation. Data description and sources can be found in Appendix 3. Panel unit root and panel cointegration tests are discussed in Appendix 4. We will analyze the effects of commodity export prices on household income inequality using the error-correction model given by equation (28) below.

$$\Delta y_{i,t} = \alpha_i + \delta' z_{i,t} + \lambda y_{i,t-1} + \beta_1' x_{i,t-1} + \beta_2 \Delta y_{i,t-1} + \beta_3' \Delta x_{i,t} + \epsilon_{i,t} \quad (28)$$

In the above equation, the subscripts $i = 1, ...N$ and $t = 1, ...T$ index the countries and years in the panel, respectively. Here $y_{i,t}$ stands for household income inequality in country i in year t, α_i is a country-specific fixed effect, and $z_{i,t}$ is an $rT \times 1$ vector of regional time dummies, where r is the number of regions.[1] The term $x_{i,t-1}$ is an $m \times 1$ vector of m variables that are expected to affect inequality both in the short run and the long run. We next discuss how the key components of equation (28) were constructed.

Our measure of household income inequality, $y_{i,t}$, was taken from Galbraith and Kum (2005) and is based on the Deininger and Squire (1996) household income inequality dataset and the UTIP-UNIDO dataset on manufacturing pay inequality. It is much more comprehensive than the Deininger and Squire inequality measure and uses a more informed filling-in of missing observations than other studies.

The vector $x_{i,t-1}$ includes, among other things, a commodity export price index to test the effect of commodity export prices, which was constructed using the methodology of Deaton and Miller (1995), Dehn (2000), and Collier and Goderis (2007a). We first collect data on world commodity prices and commodity export values for as many commodities as data availability allowed. Table 1a lists the 50 commodities in our sample. We then construct weights by dividing the individual 1990 export values for each commodity by the total value of 1990 commodity exports for every country. These weights are 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 more commodity-dependent countries, we weight the (logged) deflated index by the share of commodity exports in

[1] 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.

a country's GDP. To investigate whether the effects vary across different types of commodities, we also experiment with sub-indices for non-agricultural and agricultural commodities, which were constructed along similar lines.

To round out the vector $x_{i,t-1}$ of independent variables, we include three control variables from Barro (2000): log real GDP per capita, a measure of democracy (based on the number of political constraints), and a measure of educational attainment (expressed as the average years of primary schooling of the population aged 15 and over). These control variables are statistically significant in the specification of equation (28).[2]

Our dataset includes all countries and years for which data are available, and covers 90 countries for the period 1965 to 1999. Table 1b lists the countries in our sample along with their level of commodity exports. Table 2 reports summary statistics for the variables used in estimation.

6 Estimation results

The results of estimating equation (28) are reported in Table 3.[3] The specification in column (1) includes all regressors except the commodity export price indices. The long-run variables enter with the expected signs and are statistically significant. The level of GDP per capita enters with a negative sign and is significant at 1 percent, indicating that long-run economic development leads to lower inequality. Democratization and education also lower inequality, as the political constraints indicator and the measure of primary schooling both enter with a negative sign and are significant at 5 and 10 percent, respectively.

We next turn to the short-run variables. The lagged level of inequality enters with a negative sign and is statistically significant at 1 percent. The size of the coefficient indicates that the speed of adjustment to long-run equilibrium is around 20 percent per year. The lagged change in inequality also enters with a negative sign and is significant at 5 percent, which implies that shocks to inequality tend to be followed by shocks in the opposite direction. We experimented with additional lags of the change in inequality and with lags of the changes in the long-run variables

[2] We considered a wide range of additional control variables, including GDP per capita squared (to test the hypothesis underlying the Kuznets curve), measures of secondary and higher schooling, alternative measures of democracy, various governance indicators, measures of trade and capital account openness, inflation, external debt, political violence, commodity price volatility, measures of industrial development, natural disasters, and the political orientation (e.g. left wing, centre, or right wing) of government parties or the executive. 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. However, we do use them when we test the robustness of our results in the next section.

[3] The long-run coefficients correspond to $-(\frac{1}{\lambda}) \cdot \beta_1$ in equation (28). The short-run coefficients correspond to λ, β_2, and β_3 in equation (28).

but found these to be unimportant.

In Table 3, column (2), we include both the lagged level and the change in the commodity export price index to test the long-run and short-run effects of commodity export prices on inequality. The long-run coefficient is positive but highly insignificant, indicating that commodity export prices do not affect the long-run level of inequality. The short-run effect of a positive export price shock is negative, suggesting that commodity booms lead to lower inequality in the same year. However, the coefficient is not statistically significant so should be viewed with caution.[4] In Table 3, column (3), we drop the lagged level of the commodity export price index and find a similar result for the change in the index.

We next investigate whether the short-run effect of higher commodity prices varies across different types of commodities by replacing the change in the general index with the changes in the two sub-indices for non-agricultural and agricultural commodities. The results are reported in Table 3, column (4). Both variables enter with a negative sign but the change in the non-agricultural commodity export price index is now statistically significant at 5 percent. This indicates that a rise in the prices of non-agricultural commodities leads to a fall in inequality in the same year. The change in the agricultural index is not statistically significant but its coefficient is only slightly smaller than the coefficient for the change in the non-agricultural index.[5] In Table 3, column (5), we exclude the insignificant effect of agricultural prices. The results for all the other regressors are unchanged. In particular, the change in the non-agricultural commodity export price index again enters negative and is significant at 5 percent, as in column (4), while the size of the coefficient is almost unchanged. These results are consistent with the short run prediction of our theoretical model that inequality should fall in the short run in response to a resource boom.

We next use the results in Table 3, column (5), to graphically show the time path of inequality following a resource boom. Figure 2 shows the impulse response functions of inequality to a two standard deviation shock to the growth rate of non-agricultural commodity export prices for different levels of exports. Inequality falls in the year of the boom and then gradually moves back to its original pre-boom level. The decline in inequality is larger for higher levels of exports over GDP.

An example of a country that depends on non-agricultural commodities is the Republic of

[4] We experimented with additional lags of the change in the commodity export price index but found these to be unimportant. We also allowed for a possible effect of higher oil prices on oil *importing* countries by including an oil import price index, but found no evidence of any systematic effect. All our results go through when including this oil import price index.

[5] An F-test did not reject the null hypothesis of equal coefficients with a p-value of 0.96.

Congo. In 1990, Congo's oil exports represented 28 percent of its GDP.[6] The results in Table 3, column (5), which are illustrated in Figure 2, imply that for a country like Congo, an increase in the growth rate of the oil price of two standard deviations (33 % points) leads to a 0.3 point lower Gini index of inequality in the year of the shock.[7] From that point onward, inequality increases steadily over time until the initial impact of the boom on inequality disappears. The speed of adjustment is such that around five years after the shock, two thirds of its initial impact has died out. These empirical results on the time path of inequality after a commodity boom lend strong support to the predictions of our theoretical model.

We next test the robustness of our finding that higher non-agricultural commodity prices lower inequality in the short run. Table 4, column (1), shows the results when excluding the regional time dummies of equation (28). The short-run effect of non-agricultural commodity export prices remains statistically significant at 5 percent, while the size of the coefficient is somewhat smaller. To investigate the possibility that the effect of a change in the growth rate of non-agricultural commodity prices, from for example -25 to 0 percentage points, differs from the effect of a change from 0 to +25 percentage points, we separate the change in the non-agricultural commodity export price index into two variables that capture positive and negative changes in the index. The results are reported in Table 4, column (2). The coefficient for the positive shocks is again negative, significant at 5 percent, and somewhat larger than in earlier specifications. The coefficient for the negative shocks is also negative, although not significant, and somewhat smaller. The difference between the coefficients is not statistically significant, however, which suggests the absence of any substantial non-linearity. Nonetheless, the significant coefficient for the positive shocks is reassuring, as it means that our finding that higher non-agricultural commodity prices lower inequality is not solely driven by episodes of decreasing prices, but very much stems from boom episodes. As a further robustness test, in Table 4, column (3), we replace the fixed effects estimator by an ordinary least squares estimator. The short-run effect of higher non-agricultural commodity export prices is robust to this alternative specification. The coefficient is again significant at 5 percent and gains in size compared to the benchmark specification in Table 3, column (5).

We have so far ignored the possible endogeneity of the non-agricultural commodity export price index. As argued by Deaton and Miller (1995), international commodity prices are typically

[6] In our sample, the share of non-agricultural commodity exports in GDP varies between 0.00 for countries like Japan and Benin, and 0.34 for Zambia.

[7] For an increase in the growth rate of the oil price of 50 or 100 % points, this effect amounts to 0.5 and 0.9 Gini point, respectively. The current commodity boom has seen large cumulative price increases in non-agricultural commodities. Between 1996 and 2006 the price of oil more than tripled and the price of other non-agricultural commodity exports more than doubled (Collier and Goderis, 2007b).

not affected by individual countries and are therefore not likely to be endogenous with respect to the growth of individual countries. They further argue that by fixing the shares of individual commodity exports in total exports, any supply responses to price changes are not affecting the index. These arguments were made in relation to economic growth, but they are likely to hold in the case of household income inequality as well. However, even though the assumption of exogenous world commodity prices is probably justified for many countries in our sample, some countries that are major exporters of one or more commodities may have an influence on the world price of those commodities. For those countries, the exogeneity assumption is not justified and might lead to biased estimates. To address this concern, we express each country's exports of a given commodity as a share of the total world exports of that commodity and repeat this for all other commodities in our sample. This yields a list of commodity export shares that reflect the importance of individual exporters in the global markets for individual commodities. We found that of the 90 countries in our sample, 22 countries export at least one commodity for which their share in world exports exceeds 20 percent, while 34 countries export at least one commodity for which their share in world exports exceeds 10 percent. We investigate whether the inclusion of these major exporters in our sample affected our results by re-estimating the specification in Table 3, column (5), for two subsamples: one in which we exclude the group of 22 major exporters with exports larger than 20 percent, and one in which we exclude the group of 34 major exporters with exports larger than 10 percent.

Although the exclusion of these two groups decreases the sample size by almost a third and half, respectively, our finding that higher non-agricultural commodity export prices lower inequality in the short run is robust. The coefficient of the change in the non-agricultural price index only marginally changes from -3.26 (5 % significance) to -3.30 (10 % significance) and -3.76 (10 % significance) for the two subsamples, respectively. Hence, our results do not seem to be biased by countries that are major exporters of one or more commodities and that may influence world prices of these commodities.

However, this does not take away all concerns of endogeneity. As explained in section 5 above, the construction of the commodity price indices involves weighting them by the ratio of commodity exports over GDP. This ratio might be endogenous as it is potentially driven by factors that also affect inequality, such as governance or government policies.[8] For example, a resource-rich country with bad policies that cause low growth, and high poverty and inequality, will have a relatively

[8] The possible endogeneity of resource dependence is addressed by Brunnschweiler and Bulte (forthcoming) in relation to economic growth.

high commodity exports to GDP ratio due to the lack of development of its non-resource sectors. This implies that our procedure attaches higher weights to poor countries, which could bias the results. In as far as this potential endogeneity arises from time-invariant omitted variables, the fixed effects in our regression solve the problem. But any time-varying omitted variables could lead to a biased estimate of the effect of non-agricultural commodity export prices. To address this concern, we need to instrument for the share of non-agricultural commodity exports in GDP.

Following Collier and Goderis (2007a), we use estimates of the stock of sub-soil assets[9] (minerals) in 2000 developed by the World Bank (2006), as an instrument. These estimates are based on the net present value of a country's expected benefits over a horizon of 20 years and include 13 commodities, 12 of which are also included in our non-agricultural commodity index. The ratio of non-agricultural commodity exports over GDP does not enter our specifications by itself but only as a weight of the non-agricultural price index. We therefore construct an instrument for the index in the following way (following Collier and Goderis, 2007a). We construct the log of the deflated non-agricultural commodity export price index, as before, but now instead of weighting it by the ratio of non-agricultural commodity exports over GDP, we weight it by the 2000 value of sub-soil assets in current US dollars per capita. We then use the difference of this newly constructed variable as an instrument for the difference of the non-agricultural export price index in Table 3, column (5). For the value of sub-soil assets to be a valid instrument, it should be correlated with the ratio of non-agricultural exports over GDP, and it should not be correlated with the error term. The former is likely to hold, as commodity exports (net of imports) are only possible if those commodities are available in a country. The latter requires that the instrument does not itself affect inequality, other than through its effect on the endogenous variable (exclusion restriction), does not depend on inequality, and is not correlated with omitted variables that affect inequality. The first requirement is likely to be fulfilled as it is hard to see how a country's resource abundance could affect how it responds to commodity export prices, other than through its relationship with the level of commodity exports. The other two requirements are less likely to be fulfilled. Unequal countries tend to be poorer and are less likely to invest in geological exploration and more likely to overexploit the discovered stock of resources. As a result, their stock of discovered resources in the ground may be lower than in more equal and richer countries, everything else equal. This means that weighting the non-agricultural export price index by the value of sub-soil assets per capita may imply giving higher weights to richer countries. Although this could potentially bias

[9] Brunnschweiler and Bulte (forthcoming) and Arezki and Van der Ploeg (2007) use the World Bank estimates of natural capital to proxy resource abundance.

the results, the direction of the bias is likely to be opposite to the direction of the bias in the uninstrumented regressions, where higher weights were given to poor countries. Comparing the coefficients of the instrumented and uninstrumented regressions can therefore shed light on the size of the potential bias and the numerical range within which the actual coefficient is likely to be located.

The results of the fixed effects instrumental variables estimation are reported in Table 4, column (4).[10] The change in the non-agricultural commodity index again enters negative but is no longer statistically significant. However, the size of the coefficient is similar to the coefficient in Table 3, column (5), and even slightly larger. Given that any potential biases in the OLS and IV estimates are likely to have opposite signs, this suggests that the bias in the OLS estimates is likely to be small and, if anything, leads to an underestimation of the short-run effect of non-agricultural commodity export prices. In fact, a Davidson-MacKinnon test of exogeneity did not reject the null hypothesis of exogeneity (p-value is 0.71). In Table 4, column (5), we repeat this IV estimation procedure but using the 1994 estimates of sub-soil assets in current US dollars per capita developed by the World Bank (1997). The coverage of the 1994 estimates is lower than the coverage of the 2000 estimates but the time lag with the 1990 share of commodity exports over GDP is smaller. The results are similar to the results of column (4). The change in the non-agricultural commodity index again enters negative, although insignificant, and the size of the coefficient is similar and again slightly larger than the coefficient in Table 3, column (5). The Davidson-MacKinnon test of exogeneity again fails to reject the null of exogeneity (p-value is 0.66). All in all these results indicate that the uninstrumented estimate of the effect of non-agricultural commodity prices in Table 3, column (5), is consistent.[11]

Finally, we performed a large number of other robustness checks by adding additional control variables to the specification in Table 3, column (5). For each of these variables we run two regressions. The first only includes the lagged level of the variable, while the second includes the lagged level, the contemporaneous difference, and any additional significant lagged differences. The variables we consider are GDP squared, secondary and higher schooling, alternative indicators of democracy, several institutional quality indicators including the black market premium, civil liber-

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

[11] To further investigate possible endogeneity, we repeated the IV estimations in Table 4, columns (4) and (5), for the two subsamples in which we exclude the two groups of major exporters that may influence world prices. The coefficients for the change in the non-agricultural were similar or even larger: −3.94 and −4.05 for the group without the countries with exports > 20 %, and −7.17**and −6.81** for the group without countries with exports > 10 %. These results further support the consistency of the OLS estimates.

ties, political rights, checks and balances, and the International Country Risk Guide (ICRG) rating, and several indicators that capture the type of government (left wing, centre, or right wing). We also add several variables from the empirical growth literature, including capital account openness, international reserves, trade openness, inflation, external debt, assassinations, financial development, civil war, coup d'etat, natural disasters, commodity export price volatility, and industrial development. In total, we perform 50 additional robustness checks. In all these regressions, the coefficient of the change in the non-agricultural index has a mean value of -3.43 and varies between -5.29 and -2.15, which is close to the value in Table 3, column (5). It is statistically significant at 5 percent in 32 of the cases, significant at 10 percent in 11 other cases, and insignificant in 7 out of the 50 cases. In 6 of the latter 7 cases, running additional regressions revealed that the coefficient turned insignificant due to the smaller sample and not because of the inclusion of the additional regressor. We also ran the two regressions including all the additional regressors together. Despite the severely limited number of observations, the coefficient remained statistically significant at the 10 percent level. These results indicate that our finding that inequality falls in the year of the boom is robust to the inclusion of additional control variables.[12]

7 Conclusions

This paper has analyzed the time path of income inequality following a resource boom. We first presented a model that predicts that inequality will fall in the short run immediately after a boom and then increases steadily over time, until the initial impact of the boom on inequality disappears. We then tested the predictions of the model empirically and found strong evidence in support of the theory. Resource booms, especially mineral booms, lower inequality in the year of the boom. This effect then gradually diminishes over time until inequality returns to its pre-boom level in the long run.

Our results have important policy implications. They indicate that the often voiced concern that resource booms systematically increase inequality is not supported by the data. Hence, although policy makers in resource-rich economies face important challenges as to how resource booms should be managed, concerns about general income distribution do not seem to be particularly justified.

Several comments are in order. First, although we find that general income inequality falls

[12] To save space, we do not report these additional estimation results but they are available upon request from the authors.

immediately after a boom, this does by no means imply that a boom does not put certain groups within society at a disadvantage. Ross (2007) argues that some types of workers may lack the skills to move sectors and will become unemployed as a result of booms. Obviously, policies aimed at training or compensation for these workers have clear social benefits.

Secondly, this paper has ignored possible other transmission mechanisms through which booms may affect inequality. One example is public sector employment. Resource windfalls often generate new government jobs, which may reduce income inequality (Ross, 2007). However, these jobs are often created by governments in weak governance countries to buy political support (Robinson et al. (2006). Such inefficient redistribution, even when lowering inequality, often comes at a cost of lower growth.

Appendix 1: Proofs

Proof. *Proof of Theorem 1*: In general, the change in the overall return to capital is given by the change in sectoral profits weighted by the share of profits in each sector,

$$\hat{r} = \left(\frac{\pi_N}{\pi}\right)\hat{\pi}_N + \left(\frac{\pi_T}{\pi}\right)\hat{\pi}_T, \tag{29}$$

At the moment of impact of the boom, the economy is in long run equilibrium. Thus $r_N = r_T$, and the profit shares are equal to the capital shares in each sector: $\pi_N/\pi = \lambda_{KN}$ and $\pi_T/\pi = \lambda_{KT}$. Since capital is immobile, the change in profits in a sector is equal to the change in returns to the respective specific capital stocks. The change in nontraded sector profits is equal to the increase in the price of nontraded goods plus the increase in output, $\hat{\pi}_N = \hat{p}_N + \hat{X}_N$. The change in traded sector profits is equal to the change in traded sector output, in value terms, since traded goods are the numéraire: $\hat{\pi}_T = \hat{X}_T$. Substituting these relationships into the equation for $\hat{r}$ yields

$$\hat{r} = \lambda_{KN}(\hat{p}_N + \hat{X}_N) + \lambda_{KT}\hat{X}_T. \tag{30}$$

Now we invoke profit maximization and perfect labor mobility, which together imply that the change in the marginal value product of labor in the two sectors must be equal: $\hat{w} = \hat{p}_N + \hat{X}_N - \hat{L}_N = \hat{X}_T - \hat{L}_T$. The full utilization of labor in the economy, in changes, can be stated as the equation $\lambda_{LN}\hat{L}_N + \lambda_{LT}\hat{L}_T = 0$. Combining these two equations to solve for $\hat{p}_N + \hat{X}_N$, and substituting into the equation for $\hat{r}$, we can write the change in the rate of return on capital, in the short run, in terms of a coefficient times the change in the traded sector labor force:

$$\hat{r} = \left(\theta_{LT} - \frac{\lambda_{KN}}{\lambda_{LN}}\right)\hat{L}_T \tag{31}$$

Here we have used the fact that in the short run, $\hat{X}_T = \theta_{LT}\hat{L}_T$, since capital is fixed. The results of Corden and Neary (1982), section II, tell us that $\hat{L}_T < 0$ in the short run. Thus the sign of the change in the overall return to capital, $\hat{r}$, is determined entirely by the sign of the coefficient $\theta_{LT} - \frac{\lambda_{KN}}{\lambda_{LN}}$.

To compute the change in inequality, $\hat{I}$, it remains to compute the change in $\hat{w}$. Using the change in the marginal product of labor in the T sector, this is given by

$$\hat{w} = \hat{X}_T - \hat{L}_T = -\theta_{KT}\hat{L}_T. \tag{32}$$

Substituting the expressions for $\hat{r}$ and $\hat{w}$ into the equation for $\hat{I}$, we obtain

$$\hat{I} = \left(1 - \frac{\lambda_{KN}}{\lambda_{LN}}\right)\hat{L}_T. \tag{33}$$

The sign of $1 - \frac{\lambda_{KN}}{\lambda_{LN}}$ is determined entirely by the production side of the model, and moreover, is equal to the sign of $\theta_{LN} - \theta_{LT}$, as shown by e.g. Jones (1965). ■

Proof. *Proof of Theorem 2*: We will first derive the paths taken by the the returns, wage, and the real exchange rate when capital transitions. Using the changes in the returns and the wage, we will derive the path taken by inequality I.

From the comparative static results for the specific capital case, we know that we will have $r_N/r_T > 1$ when capital begins to transition. Thus, capital will begin to move from the traded to the nontraded sector, and seen from the fact that $\dot{K}_T = -\nu(r_N/r_T) < 0$. Letting $\chi = X_T/(X_T + A_T R)$, we obtain the the following comparative static relationship upon differentiating the log of the nontraded goods market clearing condition, holding R constant:

$$\sigma\hat{p}_N + \hat{X}_N = \chi\hat{X}_T. \tag{34}$$

The comparative static equations for changes in the output of the N and T sectors, respectively, with constant productivity, are given by

$$\hat{X}_N = \theta_{KN}\hat{K}_N + \theta_{LN}\hat{L}_N \text{ and } \hat{X}_T = \theta_{KT}\hat{K}_T + \theta_{LT}\hat{L}_T. \tag{35}$$

Labor market clearing requires that $\lambda_{LN}\hat{L}_N + \lambda_{LT}\hat{L}_T = 0$, and capital market clearing requires that $\lambda_{KN}\hat{K}_N + \lambda_{KT}\hat{K}_T = 0$. The equalization of the wage between sectors requires that

$$\hat{w} = \hat{p}_N + \hat{X}_N - \hat{L}_N = \hat{X}_T - \hat{L}_T. \tag{36}$$

Finally, the static equations for changes in the sector specific returns in response to shocks are

$$\hat{r}_N = \hat{p}_N + \hat{X}_N - \hat{K}_N \text{ and } \hat{r}_T = \hat{X}_T - \hat{K}_T, \tag{37}$$

for the N and T sectors, respectively.

The nine static relationships we have just stated, in the changes of the ten variables

$$\{X_N, X_T, L_N, L_T, K_N, K_T, p_N, r_N, r_T, w\} \tag{38}$$

plus the dynamic equation $\dot{K}_T = -\nu(r_N/r_T)$, which gives the rate of change of the traded sector capital stock with respect to time, specify the system.

The solution approach is first to use the comparative static equations to solve for the change in the traded sector labor force, $\hat{L}_T$, in terms of the change in the traded sector capital stock, $\hat{K}_T$. We then substitute the static equations for changes in outputs and labor forces in the sectors to obtain a solution for $\hat{p}_N$, the static change of the real exchange rate. Finally, we will use the dynamic equation for $\dot{K}_T$, in combination with the comparative static equations above, in order to derive the changes of the variables as time evolves. This will allow us to verify that the system is indeed dynamically stable.

Proceeding along those lines, use the the labor market and capital market conditions to solve for $\hat{L}_N$ in terms of $\hat{L}_T$, and $\hat{K}_N$ in terms of $\hat{K}_T$:

$$\hat{L}_N = -\frac{\lambda_{LT}}{\lambda_{LN}}\hat{L}_T \text{ and } \hat{K}_N = -\frac{\lambda_{KT}}{\lambda_{KN}}\hat{K}_T \tag{39}$$

Using the wage equalization equation in changes, we can write the change in the real exchange rate as:

$$\hat{p}_N = \hat{X}_T - \hat{L}_T - \hat{X}_N + \hat{L}_N. \tag{40}$$

The change in the output of the N sector is terms of the changes in T sector capital and labor is:

$$\hat{X}_N = -\theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}\hat{K}_T - \theta_{LN}\frac{\lambda_{LT}}{\lambda_{LN}}\hat{L}_T. \tag{41}$$

Now, using these equations, and the equation for the change $\hat{X}_T$ in traded sector output, we can express $\hat{p}_N$ exclusively in terms of $\hat{K}_T$ and $\hat{L}_T$:

$$\hat{p}_N = \left(\theta_{KT} + \theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}\right)\hat{K}_T + \left(\theta_{LT} + \theta_{LN}\frac{\lambda_{LT}}{\lambda_{LN}} - \frac{1}{\lambda_{LN}}\right)\hat{L}_T \tag{42}$$

Now that we can express all variables in the comparative static equation (34) for the N goods market in terms of the two variables $\hat{L}_T$ and $\hat{K}_T$, we substitute the foregoing relationships into that equation and solve for $\hat{L}_T$ as a function of $\hat{K}_T$. This yields:

$$\hat{L}_T = B\hat{K}_T, \tag{43}$$

where the function

$$B \equiv \frac{(\sigma-\chi)\theta_{KT} - (1-\sigma)\theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}}{\sigma/\lambda_{LN} + (\chi-\sigma)\theta_{LT} + (1-\sigma)\theta_{LN}(\frac{\lambda_{LT}}{\lambda_{LN}})}. \tag{44}$$

The properties of this function are important and merit discussion. First, note that the denominator of B will be positive for $\sigma \leq 1$. Second, note however that there exists a value $\sigma^* < 1$ at which the numerator is equal to zero, with $B < 0$ for $\sigma < \sigma^*$. To be precise,

$$\sigma^* = \frac{\chi\theta_{KT} + \theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}}{\theta_{KT} + \theta_{KN}\frac{\lambda_{KT}}{\lambda_{KN}}} \tag{45}$$

This effect will have clear implications for the sign of inequality. Before we discuss that, let us first verify that the system is dynamically stable. The change that is driving the system is the movement of capital from the T sector to the N sector: $\dot{K}_T = -\nu(r_N/r_T)$. Thus we can view the static effects of changes in K_T as being parameterized by time t, with the rate of movement of K_T given by $-\nu$. With this understanding, let $K_T(t)$ denote the value of T sector capital at time t. As K_T changes, the changes in the all other variables of the model are governed by the static equations just derived. The system will be dynamically stable if

$$\frac{\dot{r}_N}{r_N} - \frac{\dot{r}_T}{r_T} < 0.$$

To verify this condition holds, substitute the dynamic versions of the equations for the changes in the sector specific returns to capital,

$$\frac{\dot{r}_N}{r_N} = \frac{\dot{p}_N}{p_N} + \frac{\dot{X}_N}{X_N} - \frac{\dot{K}_N}{K_N} \text{ and } \frac{\dot{r}_T}{r_T} = \frac{\dot{X}_T}{X_T} - \frac{\dot{K}_T}{K_T}, \tag{46}$$

into the above expression, and invoke the equalization of wages between sectors to write

$$\frac{\dot{r}_N}{r_N} - \frac{\dot{r}_T}{r_T} = \frac{\dot{L}_N}{L_N} - \frac{\dot{L}_T}{L_T} + \frac{\dot{K}_N}{K_N} - \frac{\dot{K}_T}{K_T}. \tag{47}$$

Next, use the static equation (39), the dynamic equation for $\dot{K}_T$, and the static relationship (43) between L_T and K_T, to write

$$\frac{\dot{r}_N}{r_N} - \frac{\dot{r}_T}{r_T} = \frac{\nu}{K_T}\left(\frac{B}{\lambda_{LN}} - \frac{1}{\lambda_{KN}}\right) < 0 \tag{48}$$

for $B < \lambda_{LN}/\lambda_{KN}$, which is guaranteed from the fact that $B < 1$ for $\sigma \leq 1$.

Now let us proceed to our inequality result. The differential equation governing inequality is

$$\frac{\dot{I}}{I} = \frac{\dot{r}}{r} - \frac{\dot{w}}{w}.$$

Using the static equation for the change in r,

$$\hat{r} = \left(\frac{\pi_N}{\pi}\right)\hat{\pi}_N + \left(\frac{\pi_T}{\pi}\right)\hat{\pi}_T, \tag{49}$$

we use the comparative static equations for the medium run, along with the dynamic equation for $\dot{K}_T$, to obtain

$$\frac{\dot{r}}{r} = \left[\theta_{KT} + (\theta_{LT} - (\frac{\pi_N}{\pi_N})(\frac{1}{\lambda_{LN}}))B\right]\frac{\dot{K}_T}{K_T} \tag{50}$$

Similarly, from the static equation for w,

$$\hat{w} = \hat{X}_T - \hat{L}_T,$$

we use the comparative static equation for T sector output and the comparative static equation (39) to write

$$\frac{\dot{w}}{w} = \theta_{KT}(1 - B)\frac{\dot{K}_T}{K_T} \tag{51}$$

With these results in hand, the the differential equation for inequality I can be written as

$$\frac{\dot{I}}{I} = \left[1 - \frac{\pi_N}{\pi}\left(\frac{1}{\lambda_{LN}}\right)\right] B \frac{\dot{K}_T}{K_T} \tag{52}$$

For inequality to be falling during the medium run, we need to have

$$\left[1 - \frac{\pi_N}{\pi}\left(\frac{1}{\lambda_{LN}}\right)\right] B > 0.$$

We first show that the term in brackets is positive if and only if the N sector is relatively labor intensive. To see this, we must prove that $\pi_N/\pi < L_N/L$ for $\theta_{LN} > \theta_{LT}$. This is established by rearranging the inequality to read $\frac{\pi_N}{L_N} < \frac{\pi}{L}$, and using the equation $\pi = \pi_N + \pi_T$, where $\pi_N = \theta_{KN} p_N X_N$ and $\pi_T = \theta_{KT} X_T$, to obtain the inequality

$$\theta_{KN}\frac{p_N X_N}{L_N} < \theta_{KN}\frac{p_N X_N}{L_N}\lambda_{LN} + \theta_{KT}\frac{X_T}{L_T}\lambda_{LT}$$

Rearrange this to obtain

$$\theta_{KN}\frac{p_N X_N}{LN} < \theta_{KT}\frac{X_T}{L_T},$$

where we have used the fact that $1-\lambda_{LN} = \lambda_{LT}$. Now use the fact that the wage $w = \theta_{LN} p_N X_N / L_N = \theta_{LT} X_T / L_T$ to obtain the result. We have already seen that $B > 0$ for σ sufficiently large, and $B < 0$ for σ sufficiently small. This gives us our result. ∎

Appendix 2: The Comparative Statics of a change in ϕ in the LR

When both labor and capital markets are in equilibrium, the allocations of labor and capital will lie along the contract curve. Any changes in the allocations as a result of shocks to $\phi = A_T/A_N$ must also lie along the contract curve, so that

$$\hat{L}_T - \hat{K}_T = \hat{L}_N - \hat{K}_N.$$

By construction, $\hat{\phi} = \hat{A}_T - \hat{A}_N$. Using the contract curve relationship and the comparative static equations for output,

$$\hat{X}_T = \hat{A}_T + \theta_{KT}\hat{K}_T + \theta_{LT}\hat{L}_T \text{ and } \hat{X}_N = \hat{A}_N + \theta_{KN}\hat{K}_N + \theta_{LN}\hat{L}_N,$$

the change in the real exchange rate can be written as

$$\hat{p}_N = \hat{X}_T - \hat{L}_T - (\hat{X}_N - \hat{L}_N) = \hat{\phi} + (\theta_{LN} - \theta_{LT})(\hat{K}_T - \hat{L}_T)$$

Denoting by x_T and x_N the T and N sector outputs that correspond to $A_T = A_N = 1$, for given capital and labor allocations, we can write the N sector goods market equilibrium equation in changes as

$$\hat{x}_N + \sigma \hat{p}_N = \hat{\phi} + \chi \hat{x}_T.$$

From the comparative static equations corresponding to labor and capital market clearing, combined with the contract curve relationship, we obtain

$$\hat{K}_T = \frac{\lambda_{KN}}{\lambda_{LN}} \hat{L}_T.$$

Using the fact that

$$\hat{x}_T = \theta_{KT}\hat{K}_T + \theta_{LT}\hat{L}_T \text{ and } \hat{x}_N = \theta_{KN}\hat{K}_N + \theta_{LN}\hat{L}_N,$$

we thus obtain

$$\hat{x}_T = \left(\theta_{KT}\frac{\lambda_{KN}}{\lambda_{LN}} + \theta_{LT}\right)\hat{L}_T \text{ and } \hat{x}_N = -\left(\theta_{KN}\frac{\lambda_{KT}}{\lambda_{LN}} + \theta_{LN}\frac{\lambda_{LT}}{\lambda_{LN}}\right)\hat{L}_T,$$

From this we see immediately that $\hat{x}_N - \chi\hat{x}_T$ will be equal to a negative coefficient times $\hat{L}_T$. Substituting our equation for $\hat{p}_N$ into the N good market comparative static equation and rearranging, we can write

$$\hat{x}_N - \chi\hat{x}_T + \sigma(\theta_{LN} - \theta_{LT})(\frac{\lambda_{KN} - \lambda_{LN}}{\lambda_{LN}})\hat{L}_T = (1-\sigma)\hat{\phi}.$$

The LHS is a negative coefficient times $\hat{L}_T$. The RHS is positive for $\phi < 1$ and negative for $\phi > 1$. Thus we have proved that $d\lambda_{LN}/d\phi > 0$ for $\sigma < 1$ and $d\lambda_{LN}/d\phi < 0$ for $\sigma > 1$, as required.

Appendix 3: Data description and sources

Inequality Gini index of gross household income inequality, taken from The University of Texas Inequality Project (variable "EHII2.3", see Galbraith and Kum (2005) for details).

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

Political constraints indicator based on the number of independent branches of government with veto power over policy change, taken from Henisz (2000) (variable "POLCONV").

Primary schooling average years of primary schooling of the population aged 15 and over, taken from Barro and Lee (2000). Since this variable is only available for 1960, 1965, 1970, 1975, 1980, 1985, 1990, 1995, and 1999, we fill in the missing years by linear interpolation.

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 (rescaled so that 1980 = 100) index by the share of commodity exports in GDP (GDP in current US dollars, WDI). The sub-indices for non-agricultural and agricultural commodities are constructed in the same way.[13]

[13] 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 4: Panel unit root and panel cointegration tests

The long-run equilibrium relationship between inequality and its determinants, which underlies the error-correction specification in equation (28), can be written as follows[14]:

$$y_{i,t} = -\frac{1}{\lambda}(\gamma_i + \theta' \tau_{i,t} + \beta_1' x_{i,t} + u_{i,t}) \tag{53}$$

where γ_i is a country-specific fixed effect and $\tau_{i,t}$ is an $N \times 1$ vector of N country-specific time trends.[15] The error-correction specification in equation (28) is appropriate if $y_{i,t}$ and $x_{i,t}$ are cointegrated, i.e. have a common stochastic trend which is cancelled out by the linear combination. This requires that the individual variables are integrated of order 1, i.e. non-stationary in levels and stationary in first differences, and that the residuals of a regression of $I_{i,t}$ on $X_{i,t}$ are stationary. To test this, we first performed panel unit root tests on the levels and the differences of the variables in $y_{i,t}$ and $x_{i,t}$ and then performed a panel cointegration test. The results are reported in Table A.1.[16] 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). Under the null hypothesis all groups have a unit root. Rejection of the null implies that one or more groups do not have a unit root. The difference between the IPS and MW tests is that the first is parametric and uses the t-statistics of the individual unit root tests, while the second is non-parametric and uses the p-values of the individual unit root tests. While IPS can be applied to balanced panels only, MW can be used for both balanced and unbalanced panels. The top part of Table A.1 reports the IPS test results for a balanced sub-sample of 17 countries and 35 years of data, and the MW test results for both the sub-sample and the full sample. The test results support the hypothesis that the variables are integrated of order 1. The null of a unit root is always rejected for the differences and is not rejected for most of the level tests. Only for GDP per capita and political constraints, the MW test rejects the null of a unit root. However, this does not mean that all series in the panel are stationary, but that at least one of the series is stationary. It is therefore possible that

[14] This appendix draws heavily from Collier and Goderis (2007a).

[15] Since we left the country-specific fixed effect in equation (28) unrestricted, it captures a country-specific constant in both the levels and the differenced equation. The first is represented by γ_i in equation (53). The second implies a country-specific linear time trend in the levels equation (53), which is captured by $\theta' \tau_{i,t}$. We leave the regional time dummies out of equation (53) to limit the number of regressors in the cointegration test.

[16] Although we initially allow for a long-run effect of commodity export prices on inequality in our estimations, we do not find any evidence of such an effect. Therefore, the vector of cointegrating variables in our unit root and cointegration tests does not include the commodity export price index.

the tests reject non-stationarity while most of the series are in fact non-stationary. To investigate this possibility, we performed augmented Dickey-Fuller tests for individual countries. Table A.1 reports the number of countries for which these tests reject the null of stationarity at 5%, as a ratio of the total number of countries in the sample. For the vast majority majority of countries, the ADF tests do not reject a unit root for the level variables, but do reject a unit root for the differenced variables. We can therefore assume that the variables are integrated of order 1.

We next perform a panel cointegration test, as suggested by Pedroni (1999). We first run the regression in equation (54) for each country in our sample separately:

$$y_t = \alpha_0 + \alpha_1 t + \alpha_2 x_t + \varepsilon_t \tag{54}$$

where t is a time trend. This specification allows for country-specific fixed effects, country-specific time trends and country-specific coefficients for the long-run determinants of inequality (GDP per capita, political constraints, and primary schooling). We collect the residuals from these regressions and run ADF regressions for each country. Following Pedroni (1999), we allow the lag order of the dependent variable in the ADF regressions to vary across countries by including the lags that are statistically significant at 5 percent. We then calculate the mean ADF t-statistic, derive the group t-statistic, and express it in the form of equation (2) on p. 665 in Pedroni (1999). The bottom part of Table A.1 reports this standard normally distributed group t-statistic. We reject the null hypothesis of no cointegration at 1 percent significance and thus conclude that the variables are cointegrated.[17]

[17] Baltagi and Kao (2000) explain the intuition behind the panel cointegration test: "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".

Table A.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	
Inequality	-1.39	0/17	-3.89***	13/17	18.03	0/17	219.8***	13/17	160.6	6/86	964.4***	63/83
GDP per capita (log)	-1.04	1/17	-3.63***	13/17	24.58	1/17	173.0***	13/17	251.4***	11/88	747.5***	54/88
Political constraints	-0.25	2/16	-4.72***	16/16	47.73*	2/16	489.0***	16/16	272.6***	8/73	1857.7***	65/74
Primary schooling	-	-	-	-	-	-	-	-	-	-	-	-
	Panel cointegration test											
	Pedroni, full sample											
Group t-Statistic, N(0,1)	-2.77***											

Notes: Table A.1 reports the results of the panel unit root and panel cointegration tests. For the panel unit root tests, we report 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 (53) includes a time trend, we also ran the panel unit root tests for inequality 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. The number of lags is 1, 1, and 0 for inequality, GDP per capita, and political constraints, respectively. Primary schooling was constructed using interpolation to fill in missing observations. As a result, panel unit root tests are inappropriate. *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively.

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Table 1a Commodities

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

n = non-agricultural commodities

Table 1b Sample countries and their shares of commodity exports in GDP

Zambia (0.35)	Cameroon (0.15)	Senegal (0.06)	Philippines (0.03)	Hong Kong (0.01)
Venezuela (0.32)	Mauritius (0.15)	Burundi (0.06)	South Africa (0.03)	France (0.01)
Liberia (0.30)	Syria (0.15)	Peru (0.06)	Poland (0.03)	Spain (0.01)
Congo, R. (0.29)	Algeria (0.15)	El Salvador (0.06)	Ireland (0.03)	India (0.01)
Swaziland (0.22)	Togo (0.14)	New Zealand (0.06)	Greece (0.02)	Turkey (0.01)
Singapore (0.21)	Iran (0.14)	Kenya (0.06)	China (0.02)	Haiti (0.01)
Ecuador (0.21)	Bolivia (0.12)	Jordan (0.05)	Egypt (0.02)	US (0.01)
Malaysia (0.21)	Fiji (0.12)	Canada (0.05)	Brazil (0.02)	Israel (0.01)
Malawi (0.20)	Ghana (0.11)	Panama (0.05)	Pakistan (0.02)	Korea (0.01)
Honduras (0.20)	Costa Rica (0.11)	Netherlands (0.05)	Portugal (0.02)	Bangladesh (0.01)
Tr. & Tob. (0.19)	Colombia (0.11)	Thailand (0.04)	Hungary (0.02)	Austria (0.01)
Iceland (0.19)	Zimbabwe (0.09)	Uruguay (0.04)	Barbados (0.02)	Italy (0.00)
P. N. Guin. (0.18)	Guatemala (0.08)	Mexico (0.04)	C. Afr. Rep. (0.02)	Kuwait (0.00)
Jamaica (0.18)	Dom. Rep. (0.07)	The Gambia (0.04)	Finland (0.02)	Lesotho (0.00)
Nicaragua (0.17)	Sri Lanka (0.07)	Uganda (0.04)	UK (0.02)	Nepal (0.00)
Norway (0.17)	Australia (0.07)	Rwanda (0.04)	Cyprus (0.02)	Benin (0.00)
Chile (0.16)	Botswana (0.07)	Denmark (0.04)	Sweden (0.02)	Belgium (0.00)
Indonesia (0.15)	Tunisia (0.06)	Argentina (0.03)	Mozambique (0.01)	Japan (0.00)

Table 2 Summary statistics

	Obs.	Mean	St. Dev.	Min.	Max.
Household income inequality (Gini index)	1988	41.39	6.70	24.07	59.09
Δ Household income inequality (Gini index)	1988	0.11	1.44	-14.72	11.06
Real GDP per capita (log)	1988	8.02	1.48	4.69	10.73
Political constraints	1988	0.43	0.33	0.00	0.89
Primary schooling (average number of years)	1988	3.94	1.63	0.47	7.70
Commodity export price index	1958	0.33	0.33	0.00	1.87
Unlogged unweighted index (1980=100)	1958	85.70	26.88	17.60	224.48
Δ Commodity export price index	1958	-0.00	0.02	-0.23	0.31
Δ Unlogged unweighted index (1980=100)	1958	-0.79	13.02	-103.10	76.40
Commodity exports to GDP (ratio)	1988	0.08	0.08	0.00	0.35
Δ Non-agricultural commodity export price index	1958	-0.00	0.02	-0.23	0.31
Δ Unlogged unweighted non-agric. index (1980=100)	1958	-0.64	13.11	-116.89	88.35
Non-agricultural commodity exports to GDP (ratio)	1988	0.04	0.07	0.00	0.34
Δ Agricultural commodity export price index	1958	-0.00	0.01	-0.07	0.06
Δ Unlogged unweighted agric. index (1980=100)	1958	-1.08	15.09	-87.06	88.86
Agricultural commodity exports to GDP (ratio)	1988	0.03	0.05	0.00	0.22

Table 3 Estimation results: baseline specifications

	(1)	(2)	(3)	(4)	(5)
			Estimates of long-run coefficients		
GDP per capita (log)	-4.02***	-4.12***	-4.11***	-4.11***	-4.12***
	(1.49)	(1.48)	(1.48)	(1.49)	(1.48)
Political constraints	-3.02**	-3.15**	-3.14**	-3.14**	-3.14**
	(1.26)	(1.25)	(1.26)	(1.26)	(1.25)
Primary schooling	-1.45*	-1.52*	-1.51*	-1.52*	-1.52*
	(0.82)	(0.86)	(0.82)	(0.82)	(0.82)
Commodity export price index		0.48			
		(6.60)			
			Estimates of short-run coefficients		
Inequality$_{t-1}$	-0.20***	-0.20***	-0.20***	-0.20***	-0.20***
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
Δ Inequality$_{t-1}$	-0.14**	-0.14**	-0.14**	-0.14**	-0.14**
	(0.06)	(0.06)	(0.06)	(0.06)	(0.06)
Δ Commodity export price index$_t$		-2.40	-2.45		
		(1.75)	(1.62)		
Δ Non-agricultural export price index$_t$				-3.23**	-3.26**
				(1.53)	(1.54)
Δ Agricultural export price index$_t$				-2.88	
				(6.93)	
Country fixed effects	YES	YES	YES	YES	YES
Regional time dummies	YES	YES	YES	YES	YES
Observations	1988	1958	1958	1958	1958
R-squared (within)	0.23	0.23	0.23	0.23	0.23

Notes: The dependent variable is the change in inequality in year t. Robust standard errors are clustered by country and are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels.

Table 4 Estimation results: sensitivity analysis

	(1)	(2)	(3)	(4)	(5)
		Estimates of long-run coefficients			
GDP per capita (log)	1.75	-4.13***	-2.29	-5.22***	-5.69***
	(2.20)	(1.48)	(1.70)	(1.19)	(1.24)
Political constraints	0.07	-3.13**	-0.21	-2.46**	-3.40***
	(2.20)	(1.25)	(3.80)	(0.98)	(1.02)
Primary schooling	1.79**	-1.53*	-0.17	-1.28**	-1.17*
	(0.86)	(0.82)	(0.93)	(0.63)	(0.70)
		Estimates of short-run coefficients			
Inequality_{t-1}	-0.13***	-0.20***	-0.05***	-0.22***	-0.23***
	(0.03)	(0.03)	(0.01)	(0.02)	(0.02)
Δ Inequality_{t-1}	-0.15**	-0.14**	-0.19***	-0.07***	-0.15***
	(0.06)	(0.06)	(0.06)	(0.03)	(0.03)
Δ Non-agri export price index_t	-2.37**		-3.75**	-4.23	-4.30
	(1.18)		(1.73)	(2.97)	(2.62)
Δ Non-agri export price index_t-positive		-4.16**			
		(1.77)			
Δ Non-agri export price index_t-negative		-2.24			
		(2.93)			
Method	FE	FE	OLS	FE-IV	FE-IV
Regional time dummies	NO	YES	YES	YES	YES
Observations	1958	1958	1958	1802	1462
R-squared (within)	0.09	0.23	0.17	0.24	0.29

Notes: The dependent variable is the change in inequality in year t. Columns (4) and (5) report the second-stage results of an instrumental variables regression in which we instrument for the change in the non-agricultural commodity export price index, using World Bank estimates of sub-soil assets in 2000 and 1994, respectively. Standard errors are reported in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels.

Figure 1 The Edgeworth-Bowley box with dynamically stable capital and labor allocations

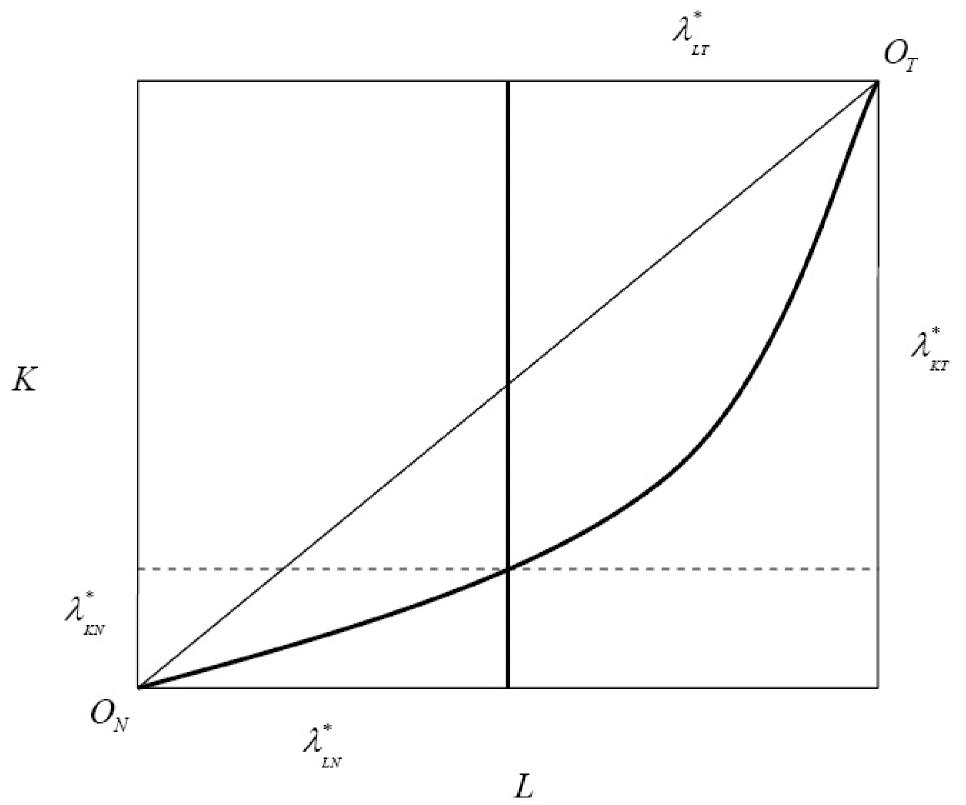

Figure 2 The time path of inequality following a resource boom

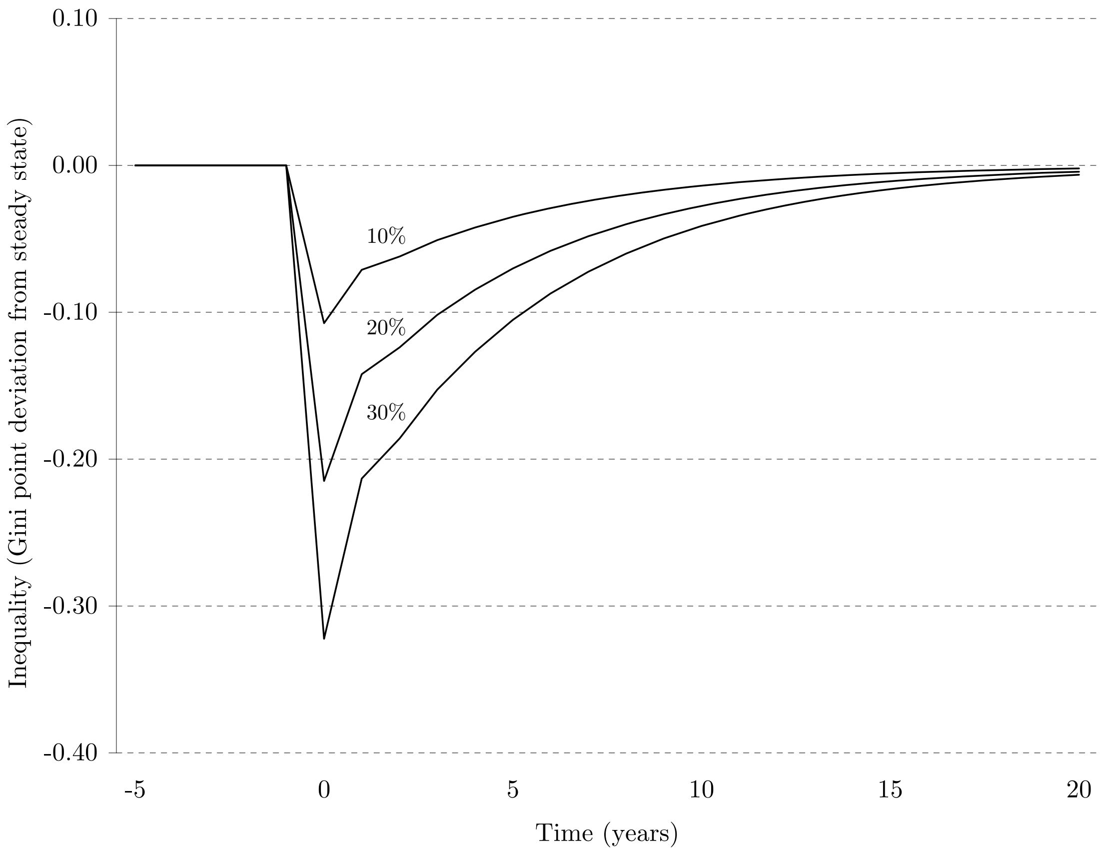

Figure 2 is based on the estimation results in Table 3, column (5), and shows the impulse response functions (IRF's) of inequality to a two standard deviation (33 % points) shock to the growth rate of non-agricultural commodity export prices in period 0. The three IRF's correspond to the cases in which the share of non-agricultural commodity exports in GDP is 10%, 20%, and 30%, respectively.