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Abstract

While much ink has been spilled over the study of income inequality and the Dutch disease in isolation from each other, little attention has been paid to the association between these subjects of interest. From this perspective, the present paper develops a two-sector growth model including two groups of workers (skilled and unskilled) with different consumption baskets. The model is induced by a relative real wage between sectors and between workers in the short-term (comparative static), while it is driven by the relative productivity growth and also a change in the relative consumption expenditure, resulting from an income inequality change, in the long-term. The main findings are twofold. First, a natural resource boom reduces income inequality if the relative real wage of skilled to unskilled workers is stronger than their relative share on windfall income benefit (subsidies). Second, falling income inequality exacerbates the intensity of the Dutch disease if skilled workers, with respect to unskilled workers, allocate a larger expenditure share for traded goods. Using the dynamic panel data approach for a sample of 79 countries over the period 1975-2014, I evaluate the theory's predictions. The empirical study represents some clear evidence in supporting the crucial role of income inequality in the economic performance of resource-dependent countries.

Keywords: Windfall Income; The Dutch Disease; Income Inequality

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1 Introduction

Resource economies have usually failed to show better economic performance than the others (Frankel, 2010). In practice, these economies exhibit slower growth (Sachs and Warner, 1995) and higher levels of income inequality (Boyce and Ndikumana, 2012). The questions that have attracted growing attention of researchers are the impact of windfall income on economic growth and income inequality. Surprisingly, these variables of interest have been studied in isolation from each other and a little attention has been paid to study these variables in a unified framework. Literature reveals that economic growth and income inequality are endogenous variables and their co-movements affect the underlying economic forces to which they are both responding (Turnovsky, 2011) [1]. Therefore, these variables of interest need to be simultaneously studied since their relationship seems to be associative and not causal (Ehrlich and Kim, 2007).

The main contribution of this paper stems from the question of how the impact of a resource boom on the income inequality-growth nexus modifies the standard view on the Dutch disease. In this vein, I investigate first the impact of a resource boom on income inequality, and then the feedback of a change in income inequality on sectoral growth rate (i.e. the Dutch disease), in terms of both theory and empirical evidence.

A useful starting point for our discussion is to ask why do countries with resource wealth tend to grow less rapidly? A conventional answer refers to the theory of the Dutch disease. The main idea surrounding the original theory, proposed by (Corden and Neary, 1982), rests on the following triple-step reasoning: in a comparative static, a natural resource boom, first, increases the marginal product of labor in the natural resource sector and leads to rise the relative real wage between sectors. Accordingly, labor forces move from both manufacturing and non-traded sectors to the natural resource sector (i.e. the so-called resource movement effect). Second, the windfall income raises the national income and so tends to increase the demand for imported goods and the domestic absorption for both traded and non-traded goods. Third, the real exchange rate (the relative price of non-traded to traded goods) appreciates to confront the expanded demand for non-traded goods. Consequently, labor forces shift away from the manufacturing sector and into the non-traded sector to respond the gap between supply and demand sides (i.e. the so-called spending effect). Briefly, a natural resource boom leads to worsening the competitiveness of non-resource sectors through an increase in the real exchange rate and thus shrinking the traded sector.

Later on, Sachs and Warner (1995); Torvik (2001) challenged this strand of the literature by

[1] The possible impact of income inequality on economic growth has received more attention in recent decades. The empirical evidence for this case is inconclusive. For instance, Persson and Tabellini (1994); Alesina and Rodrik (1994); Castelló and Doménech (2002); Easterly (2007) and Assa (2012) all suggest that income inequality has a negative impact on the growth rate, while Perotti (1996); Li and Zou (1998); Forbes (2000) and Naguib (2015) show that an increase in income inequality accelerates the rate of economic growth. In contrast, Banerjee and Duflo (2003) show that the growth rate is an inverted U-shaped function of net changes in inequality. Grijalva (2011) concludes that an increase in income inequality level declines the rate of economic growth, however, this effect seems to vanish over time. Further, Barro (2000, 2008) argue that a higher level of income inequality decelerates growth rate in developing countries while it accelerates the growth rate in developed countries.

developing dynamic versions of the Dutch-Disease model in which productivity growth was driven by learning by doing (*LBD*) [2]. This rests on the point that the traded sector benefits more from learning by doing and thus the non-resource traded sector hit by reducing competitiveness is not fully recovered once the resource income runs out (Van der Ploeg, 2011). The hypothesis of an adverse effect of resource dependence [3] on economic growth has been empirically supported by Sachs and Warner (1995, 2001), Rodriguez and Sachs (1999) and Gylfason et al. (1999) [4]. The second question, i.e. effect of a resource boom on inequality, has been investigated theoretically and empirically for decades. Bourguignon and Morrisson (1990) argue that the mineral resource endowment is one of the main determinants of income inequality in developing countries. According to Gylfason and Zoega (2002), resource dependence leads to both lower growth rates and higher inequality. In a simple model, they demonstrate that education, which raises the return to work through a higher productivity level, can simultaneously enhance both equality and growth and thereby reduces the adverse effects of resource rents on these economies. Briefly, investment on education may help economic growth and reduce inequality. Goderis and Malone (2011) propose a two-sector growth model in which two kinds of labor, skilled and unskilled, under a learning by doing model (*LBD*) drive economic growth. They suppose that the traded sector is relatively more skilled labor intensive than the non-traded sector. Changes in income inequality are driven by distribution of the resource income as well as factor reallocation across sectors. Their theoretical findings have been supported by a panel data approach on a dataset including 90 countries over the period 1965-1999. They show that income inequality falls in the short term immediately after a resource boom and then rises steadily over time until the initial impact of the resource boom disappears. In this vein, the impact of the oil rent on income inequality has also been investigated by Mallaye et al. (2015). They study a dynamic panel data model on a dataset including 40 developing countries over the period 1996-2008. Their findings suggest that the oil rent reduces inequality in the short run, while this effect vanishes over time as the oil revenue increases. Using pooled *OLS* regressions on an unbalanced panel of 55 country observations for the years 1975-2008, Parcero and Papyrakis (2016) found that oil is associated with lower income inequality for economies with moderate levels of oil dependence/abundance

[2] Whereas for the former, Sachs and Warner (1995), *LBD* is generated in the traded sector with a perfect spillover to the non-traded sector, for the latter, Torvik (2001), both traded and non-traded sectors contribute the learning process with a spillover between two sectors.

[3] Two different criteria are usually used to assess the economies depending on the natural resource: "resource-dependence" refers to the value of the resource as a share of GDP or total national wealth and "resource-abundance" refers to the per capita value of the stock of natural resource wealth. The literature has voided the misconception that resource abundance should be interpreted as a rule that resource-rich countries are doomed to failure (Frankel, 2010). For example, comparing Sierra Leone and Botswana as two diamond-rich countries show that Botswana has expanded at an average rate of 7 % over the recent 20 years, while the growth rate of Sierra Leone has dropped 37 % between 1971 and 1989 (Humphreys et al., 2007). Moreover, recent empirical studies across a comprehensive sample of countries indicate that natural resource abundance plays a positive role in economic performance (see. Brunnschweiler and Bulte (2008); Alexeev and Conrad (2009); Esfahani et al. (2013); Cavalcanti et al. (2011)).

[4] The scarcity paradox may be observed in natural resource-rich countries due to keeping the energy price at a low level for a long period. In this vein, Bretschger (2015) shows that higher energy price through lower energy use leads to a reallocation of inputs toward capital accumulation that, in turn, makes the growth rate faster.

and greater income inequality for the very oil-dependent economies.

Although the literature has deepened our understanding of the Dutch disease and of the determinants on income inequality in the natural resource countries, surprisingly there are few serious attempts, as far as I know, to clarify how income inequality induced by a natural resource boom is associated with the economic growth [5]. Therefore, the main contribution of this paper is to clarify the role of income inequality on the standard Dutch disease theory.

In this vein, I extend the framework proposed by Goderis and Malone (2011) which in turn is based on the model developed by Torvik (2001). The core mechanism of the framework is learning by doing model (*LBD*) in which both sectors contribute to learning and there are learning spillovers between the sectors. My model departs from Goderis and Malone (2011) as follows. First, I will relax the core assumption of their model in which the traded sector is relatively skilled workers intensive. Second, productivity growth is driven by the learning-by-doing of skilled workers rather than unskilled workers. Further, I capture the feedback effect of income inequality on the Dutch disease by assuming different consumption baskets between workers' groups (the novel aspect of the model).

The theory predicts that in a transition path, the effect of windfall income on income inequality depends on the difference between the relative real wage across workers' groups and their relative share on windfall income benefit (e.g. subsidies). Moreover, the impact of income inequality on the intensity of the Dutch disease depends on how workers spend their income on traded and non-traded goods. I then present a structural empirical analysis to investigate whether the theoretical predictions are consistent with the empirical findings. In this respect, a dynamic panel data regression is applied. I first collect available data for 79 countries over the period 1975-2014 and then estimate the core findings of the theory by the generalized method of moments (GMM). Using regressions for a sample database, I find some clear evidence in supporting the theoretical findings. Briefly, the empirical estimations demonstrate that on average income inequality reduces as a natural resource-dependence proxy increases and also falling income inequality is associated with a larger deceleration in sectoral economic growth.

The rest of this paper is organized as follows: Section 2 presents the model and its analysis in the short-run (comparative static) and long-run (dynamics). Section 3 undertakes the empirical study to examine the theoretical predictions. Section 4 concludes the paper.

[5] Scognamillo et al. (2016) present an empirical study to analyze a sample database including 43 countries over the period 1980-2012. The findings show that the resource-dependence among high-income countries is negatively correlated with the Gini coefficient and the correlation between resource dependence and per capita GDP is insignificant, while the resource-dependence among low-income countries is associated with a higher level of Gini coefficient and a lower level of per capita GDP. Further, Behzadan et al. (2017) highlight the importance of resource rent distribution in creating the Dutch disease. Their empirical analysis suggests that less equality in the distribution of natural resource rents leads to arising stronger Dutch disease effects.

2 The model

Consider a two-sector economy: the traded and non-traded sectors, which are respectively indexed by T and N. I assume there is no asset accumulation and the windfall income R is a constant exogenous gift over time, resulting from discovery or a resource price appreciation. The model consists of two groups of households: skilled (S) and unskilled (L) workers. Each of whom, populated by a continuum of symmetric-identical households, supplies labor inelastically. Each group's labor supply is normalized to unity.

$$S_T + S_N = 1 \tag{1a}$$

$$L_T + L_N = 1 \tag{1b}$$

S_T and S_N represent the skilled labor force in the traded and non-traded sectors, respectively. Similarly, L_T and L_N denote the unskilled labor force in the traded and non-traded sectors. To put some structure on the analysis, I assume labor can move freely across sectors. Moreover, production in each sector (that is X_T and X_N) employs both skilled and unskilled labor forces and operates under constant returns to scale. For simplicity, a *Cobb-Douglas* production function is assumed in each sector.

$$X_T = A_T S_T^{\beta} L_T^{1-\beta} \tag{2a}$$

$$X_N = A_N S_N^{\alpha} L_N^{1-\alpha} \tag{2b}$$

A_J, $J = \{T, N\}$ denotes total factor productivity in sector J. The price of traded goods is normalized to unity. Thereby, the price of non-traded goods, denoted by P, is identified as the real exchange rate. The economy's total income (Y) is given by the value of produced traded and non-traded goods plus the value of windfall income.

$$Y = X_T + PX_N + A_T R \tag{3}$$

A windfall income boom will be considered as an increase in R. As in Torvik (2001) and Goderis and Malone (2011), the value of windfall income is indexed on productivity in the traded sector. This assumption allows us to prevent windfall income from vanishing relative to national income as the economy grows [6].

On the demand side, I make the assumption that preferences differ by groups of households. A representative household of group $i = S, L$ maximizes a *CES* utility function in his consumption of traded and non-traded goods subject to its budget constraint $(PC_N^i + C_T^i = Y^i)$. The utility

[6] Two alternatives are to measure the real value of windfall income in the productivity unit of the non-traded sector or in a given fraction of total income. However, the steady-state solution is independent of these choices (Torvik, 2001).

function is given by:

$$U^i\left(C_N^i, C_T^i\right) = \left[(1-\theta_i)^{\frac{1}{\sigma}}\left({C_T}^i\right)^{\frac{\sigma-1}{\sigma}} + (\theta_i)^{\frac{1}{\sigma}}\left({C_N}^i\right)^{\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}} \quad ; \quad i = S, L. \tag{4}$$

σ is the elasticity of substitution between traded and non-traded goods. θ_i represents the consumption share of non-traded goods for group $i = S, L$ and Y^i is total income for the representative household of group i. The demand for traded and non-traded goods is denoted by C_T^i and C_N^i. The aggregate price index of each group of households $e_i(P)$ is determined by the solution of consumer's problem. As in Obstfeld et al. (1996), it is given:

$$e_i(P) = \left[(1-\theta_i) + \theta_i P^{1-\sigma}\right]^{\frac{1}{1-\sigma}} \quad ; \quad i = S, L. \tag{5}$$

Households' groups allocate their total income for traded and non-traded goods according to [7]:

$$C_N^i = \theta_i\, e_i^{\sigma-1}\, P^{-\sigma} Y_i \quad ; \quad C_T^i = (1-\theta_i)\, e_i^{\sigma-1}\, Y_i \quad ; \quad i = S, L. \tag{6}$$

In contrast with Goderis and Malone (2011) in which households have identical tastes, this model implies that the consumption distribution plays a key role in the economy's response to a resource boom. Finally, total consumption expenditure (C) equals the skilled workers' consumption expenditure (C^S) plus the unskilled workers' consumption expenditure (C^L). As regards there is no saving, the budget constraint of the entire economy is given by the equality between aggregate income and aggregate consumption expenditure (i.e. $Y = C = C^S + C^L$). To close the model, the market for non-traded goods must clear ($X_N = C_N$). The current account balance will always be in equilibrium by Walras' law.

Before moving to the description of the economy's response to an increase in the windfall income, a note about income inequality seems to be useful. There are two main sources leading to a change in income inequality between skilled and unskilled workers. The first is the relative factor intensity between sectors and the second is the relative share of windfall income benefit. Given that labor is mobile across sectors, an equalization between the sectoral marginal product of each factor gives the real wage of skilled workers (w_S) and unskilled ones (w_L),

$$w_S = \frac{\partial X_T}{\partial S_T} = P\frac{\partial X_N}{\partial S_N} \qquad w_L = \frac{\partial X_T}{\partial L_T} = P\frac{\partial X_N}{\partial L_N}. \tag{7}$$

Regarding the second driver of income inequality change, I make the assumption that windfall income benefits are directly distributed between household's groups (e.g. subsidies on energy price, education, and health care services). Skilled workers appropriate a share of the windfall

[7] For especial case $\sigma = 1$, we have a log-linear utility function as follow:
$U\left(C_N^i, C_T^i\right) = \left[\theta_i \log C_N^i + (1-\theta_i)\log C_T^i\right] \quad ; \quad i = S, L.$
The standard result is that each group of households allocates a given fraction of their total income for traded and non-traded goods. $C_T^i = (1-\theta_i) Y^i \qquad PC_N^i = \theta_i Y^i \qquad i = S, L$

income benefit, equal to π, while the remaining share $1-\pi$ accrues to unskilled workers. Therefore, income inequality which is also read as consumption expenditure inequality is defined as the ratio of skilled workers' income to total income,

$$I \equiv \frac{w_S + \pi A_T R}{Y} = \frac{C^S}{C}. \tag{8}$$

The income inequality definition allows to capture in a simple way the feedback of a change in income inequality on the economic performance of a resource-dependent economy. Using Equations 6 and 8, we have [8]:

$$\frac{P^{\sigma} C_N}{C} = \theta_L e_L^{\sigma-1} + \left[\theta_S e_S^{\sigma-1} - \theta_L e_L^{\sigma-1}\right] I. \tag{9}$$

It is important to note that the composition of demand plays no role in the response of an economy to a windfall income boom if skilled and unskilled workers have identical tastes (i.e. $\theta_S = \theta_L$). If tastes differ (i.e. $\theta_S \neq \theta_L$), a change in income inequality shifts demand away from one good and into another good. For instance, if $\theta_S > \theta_L$, Equation 9 indicates that rising income inequality raises the expenditure share of non-traded goods.

This structure of the model implies that there are three dimensions at work to analyze the economic performance of a resource-dependent country. The first one is factor intensity. The case where the traded sector is relatively skilled workers intensive ($\beta > \alpha$) seems to be consistent with countries exporting raw materials and having low value-added services or a large informal service sector. These countries are normally classified as lower-middle-income or low-income countries. Nevertheless, the case where the non-traded sector is relatively skilled workers intensive ($\beta < \alpha$) is inconclusive in resource-developed economies with high value-added services and low absorptive capacity constraints [9]. These countries are normally classified as upper-middle-income or high-income countries. Time-series evidence in high income countries shows that labor and value-added shares in the service sector increase along the development process (Herrendorf et al., 2013). These also coincide with a well-known upward trend between the human capital index and income level (Benhabib and Spiegel, 1994). Hence, the assumption that the non-traded sector is relatively more intensive in skilled workers seems to be plausible in high-income countries. But a large share of skilled-intensive services in high-income countries is export-oriented (e.g. services in finance, business, communication) and must be classified as traded goods, not as non-traded goods. As a conclusion, the assumption that the traded sector is relatively skilled workers intensive is more likely reliable for most of the countries. However, the other case is interesting to be discussed.

[8] For a linear utility function (i.e. $\sigma = 1$), the ratio of non-traded goods' expenditure to total expenditure is equal to $\frac{PC_N}{C} = \theta_L + (\theta_S - \theta_L) I$.

[9] The absorptive capacity refers to the declining marginal rate of return to aid (windfall income in this paper) as the value of aid increases. Given constraints on capacity such as a shortage of human capital and infrastructure (i.e. public goods), the unit cost of additional public goods and services rise and so the incremental returns to aid fall (Bourguignon and Sundberg, 2006).

The distribution of natural resource benefits among workers' groups is the second dimension of the model. Empirical evidence states that institutional quality [10] and political power may be key determinants of how windfall income is distributed between income classes. A natural recourse rent is more likely distributed unevenly in a poor institution country, while it might be distributed fairly in others (e.g. Torvik (2002) and Chaturvedi (2016)). In addition, International Energy Agency reported that major energy subsidizers are oil-exporter countries (IEA, 2019) [11]. In the same vein, there are many evidences to confirm an unequal distribution of natural resource benefit (e.g. subsidies) between income groups. Sdralevich et al. (2014) show that the poorest quintiles of some *MENA* countries (Egypt, Jordan, Mauritania, Morocco, and Yemen) receive about 1-7% of total diesel subsidies in 2012 compared to the richest quintiles benefit of about 42-77%. Gaddah et al. (2015) also examined the incidence of public education subsidies in Ghana in 2005. They found that the poorest quintile gains less than 15% of total education benefits, while the richest quintile benefit is more than 25%. Further, the total education benefits are going less to the poor than the rich in Indonesia at 1989 and Cote d'Ivoire at 1995 (Demery, 2000), increasing from 15% for the lowest quintile to 29% for the highest quintile in Indonesia and from 13% for the lowest quintile to 35% for the highest quintile in Cote d'Ivoire. Tiongson et al. (2003) have done a comprehensive benefit incidence analysis on education and health spending over a period of 1960-2000. Their analysis of education spending for 37 developing and transition economies represents that, on average, 27% of the benefits accrue to the richest quintile compared to 16% for the poorest quintile. Also, their analysis of health spending for 26 economies shows that about 23% of benefits accrue to the richest quintile, while the poorest quintile receives only about 17% of benefits [12].

The third dimension of the model is the gap in consumption expenditure shares for non-traded goods between workers' groups. The hypothesis suggesting that the poor with respect to the rich spend more for the non-traded goods (e.g. housing, transportation, education, and health care) and less for the traded goods (e.g. food and clothing) seems to be consistent with stylized facts [13]. The literature on the difference in unit price for homogeneous commodities have confirmed that the poor allocate lower unit value expenditure for good purchase (i.e. traded goods) (Leibtag et al., 2003; Broda et al., 2009). In contrast, the literature on health spending (i.e. non-traded

[10] Mehlum et al. (2006) argue that the resource is a blessing when institutions are good and is a curse when institutions are bad.

[11] The report reveals that Iran is the largest energy subsidizers and China, Saudi Arabia, Russia, India, Indonesia, Egypt, Algeria, Venezuela, Iraq and Kazakhstan are in the next ranks.

[12] Tiongson et al. (2003) also suggest the analysis of public spending on education and health for five regional groups: sub-Saharan Africa, Asia, and Pacific, Western Hemisphere (excluding Canada and the United States), Middle East and North Africa, and transition economies. In all country groups, the public spending benefit on education and health is appropriated less by the poor than by the rich. Only in Western Hemisphere, the public spending benefit on health is going more to the poor than the rich.

[13] According to a Consumer Expenditure Survey for the *US* economy in 2014, conducted by the Bureau of Labor Statistics, Low-income households allocated 81% of their budgets on basic needs (housing, transportation, health care, food, and clothing). While High-income households spent only 66% of their budgets on basic needs. Additionally, about 80% of their budgets on basic needs was allocated for non-traded goods (about 64% of the poor's budget and about 53% of the rich's budget).

goods) demonstrates that the poor pay a larger share of their income for health care. For example, in Thailand, the health expenditure share for the poor is 21% of their budget, whereas it is 2% for the rich (Pannarunothai and Mills, 1997). Further, expenditure share on health care in Sierra Leone is an average of 6.9%, decreasing from 25.6% for the lowest quintile to 3.7% for the highest quintile (Fabricant et al., 1999). Inequality in expenditure on health care seems to be valid even in the developed economies. A recent study for *US* economy (Ketsche et al., 2011) states that health-care spending consume more than 22% of total income for families in the lowest-income quintile, while it consumes less than 16% for families in any other income quintile.

In the same vein, literature on transportation spending and the housing expenditure share bring additional insights. In respect of the affordability of public transport, the household travel survey in *Delhi* (India) at 1994 find that the poorest quintile spends almost 15% of their income on public transport, while the highest quintile spends less than 10% of their income on public transport (Badami et al., 2004). Further, a survey, undertaken by the World Bank and the Center for Economic Studies at the Argentina Business University in 2002, shows that bottom quintile families allocate 31.6% of their income on travel to work, while top quintile families allocate 7.5% of their income on it (Carruthers et al., 2005). In respect of the housing expenditure share, a recent study in Germany over a period of 1993-2013 demonstrates an increase in expenditure share on housing for the lowest quintile (from 27% in 1993 to 39% in 2013) and a decrease for the highest quintile (from 16% in 1993 to 14% in 2013) (Dustmann et al., 2018).

Briefly, the evidence suggested for the gap in consumption expenditure shares might be feasible for most of the economies. Those seem to strongly support the assumption that unskilled workers (i.e. the poor) allocate a larger expenditure share for non-traded goods than skilled workers (the rich) do (i.e. $\theta_L > \theta_S$) [14].

In what follows, I discuss the economy's response to a windfall income boom in two stages. First in the short run when the productivity levels of both sectors are constant and second in the long run when productivity levels are driven by learning-by-doing (*LBD*).

2.1 Short run analysis (static model)

This section studies a comparative static response of an economy induced by a natural resource boom. A common assumption for this case is that productivity levels in both sectors are constant. Therefore, factor mobility across sectors will be the primary driver of the economy.

Given that the labor share of skilled workers drives the productivity levels (See. Section 2.2), I study the combination of the real exchange rate (P) and the labor share of skilled workers in the non-traded sector (S_N) to determine a static equilibrium of the model. The first combination of the variables of interest are found from the labor market (i.e. Equations 7) (henceforth *LL*-curve) and the second one is determined from the market-clearing condition in the non-traded sector (i.e.

[14] This assumption makes our analysis be consistent with the empirical predictions, presented in section 3.

$X_N = C_N$) (henceforth NN-curve). Defining the ratio of productivity levels $\phi \equiv \frac{A_T}{A_N}$, the results are respectively written as:

$$P = \phi \frac{\beta}{\alpha} \left(\frac{L_T}{S_T}\right)^{1-\beta} \left(\frac{S_N}{L_N}\right)^{1-\alpha} \qquad LL-curve \tag{10a}$$

$$P = \phi^{\frac{1}{\sigma}} \left(\frac{S_T^{\beta} L_T^{1-\beta} + R}{S_N^{\alpha} L_N^{1-\alpha}}\right)^{\frac{1}{\sigma}} \left[\Psi(P,I)\right]^{\frac{1}{\sigma}}. \qquad NN-curve \tag{10b}$$

Where,

$$\Psi(P,I) = \left(\frac{P^{\sigma} C_N}{C_T}\right) = \frac{\left[\theta_L e_L^{\sigma-1} + \left(\theta_S e_S^{\sigma-1} - \theta_L e_L^{\sigma-1}\right) I\right]}{1 - P^{1-\sigma}\left[\theta_L e_L^{\sigma-1} + \left(\theta_S e_S^{\sigma-1} - \theta_L e_L^{\sigma-1}\right) I\right]}. \tag{11}$$

LL-curve (Equation 10a) is an upward sloping curve. As in Torvik (2001), it says that for a given L_N and S_N, an increase in the real exchange rate makes the marginal productivity of skilled workers be larger in the non-traded sector than in the traded sector. Therefore, the labor share of skilled workers has to increase in the non-traded sector to re-establish the equilibrium in the labor market.

Another combination of P and S_N is NN-curve (Equation 10b). NN-curve is a downward sloping curve. This can be inferred as in Torvik (2001). For a given L_N and P, an increase in S_N enlarges the supply of non-traded goods. To re-establish the equilibrium in the non-traded market, the real exchange rate depreciates to expand demand for non-traded goods.

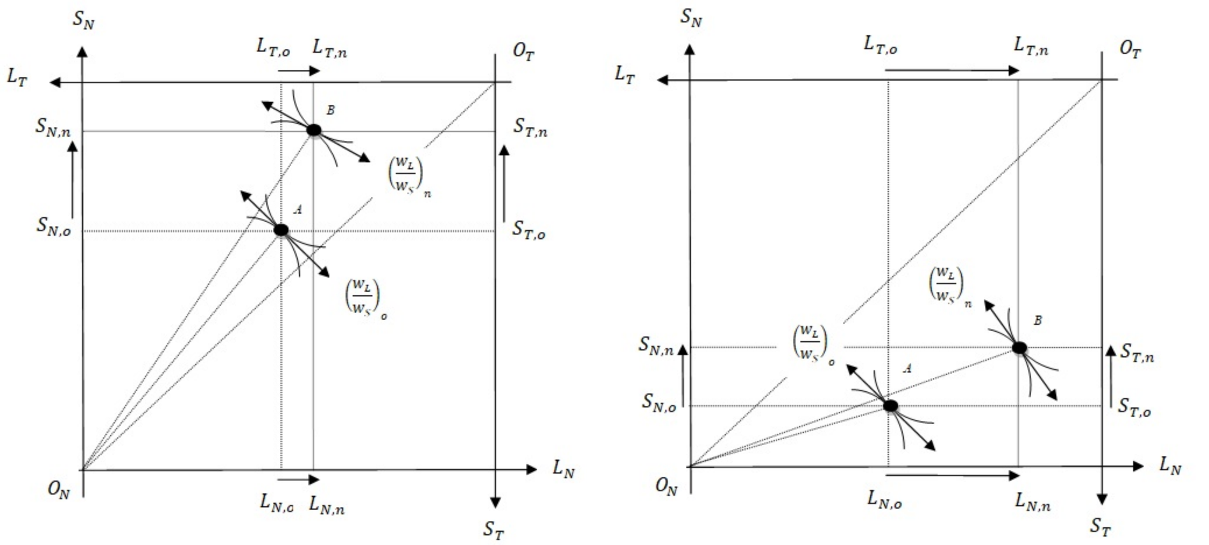
Before attention is turned to the static response of the model, let me look closely at Equation 11. $\Psi(P,I)$ highlights the key role of income inequality in the response of an economy to a natural resource boom. This is in contrast to Goderis and Malone (2011) in which $\theta_S = \theta_L$ [15]. Equations 11 is simplified to $\Psi(I) = \frac{\theta_L + (\theta_S - \theta_L)I}{1-[\theta_L + (\theta_S - \theta_L)I]}$ when $\sigma = 1$. Further, it indicates that the feedback of income inequality change on the real exchange rate depends on the gap in expenditure shares between skilled and unskilled workers. When skilled workers with respect to unskilled workers allocate a larger expenditure share for traded goods (i.e. $\theta_S < \theta_L$), rising income inequality depreciates the real exchange rate by shifting demand from non-traded to traded goods [16].

An increase in windfall income (R) expands total national income (Y) and so tends to increase the demand for both traded and non-traded goods. The expanded demand for traded goods will be compensated by more imported goods. While the real exchange rate (P) must appreciate in order to confront the expanded demand for non-traded goods. Graphically NN-curve shifts up, while LL-curve is not affected by the windfall income boom. The new static equilibrium is placed at a higher level of the real exchange rate and a larger labor share of skilled workers

[15] When $\theta_S = \theta_L = \theta$, Ψ will be constant and equals to $\frac{\theta}{1-\theta}$. It implies that a change in income inequality plays no role in the economic performance

[16] From Equations 10b, the effect of income inequality on the real exchange rate is: $\frac{dP}{dI} = \frac{P}{\Psi} \frac{\theta_S - \theta_L}{(1-[\theta_L + (\theta_S - \theta_L)I])^2}$.

employed by the non-traded sector.
In the economy's response to a real exchange rate appreciation, there are two conflicting forces at work. The first force refers to labor reallocation due to a change in real wage between sectors (henceforth called the spending effect). While the second force refers to labor reallocation due to a change in income level between household's groups (henceforth called the inequality effect).
Regarding the spending effect, a real exchange rate appreciation leads to increasing the real wage of both skilled and unskilled workers in the non-traded sector with respect to that in the traded sector. Hence, it signals to both workers' groups to move from the traded to the non-traded sector. Consequently, the production sector of traded goods shrinks and the production sector of non-traded goods expands (i.e. the Dutch disease).
Regarding the inequality effect, I first look at the effect of a real exchange rate appreciation on the relative real wage. Figure 1 illustrates this effect through an Edgeworth box with skilled labor on the vertical axis and unskilled labor on the horizontal axis. Workers employed by the non-traded sector are measured as the distance from point O_N. Similarly, workers in the traded sector are measured from point O_T. The contract curve lies above the diagonal line if the non-traded sector is relatively intensive in skilled labor (see. Figure 1-a), while it lies below the diagonal line if the non-traded sector is relatively intensive in unskilled labor (see. Figure 1-b). The economy moves from point A to point B as the windfall income increases. Subscripts o and n refer to "old" and "new" equilibrium points, respectively. Although both types of workers shift away from the traded sector and into the non-traded sector, to what extent they shift depends on factor intensity between sectors.
Real exchange rate appreciation increases the relative factor price of skilled to unskilled workers



(a) $\alpha > \beta$ (b) $\alpha < \beta$

Figure 1: Factors movement in response to a windfall income

(i.e. $\frac{w_S}{w_L}$) if skilled workers are used intensively in the non-traded sector. The converse occurs if skilled workers are used intensively in the traded sector. These are clarified in Figure 1 by a comparison between the slope line of relative real wage in point (B) and point (A). Whereas for the former case, the slope line will be flatter (see. Figure 1-a), for the latter case it will be steeper (see. Figure 1-b). These are direct results of *Stolper-Samuelson theorem* (Stolper and Samuelson, 1941). The theorem represents that the relative price appreciation increases the relative demand of factor used intensively in the non-traded sector to another factor. This, in turn, increases the relative factor price. Let $I^{NR} \equiv \frac{w_S}{w_L+w_S}$ denote the Market-based (non-resource) income inequality , I summarize the above argument as the following proposition.

Proposition 1: *In a comparative static, for any* $\alpha, \beta > 0$,
$a)$ *if* $\alpha > \beta$, $\frac{dI^{NR}}{dR} > 0$
$b)$ *if* $\alpha < \beta$, $\frac{dI^{NR}}{dR} < 0$.

Proof: The results are derived from the graphical framework (Figure1) and given that $\frac{dP}{dR} > 0$.

Now, I am interested in how the distribution of resource benefits between workers' groups affects inequality. I plausibly assume that skilled workers have more political power than unskilled ones [17] and so skilled workers appropriate a larger share of resource benefit (e.g. subsidies). This assumption is not far from what is being observed in many natural resource economies. For example in most *Middle East* and low-income *sub − Saharan* countries, an elite group controls the government as well as the resource rent (Lam and Wantchekon, 2002). Further, It is more likely that a larger share of subsidies (e.g. on energy price) is received by the richest quintiles (Sdralevich et al., 2014).

When windfall income is distributed unevenly in favor of skilled workers, two cases may occur in income inequality through a windfall income boom. Whereas, for the first case in which skilled workers are used intensively in the non-traded sector income inequality will rise, for the second one in which skilled workers are used intensively in the traded sector income inequality change will be ambiguous. The results are derived from the response of the two main sources of income inequality to a resource boom. For the first case, both forces drive income inequality in the same direction. On the one hand, a larger share of windfall income is appropriated by skilled workers and on the other hand, their relative real wage increases. In conclusion, both sources of income, going to skilled workers, tends to rise income inequality. For the second case, these two forces drive income inequality in opposite directions. Windfall income is still distributed unevenly in favor of the skilled workers and so income inequality tends to go up. Nevertheless, a resource boom reduces the relative real wage of skilled workers, so that income inequality tends to go down. Consequently, the effect of a resource boom on income inequality change is ambiguous and depends on which one of these two driving forces is stronger.

To address total income inequality changes, let $y^{NR} \equiv w_L + w_S = PX_N + X_T$ and $y^R \equiv A_T R$

[17] Well-known evidence, implying that the political participation and so political voice is larger for the rich (skilled workers) than for the poor (unskilled workers), are documented by Petrocik and Shaw (1991), and Benabou (2000).

denote total non-resource income and total resource income, respectively. Therefore, the total income of the economy is $y = y^{NR} + y^{R}$. Now, the above discussion is summarized as follows,

Proposition 2: *In a comparative static, for any $\alpha, \beta > 0$ and $\pi > 0$,*
a) $\alpha > \beta$ and $\pi > I^{NR}$ are sufficient conditions to ensure that $\frac{dI}{dR} > 0$,
b) $\alpha < \beta$ and $\pi < I^{NR}$ are sufficient conditions to ensure that $\frac{dI}{dR} < 0$.

Proof: The results are found directly by computing $\frac{dI}{dR} = \frac{d\left(y^R/y\right)}{dR}\left[\pi - I^{NR}\right] + \left(1 - \frac{y^R}{y}\right)\frac{dI^{NR}}{dR}$ and applying Proposition 1.
The final point in the comparative static argument is how the gap in expenditure shares between skilled and unskilled workers can motivate the intensity of labor reallocation (i.e. the intensity of the Dutch disease). For this purpose, we need to know to what extent the gap in expenditure shares affects the real exchange rate appreciation. Since the larger real exchange rate appreciation, the more movement of labor from the traded to the non-traded sector.

Proposition 3: *In a comparative static, for any $\theta_S, \theta_L > 0$,*
a) assume $\alpha > \beta$ and $\pi > I^{NR}$ holds $\Rightarrow \frac{dI}{dR} > 0$,
1) a natural resource boom leads to a stronger real exchange rate appreciation if skilled workers, with respect to unskilled workers, spend more on non-traded goods than if spending shares are equal $\Leftrightarrow$ if $\theta_S > \theta_L$, $\left(\frac{dP}{dR}\right)_{\theta_S > \theta_L} > \left(\frac{dP}{dR}\right)_{\theta_S = \theta_L}$,
2) a natural resource boom leads to a weaker real exchange rate appreciation if skilled workers, with respect to unskilled workers, spend less on non-traded goods than if spending shares are equal $\Leftrightarrow$ if $\theta_S < \theta_L$, $\left(\frac{dP}{dR}\right)_{\theta_S < \theta_L} < \left(\frac{dP}{dR}\right)_{\theta_S = \theta_L}$,

b) assume $\alpha < \beta$ and $\pi < I^{NR}$ holds $\Rightarrow \frac{dI}{dR} < 0$,
1) a natural resource boom leads to a weaker real exchange rate appreciation if skilled workers, with respect to unskilled workers, spend more on non-traded goods than if spending shares are equal $\Leftrightarrow$ if $\theta_S > \theta_L$, $\left(\frac{dP}{dR}\right)_{\theta_S > \theta_L} < \left(\frac{dP}{dR}\right)_{\theta_S = \theta_L}$,
2) a natural resource boom leads to a stronger real exchange rate appreciation if skilled workers, with respect to unskilled workers, spend less on non-traded goods than if spending shares are equal $\Leftrightarrow$ if $\theta_S < \theta_L$, $\left(\frac{dP}{dR}\right)_{\theta_S < \theta_L} > \left(\frac{dP}{dR}\right)_{\theta_S = \theta_L}$.

Proof: These results are verified by computing the vertical shift of *NN*-curve in response to a natural resource boom (see. Appendix A).
The proposition represents when skilled workers with respect to unskilled workers allocate a larger expenditure share for traded goods (i.e. $\theta_S < \theta_L$), falling income inequality increases the relative consumption expenditure of non-traded to traded goods. This, in turn, increases the pressure on the real exchange rate to appreciate more. Then more labor forces must be shifted away from the traded and into the non-traded sector. Briefly, this argument says that falling

income inequality deepens the Dutch disease. In the same way, another case in which skilled workers with respect to unskilled workers allocate a larger expenditure share for non-traded goods (i.e. $\theta_S > \theta_L$) can be analyzed. Falling income inequality will moderate the Dutch disease through a weaker pressure on the appreciating real exchange rate, resulting from less demand for non-traded goods [18].

Before attention is turned to analyze the long-run mechanism, one seems to be useful to discuss the role of the institution's quality in the modification of the transition path. In this paper, I merely discuss one of the aspects of the role of institutions in the sustainable development of resource-dependence countries. Improving the quality of institutions can be associated with the expansion and improvement of public systems such as social security, health, education, and transportation that presumably lead to decreasing unskilled workers' expenditure shares on non-traded goods relative to skilled workers [19]. Hence, a reduction in the gap in the expenditure shares between workers' groups $(\theta_L - \theta_S)$ can be translated as an improvement in the quality of the institution. The real exchange rate appreciation is slightly moderated as the gap in expenditure shares decreases due to an improvement in the quality of the institution. This, in turn, leads to less intensity of falling market-based income inequality. With an unchanged distribution of the resource rent, the total income inequality level goes down less. Accordingly, less reduction in income inequality leads to less intensity of the Dutch disease. This argument concludes that the real exchange rate appreciation, falling income inequality, and the intensity of the Dutch disease are stronger in countries with low-quality institutions than those with high-quality institutions.

2.2 Long run analysis (dynamic model)

In this section, we endogenize factor productivity (TFP) to discuss the long-run steady-state effects of a resource boom. TFP is driven by the learning-by-doing (*LBD*) model. The earlier literature assumes that *LBD* is only generated by the traded sector (Krugman, 1987) or both sectors generate the learning but there is no spillover between them (Lucas Jr, 1988), whereas Sachs and Warner (1995) represent the case where only the traded sector generates the learning and there is a perfect spillover for the non-traded sector. Nevertheless, I modify Torvik (2001) framework which covers the earlier models [20]. I assume that productivity growth is generated by

[18] Figure 3 shows the impact of a gap in expenditure shares on the magnitude of real exchange rate appreciation (i.e. the vertical movement of *NN*-curve) for the case $\frac{dI}{dR} < 0$.

[19] Improving the distribution of subsidies that can financially be provided by the resource rent is an alternative case reflecting the role of the institution's quality. But, given that the market-based (non-resource) income inequality is the main driving force of the total income inequality (regarding empirical evidence presented), I ignore the analysis of this case.

[20] The departure of my framework from those papers is that I assume there are two workers' groups rather than one group and that the workers have different identical tastes.

skilled workers in both sectors with an imperfect spillover between sectors.

$$\frac{\dot{A}_T}{A_T} = v\,S_T + \delta_N\,u\,S_N \tag{12a}$$

$$\frac{\dot{A}_N}{A_N} = u\,S_N + \delta_T\,v\,S_T \tag{12b}$$

u and v denote the productivity growth rate of one unit skilled worker employed respectively by the non-traded and traded sectors. The constant $0 < \delta_T < 1$ measures a fraction of the learning generated by skilled workers in the traded sector and spilled over to the non-traded sector. Similarly, $0 < \delta_N < 1$ measures a fraction of the learning generated by skilled workers in the non-traded sector and spilled over to the traded sector. In these equations, Krugman (1987) represents the case where $u = \delta_T = 0$, while Lucas Jr (1988) assume that $\delta_T = \delta_N = 0$. In addition, these are simplified to Sachs and Warner (1995) model when $u = 0$ and $\delta_T = 1$.

For the purpose of the long-run analysis, I need to know how the static equilibrium of labor allocation reacts to a change in the relative productivity (i.e. $\phi \equiv \frac{A_T}{A_N}$). The response of labor allocation to an increase in the relative productivity ratio ϕ is ambiguous (Torvik, 2001). Since there are two conflicting forces at work. On the one hand, with an increase in the relative productivity, the labor requirement in the traded sector falls and that in the non-traded sector goes up (i.e. Labor requirement effect). With an unchanged composition of the basket for both groups, labor must shift away from the traded sector and into the non-traded sector. Hence, ignoring the effect on demand, an increase in the relative productivity for traded goods tends to increase employment in the non-traded sector. On the other hand, traded goods are cheaper to produce than before. So it is more likely that consumers substitute non-traded goods with traded goods (i.e. Substitution effect). Hence, shifting demand in favor of traded goods pushes both groups of labor employment to move from the non-traded sector to the traded sector. In conclusion, the two effects push S_N in opposite directions.

Now we can ask which one of these conflicting forces is dominant. On the one hand, for a given labor allocation, rising the relative productivity ratio states that the supply of the traded sector expands faster than the supply of the non-traded sector. Then the real exchange rate appreciates in order to back the good market in balance. Hence the NN-curve shifts up as ϕ increase (i.e. Labor requirement effect). The vertical shift in the NN-curve equals $\frac{P}{\phi\,\sigma_c}$ (i.e. $\frac{dP}{d\phi}$ in Equation 10b), where $\sigma_c \equiv \varepsilon_{T,P} - \varepsilon_{N,P}$ denotes the gap in the price elasticity of demand for traded and non-traded goods. On the other hand, for a given real exchange rate, the marginal productivity of labor is larger in the traded sector than in the non-traded sector as ϕ increases. Then the labor force in the non-traded sector falls to back the labor market in balance. Hence an increase in productivity level pushes the LL-curve to the left (i.e. Substitution effect). The vertical shift in the LL-curve equals $\frac{P}{\phi}$ (i.e. $\frac{dP}{d\phi}$ in Equation 10a). This argument represents that the vertical movement will be larger in NN-curve than in LL-curve if $\sigma_c < 1$. I summarize this argument as follows,

Summarized argument: $S_N = S_N(\phi, R)$, $\frac{dS_N}{d\phi} > 0 \ \ if \ \ \sigma_c < 1 \quad and \quad \frac{dS_N}{dR} > 0.$

Now I ask, is there a balanced growth path along which the productivity levels of both sectors grow at the same rate? The growth rate of the relative productivity ratio is given by:

$$\frac{\dot{\phi}}{\phi} = \frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} = (1-\delta_T)\,v - [(1-\delta_T)\,v + (1-\delta_N)\,u]\,S_N(\phi, R)\,. \tag{13}$$

To prove the existence of a balanced growth path, I investigate the stability properties of the dynamic model. The rate of change in relative productivity ratio is governed by:

$$\frac{d\left(\dot{\phi}/\phi\right)}{d\phi} = -\left[(1-\delta_T)\,v + (1-\delta_N)\,u\right]\frac{dS_N(\phi, R)}{d\phi}. \tag{14}$$

Equation 14 states that there is a balanced growth path iff the skilled labor employment in the non-traded sector increases as the relative productivity ratio jumps up (i.e. $\frac{dS_N}{d\phi} > 0$). The static response of an economy to a shift in the relative productivity ratio clarifies that the stability of the dynamic steady-state depends on the gap in the price elasticity of demand ($\sigma_c \equiv \varepsilon_{T,P} - \varepsilon_{N,P}$). Regrading the summarized argument of the vertical shift in curves, the dynamic stability of the system is satisfied if $\sigma_c \leq 1$. A little gap in the price elasticity of demand (σ_c) induces the demand side to shift to the extent which is sufficient to counteract the labor movement effects.
Now let me take into account that households have identical tastes (i.e. $\theta_S = \theta_L$). This results in $\sigma_c = \sigma$. When $\sigma > 1$, the economy ends up with a specialization in one of two goods. It converges to specialization in traded (non-traded) goods if the relative productivity is larger (smaller) at initial than at steady-state. Nevertheless, $\sigma \leq 1$ is the sufficient condition to guarantee stability. There exists a unique balanced growth path for the case $\sigma < 1$, while for the case $\sigma = 1$ implying that there is a Cobb-Douglas utility function there is a set of growth path. For the latter case, the vertical shifts of the curves, moving in opposite directions, are equal. It implies that a change in ϕ has no effect on S_N (i.e. $\frac{dS_N}{d\phi} = 0$) so that the labor share of skilled workers in the non-traded sector is constant over time. Consequently, one of the goods produces faster than the other [21] and the relative price of the other good increases to keep the budget constraint in balance (for more information see. Lucas Jr (1988); Torvik (2001)).
Among literature, Lucas Jr (1988) and Sachs and Warner (1995) have balanced productivity growth, while Krugman (1987) is an unbalanced growth model. The case where $\theta_S = \theta_L$ is not an interesting case for us. Since a change in income inequality plays no role in the mechanism of the model. Therefore, I argue below the case $\theta_S \neq \theta_L$.

Proposition 4: *if* $\theta_S < \frac{1}{1+\Psi} < \theta_L$ *holds,* $\sigma < 1$ *is a sufficient condition to satisfy the dy-*

[21] In a special case where $\frac{S_T}{S_N} = \frac{u(1-\delta_N)}{v(1-\delta_T)}$, both sectors expand at the same rate.

namic stability of system, where $\Psi \equiv \frac{P^{\sigma} C_N}{C_T}$ [22].

Proof: Regarding Equation 14, the stability of the dynamic system is satisfied if $\sigma_c < 1 \Rightarrow \frac{dS_N}{d\phi} > 0 \Rightarrow \frac{d(\dot{\phi}/\phi)}{d\phi} < 0$. Hence, if we find a local condition resulting in $\sigma_c < \sigma$, we can conclude that $\sigma < 1$ will be a sufficient condition to exist a balanced growth path. The computations prove that $\sigma_c < \sigma$ is satisfied if the expenditure share on non-traded goods is larger for unskilled workers than for skilled workers so that $\theta_S < \frac{1}{1+\Psi} < \theta_L$. For an analytic proof see Appendix B.

Before I turn to dynamic equilibrium, it is useful to take a closer look at the constraint of the stability condition, $\theta_S < \frac{1}{1+\Psi} < \theta_L$. The analytic computation reveals that the constraint can be rewritten as the following inequality relation: $I_N < I < I_T$, where $I_N \equiv \frac{C_S^N}{C_N}$ and $I_T \equiv \frac{C_S^T}{C_T}$ are, respectively, the consumption inequality of non-traded and traded goods (see. Appendix B). This implies when $\sigma < 1$, the dynamic system is stable iff the consumption distribution of non-traded goods is more evenly than the consumption distribution of total goods and the consumption distribution of traded goods is more unevenly than the consumption distribution of total goods. To make clear this argument, now let me discuss a special case in which skilled workers spend only on traded goods and unskilled workers spend only on non-traded goods (i.e. $\theta_L = 1$ and $\theta_S = 0$) [23]. The consumption inequality on traded goods equals one while the other is zero (i.e. $I_T = 1$ and $I_N = 0$). Therefore, the constraint is always satisfied (i.e. $I < I_T = 1$). As a conclusion when the expenditure share on non-traded goods is much larger for unskilled workers than for skilled workers ($\theta_S \ll \theta_L$), it is more likely that the constraint ($I_N < I < I_T$) is satisfied and so $\sigma < 1$ will be a sufficient condition to satisfy the dynamic stability of system.

When the local stability condition is satisfied, the model has a stable solution for the relative productivity ratio, denoted by ϕ^*. At such a dynamic equilibrium, the skilled and unskilled labor forces are also at their steady-state values [24].

$$S_N^* = \frac{(1-\delta_T)\,v}{[(1-\delta_T)\,v + (1-\delta_N)\,u]} \qquad L_N^* = \frac{1}{1+\left(\frac{1-\beta}{1-\alpha}\right)\left(\frac{\alpha}{\beta}\right)\left[\frac{(1-\delta_N)u}{(1-\delta_T)v}\right]}. \tag{15}$$

Equation 15 clearly implies that the steady-state labor allocation of skilled and unskilled workers are independent of the windfall income boom. It, in turn, results in a constant relative real wage

[22] Note that there exists the other stability condition. For the case $\theta_S > \frac{1}{1+\Psi} > \theta_L$, the sufficient condition to guarantee the stability is $\sigma < 1 - (\gamma_T - \gamma_N)\left(\varepsilon_{N,P}^L - \varepsilon_{N,P}^S\right)$, where $\gamma_J \equiv \frac{C_J^L}{C_J}$; $J = T, N$ is the share of good J allocated to unskilled workers and $\varepsilon_{N,P}^i \equiv \frac{P}{C_N^i}\frac{\partial C_N^i}{\partial P} = -\left(\frac{\sigma(1-\theta_i)+\theta_i P^{1-\sigma}}{1-\theta_i+\theta_i P^{1-\sigma}}\right)$; $i = L, S$ is the price elasticity of non-traded goods in terms of i workers.

[23] Regarding Equation 5, the aggregate price index or the unite expenditure is equal to P for unskilled workers while it is equal to one for skilled workers.

[24] S_N^* is determined from steady-state equilibrium (i.e. $\frac{\dot{\phi}}{\phi} = 0$), while L_N^* is found by replacing S_N^* at the factor market equilibrium (i.e. $\frac{w_L}{w_S} = \frac{1-\alpha}{\alpha}\frac{S_N^*}{L_N^*} = \frac{1-\beta}{\beta}\frac{1-S_N^*}{1-L_N^*}$).

at the steady-state. The steady-state value of income inequality is also given by:

$$I^* = \frac{\beta + (\eta^*)^{1-\beta}\,\pi R}{\beta + (1-\beta)\,\eta^* + (\eta^*)^{1-\beta} R} \tag{16}$$

where $\eta^* \equiv \frac{S_T^*}{L_T^*}$. Equation 16 demonstrates that the steady-state income inequality is only driven by a natural resource boom. Therefore, factor intensity has no effect in the analysis of the balanced growth path and the distribution of windfall income between workers' groups is the only determinant of inequality change. This clarifies that a change in income inequality is a strong transmission channel in the long-run growth path. Let I^{NR^*} denotes the Market-based (non-resource) income inequality at the steady-state.

Proposition 5: *At steady-state, for any $\pi > 0$,*
a) If resource income inequality is larger than steady-state non-resource income inequality, a natural resource boom increases total income inequality $\Leftrightarrow$ if $\pi > I^{NR^*}$, $\frac{dI^*}{dR} > 0$,
b) If resource income inequality is smaller than steady-state non-resource income inequality, a natural resource boom decreases total income inequality $\Leftrightarrow$ if $\pi < I^{NR^*}$, $\frac{dI^*}{dR} < 0$.

Proof: The results are derived through $\frac{dI^*}{dR} = \frac{d\left(y^R/y\right)}{dR}\left[\pi - I^{NR^*}\right] + \left(1 - \frac{y^R}{y}\right)\frac{dI^{NR^*}}{dR}$ and given that $\frac{dI^{NR^*}}{dR} = 0$ at steady-state.

2.3 Dutch disease dynamics

The aim of this section is to study the dynamic effect of a natural resource boom on the growth rate of the relative productivity ratio. From Equation 13, the derivative of growth rate with respect to R gives:

$$\frac{d\left(\dot{\phi}/\phi\right)}{dR} = -\left[(1-\delta_T)\,v + (1-\delta_N)\,u\right]\frac{dS_N(\phi,R)}{dR}. \tag{17}$$

Given that $\frac{dS_N}{dR} > 0$, the economy throws out the steady-state equilibrium as the windfall income increases. A natural resource boom through the real exchange rate appreciation shifts the skilled labor force away from the traded sector and into the non-traded sector. Since S_N goes up relative to its steady-state value the relative productivity growth rate declines along the transition path (i.e. $\frac{\dot{\phi}}{\phi} < 0$). As a consequence, the relative productivity ratio declines to re-establish the steady-state equilibrium. This fall in ϕ induces a countervailing movement of labor from the non-traded to the traded sector that gradually brings back the allocation of labor to its long-run equilibrium level.

Figure 2 shows the adjustment balanced growth path. The locus of relative productivity ratio is a downward-sloping line in order to satisfy the stability condition (see. equation 14) [25]. When the windfall income increases, the economy jumps vertically down from the solid line to the dotted line and from there it moves towards the new dynamic equilibrium ϕ^{**}. This concludes whereas in the short-term, a labor reallocation between sectors brings about the Dutch disease, in the long-term a change in the relative productivity level consolidates the Dutch disease.

A useful undertaking is now to investigate the gap in expenditure shares between groups'

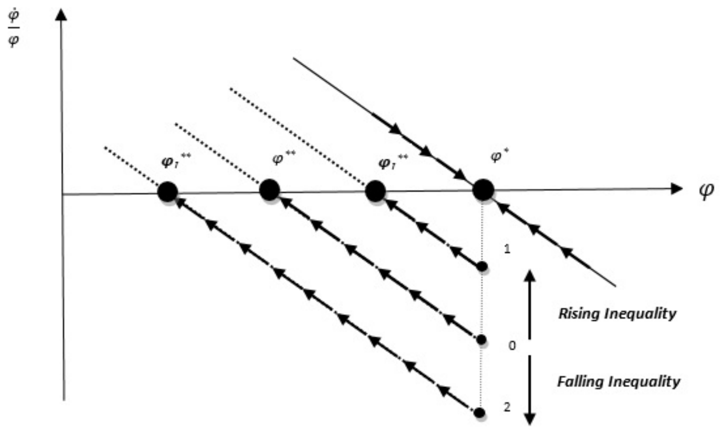


Figure 2: Adjustment balanced growth path for case $\theta_S < \theta_L$

workers on the intensity of the Dutch disease along the transition path. For a case in which both skilled and unskilled workers have identical tastes, the economy first jumps down from steady-state ϕ^* to point 0 and then it goes towards the new steady-state, denoted by ϕ^{**}. In this case, income inequality has no effect on the adjustment path. While for a case in which consumption baskets of skilled and unskilled workers are different, income inequality plays a key role in the adjustment path. Regarding Proposition 4, I only address the case $\theta_S < \theta_L$. If skilled workers allocate a larger expenditure share on traded goods ($\theta_S < \theta_L$), rising income inequality reduces the intensity of the real exchange rate appreciates along the transition path. This, in turn, makes the reallocation of skilled labor weaker than it would otherwise be. Thus the impact is smaller. The economy moves down to point 1 upon impact and from there, it moves towards the larger steady-state ϕ_I^{**}, compared with the former case reaching ϕ^{**}. The opposite occurs if income inequality falls. Briefly, when $\theta_S < \theta_L$, rising income inequality decreases the intensity of the Dutch disease and falling income inequality increases the intensity of the Dutch disease.

Proposition 6: *At steady-state, assume unskilled workers spend more on non-traded goods than skilled workers do* $\Leftrightarrow \theta_S < \theta_L$,
a) rising income inequality moderates the intensity of the Dutch disease $\Leftrightarrow$ *if* $\frac{dI^*}{dR} > 0$, $\phi_I^{**} > \phi^{**}$,

[25] Assume initially $\phi > \phi^*$. Since the labor movement effect is stronger than the substitution effect (i.e. $\sigma < 1$), the employment in the non-traded sector is larger than it would be at steady-state. It implies that productivity growth is stronger in the non-traded sector than in the traded sector and so the relative productivity growth falls over time until it reaches its steady-state value.

$b)$ *falling income inequality exacerbates the intensity of the Dutch disease* $\Leftrightarrow$ *if* $\frac{dI^*}{dR} < 0$, $\phi_I^{**} < \phi^{**}$.

Proof: Regarding Equation 10a at steady-state, $\frac{dP^*}{d\phi^*} > 0$. It implies when the economy throws out the initial equilibrium, both P^* and ϕ^* shift in the same direction. Hence, the proposition can be verified by taking a derivative of the steady-state real exchange rate, rather than the steady-state productivity ratio, with respect to R, (i.e. $\frac{dP*}{dR}$) (see. Appendix C).
Figure 3 shows how the labor share in the non-traded sector, S_N, and the real exchange rate, P, react in turn to an exogenous shock. *LL*-curve (Equation 10a) and *NN*-curve (Equation 10b) are drawn as upward and downward sloping curves, respectively. A natural resource boom raises the employment share in the non-traded sector, S_N. This, in turn, reduces the relative productivity ratio. Therefore, both the *LL*-curve and the *NN*-curve shift down along the transition path. Since the productivity level grows faster in the non-traded sector than in the traded sector [26], this movement will continue as long as the labor share in the non-traded sector backs in its initial value. Regarding Section 2.2, the *NN*-curve shifts faster than the *LL*-curve, if $\sigma < 1$. Hence, both S_N and P fall along the transition process.
Now, I am interested to study the effect of the gap in expenditure shares between workers' groups on the steady-state real exchange rate. I address the case in which skilled workers spend a larger expenditure share on traded goods (i.e. $\theta_S < \theta_L$). This case is in consistent with the stability condition (Proposition 4) and in keeping with the empirical results, presented in Section 3.

Proposition 7: *At steady-state,*
$a)$ *When resource income inequality is larger than steady-state non-resource income inequality, a natural resource boom leads to a weaker real exchange rate appreciation for case* $\theta_S < \theta_L$ *than for case* $\theta_S = \theta_L \Leftrightarrow$ *when* $\pi > I^{NR^*}$, $\left(\frac{dP^*}{dR}\right)_{\theta_S<\theta_L} > \left(\frac{dP^*}{dR}\right)_{\theta_S=\theta_L}$
$b)$ *When resource income inequality is smaller than steady-state non-resource income inequality, a natural resource boom leads to a stronger real exchange rate appreciation for case* $\theta_S < \theta_L$ *than for case* $\theta_S = \theta_L \Leftrightarrow$ *when* $\pi < I^{NR^*}$, $\left(\frac{dP^*}{dR}\right)_{\theta_S<\theta_L} < \left(\frac{dP^*}{dR}\right)_{\theta_S=\theta_L}$.

Proof: It is contained in Appendix C.
Figure 3 clarifies that falling income inequality increases the relative consumption expenditure of non-traded to traded goods. This, in turn, tends to shift a larger labor share away from the traded sector and into the non-traded sector. Hence, S_N is larger for the case $\theta_S < \theta_L$ than $\theta_S = \theta_L$ as a natural resource boom is accompanied by falling income inequality. This concludes that along the transition path the *LL*-curve and the *NN*-curve shift down more for the case $\theta_S < \theta_L$ than it would otherwise be. The new dynamic equilibrium is at a point such as E_2 for the case

[26] An alternative is that the productivity level shrinks faster in the traded sector than in the non-traded sector (see. section 2.4). This case is consistent with the empirical evidence (see. section 3.4 and 3.5).

$\theta_S = \theta_L$, while it rests in a lower point such as E_2' for the case $\theta_S < \theta_L$. Briefly, the steady-state real exchange rate depreciates more when $\theta_S < \theta_L$.

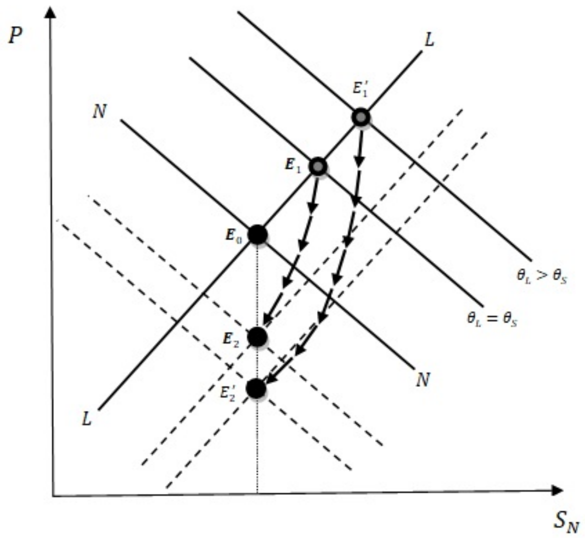


Figure 3: The development of comparative static when $\alpha < \beta$ and $\pi < I^{NR} \Rightarrow \frac{dI}{dR} < 0$

2.4 Absolute Productivity Growth

So far I have discussed how a natural resource boom affects relative productivity growth. But this is not the only issue of interest. It is worth conducting an investigation into the response of absolute productivity growth (i.e. $\dot{A}_T/A_T$ and $\dot{A}_N/A_N$), rather than the relative productivity growth (i.e. $\dot{\phi}/\phi$), to a factor reallocation.
Equations 12 verifies that the response of the absolute productivity level in each sector depends on the size of the direct effect of learning in each sector and the indirect effect of learning spilled over from another sector. Figure 4, the same as Torvik (2001), displays this argument. In response to a natural resource boom, the labor share of skilled workers in the non-traded sector increases. Therefore, if the direct effect of learning generated by the traded sector is stronger than the spillover effect (i.e. $v > \delta_N u$), the productivity growth rate in the traded sector decelerates as S_N increases. Further, equation 12b verifies that an increase in S_N accelerates the productivity growth in the non-traded sector when the direct effect is dominant (i.e. $u > \delta_T v$) and it decelerates the productivity growth when the spillover effect is dominant (i.e. $u < \delta_T v$).

Proposition 8: *For any* $u, v > 0$ *and* $0 < \delta_T, \delta_N < 1$,
a) The natural resource boom decelerates the productivity growth in the traded sector, If the direct effect of learning is stronger than the spillover effect $\Leftrightarrow$ *if* $v > \delta_N u$, $\frac{d(\dot{A}_T/A_T)}{dR} < 0$,
b) The natural resource boom decelerates the productivity growth in the non-traded sector, If the direct effect of learning is weaker than the spillover effect $\Leftrightarrow$ *if* $u < \delta_T v$, $\frac{d(\dot{A}_N/A_N)}{dR} < 0$.

Proof: The results are found through the derivative of the absolute productivity growth with respect to R, $\frac{d(\dot{A}_T/A_T)}{dR} = -(v-\delta_N u)\frac{dS_N}{dR}$ & $\frac{d(\dot{A}_N/A_N)}{dR} = -(\delta_T v - u)\frac{dS_N}{dR}$, and given that $\frac{dS_N}{dR} > 0$.

The question arising now is the effect of a change in income inequality on the productivity

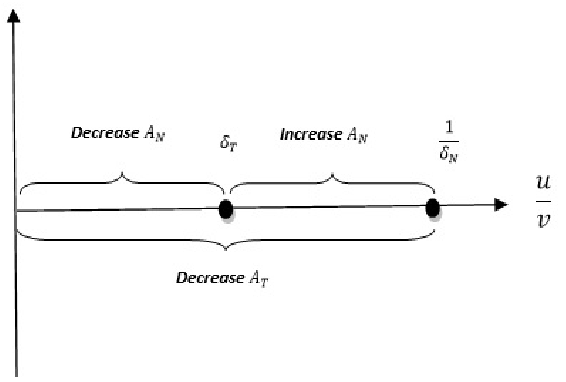


Figure 4: Absolute productivity growth change in response to an increase in S_N.

growth in each sector. I concentrate on the case $\theta_S < \theta_L$ which is consistent with the stability condition and the empirical evidence. On the one hand, a natural resource boom increases the labor share of skilled workers in the non-traded sector (S_N). In response, the relative productivity, ϕ, falls to shift down the *LL*-curve and *NN*-curve along the transition path. The reduction in the relative productivity ratio can be paraphrased as a faster expansion of the productivity level in the non-traded sector than in the traded sector. But there is an alternative. It states that the productivity level shrinks faster in the traded sector than in the non-traded sector. Which one of these two states occurs depends on the gap between the direct and spillover effect for each productivity level. Regarding Proposition 8, if the direct effect of learning generated by the non-traded sector is weaker than the spillover effect (i.e. $\frac{u}{v} < \delta_T$) [27], the productivity level should shrink faster in the traded sector than in the non-traded sector to reduce the relative productivity level (Figure 4).

On the other hand, a natural resource boom can lead to falling income inequality if resource inequality is smaller than non-resource income inequality (i.e. $\pi < I^{NR}$) [28] and skilled workers are used intensively by the traded sector than by the non-traded sector (i.e. $\alpha < \beta$). (Proposition 3). Therefore, a natural resource boom through falling income inequality makes the real exchange rate appreciation larger for the case $\theta_S < \theta_L$ than for the case $\theta_S = \theta_L$ (see. Figure 3). Consequently, S_N is larger for the case $\theta_S < \theta_L$ than the case $\theta_S = \theta_L$ as a natural resource boom is accompanied by falling income inequality. This concludes as windfall income goes up, the productivity growth deceleration in both sectors is more intensive for the case $\theta_S < \theta_L$ than for the case $\theta_S = \theta_L$. The above argument is summarized as the following proposition

[27] If this assumption holds, both productivity growth decelerates as S_N goes up (see. Figure 4).
[28] This is equal to $\frac{\pi}{1-\pi} < \frac{w_S}{w_L}$, the relative resource rent benefit of skilled workers is smaller than the relative real wage.

Proposition 9: *Assume the direct effect of learning generated by the non-traded sector is weaker than the spillover effect (i.e.* $\frac{u}{v} < \delta_T$*), for the case* $\alpha < \beta$ *and* $\pi < I^{NR} \Rightarrow \frac{dI}{dR} < 0$*, the productivity growth deceleration in both sectors is more intensive when income inequality reduces than when income inequality plays no role* $\Leftrightarrow \left(\frac{d\left(\dot{A}_J/A_J\right)}{dR}\right)_{\theta_S<\theta_L} > \left(\frac{d\left(\dot{A}_J/A_J\right)}{dR}\right)_{\theta_L=\theta_S}$*, where* $J = T, N$*.*

Proof: We know the derivatives of the productivity growth with respect to R (Proof of Proposition 8) and given that $\left(\frac{dS_N}{dR}\right)_{\theta_S<\theta_L} > \left(\frac{dS_N}{dR}\right)_{\theta_L=\theta_S}$ (Proposition 3 and Figure 3).

Before attention is turned to empirical evidence, It seems to be useful to discuss the steady-state growth rate. By inserting the steady-state labor share of skilled workers in the non-traded sector (S_N^* at Equations 15) in one of Equations 12, the steady-state growth rate is given :

$$g = \frac{uv\left(1-\delta_T\delta_N\right)}{\left(1-\delta_T\right)v + \left(1-\delta_N\right)u}. \tag{18}$$

This Equation reveals that at steady-state the productivity level in each sector and total non-resource income expand equally irrespective of the natural resource (i.e.$\frac{\dot{A}_T}{A_T} = \frac{\dot{A}_N}{A_N} = \frac{\dot{X}}{X} = g$)[29]. Also, total income grows at the same rate if R is constant over time (i.e. $\frac{\dot{Y}}{Y} = g$).

3 Empirical approach

In this section, I present an empirical approach to find some reliable evidence in order to support the proposed theory. The main contribution is to estimate the effect of a natural resource boom on inequality, and the feedback effect of inequality on the intensity of the Dutch disease. Before moving to the description of the econometric methodology and the variables, let me briefly conclude the structural relationship of the model to bridge the gap between theory and empiric. Figure 5 shows the conceptual mechanism of the proposed model.

*Arrow*1 : The real exchange rate appreciates as the windfall income increases.

*Arrow*2 : The real exchange rate appreciation induces income inequality according to a sectoral factor intensity, so that the market-based income inequality rises (falls) if the non-traded sector is relatively intensive in skilled (unskilled) workers (Proposition 1).

*Arrow*3 : Distribution of the resource income benefits (e.g. subsidies) in participation with distribution of the market-based income changes the total income inequality (Proposition 2).

*Arrow*4 : In response to an increase in the real exchange rate, the relative productivity ratio reduces.

*Arrow*5 : On the assumption that unskilled workers with respect to skilled workers spend more

[29] The result proves as follows: $\frac{\dot{X}}{X} = \frac{X_T}{X}\frac{\dot{X}_T}{X_T} + \frac{PX_N}{X}\left[\frac{\dot{P}}{P} + \frac{\dot{X}_N}{X_N}\right]$. From Equations 12 we have $\frac{\dot{X}_T}{X_T} = \frac{\dot{A}_T}{A_T} = g$ and $\frac{\dot{X}_N}{X_N} = \frac{\dot{A}_N}{A_N} = g$. It also yields $\frac{\dot{P}}{P} = \frac{\dot{\phi}}{\phi} = \frac{\dot{A}_T}{A_T} - \frac{\dot{A}_N}{A_N} = 0$. Therefore, $\frac{\dot{X}}{X} = g$

on the non-traded goods, falling (rising) income inequality exacerbates (moderates) the intensity of the Dutch disease (Proposition 6).

In what follows, I estimate the structural relationship between key variables of interest in

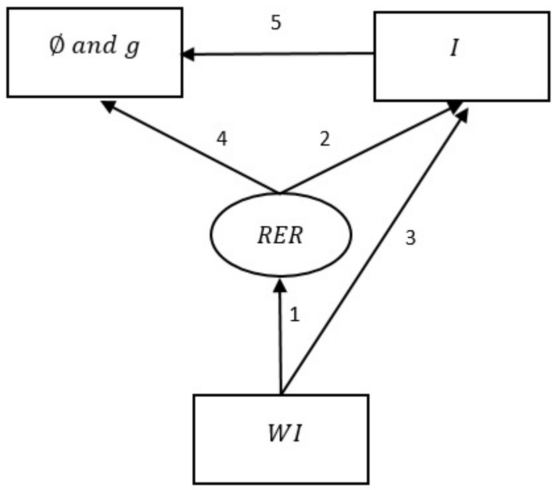


Figure 5: the conceptual mechanism of the model. WI is the windfall income (a natural resource boom). The real exchange rate and income inequality are respectively denoted by RER and I. In addition, g and ϕ denote the economic growth and the relative productivity ratio.

following sub-sections to clarify the conceptual mechanism of the model:
1) the effect of the windfall income on the real exchange rate along the transition path,
2) the response of income inequality to a change in the windfall income,
3) the combination of a natural resource boom and income inequality changes on the intensity of the Dutch disease, and
4) the impact of a natural resource boom on the sectoral economic growth along the transition path.

3.1 Data and Empirical methodology

Empirical attempts to analyze how windfall income affects economic performance have predominantly been performed using cross-sectional regression (e.g Sachs and Warner (1995)). This procedure is unable to capture the effect of a change in our variables of interest (e.g. income inequality) on the economic performance of a natural resource-dependent economy. For this reason, I employ a panel data approach to elaborate on this problem.
The unbalanced panel sample comprises 79 countries and covers non-overlapping 5-year periods of data observed from 1975 to 2014 [30]. The list of countries included in the sample database presents in Appendix D. Table 1 reports summary statistics. The average for 5-year periods of per capita GDP (Constant 2010 US $dollars$), and per capita manufacturing and service value-added (Constant 2010 US $dollars$) are sourced from the World Bank's World Development Indicator

[30] 5-year periods are: 1975-79, 1980-84, 1985-89, 1990-94, 1995-99, 2000-04, 2005-09, 2010-14. By taking into account five-year averages, I prevent the effects of short term fluctuation.

Database (WDI). Income inequality, measured by the Gini index, is gathered from the World Bank World Development Indicators (World Bank, 2013) and the University of Texas Inequality project dataset [31]. Due to the non-existence of price index for traded and non-traded goods, I use the real effective exchange rate, estimated by *Bruegel* (Darvas, 2012) as a proxy for the relative price [32].

The windfall income is measured by total natural resources rents (% of GDP) collected from

Table 1: Summary Statistics

	Obs.	Mean	Std Dev.	Min	Max
Ln (per capita GDP)	479	8.69	1.37	5.12	11.41
Ln (per capita GDP of Manufacturing)	400	6.72	1.42	2.08	9.24
Ln (per capita GDP of Service)	409	8.02	1.51	3.92	10.76
Gini index	479	42.67	6.30	26.97	56.38
Real effective exchange rate (divided by 100)	474	1.13	0.77	0.33	11.95
Windfall income proxy	479	5.06	6.96	0.01	47.33
Commodity Price index	181	1.28	0.65	1.00	6.07
Population growth	479	1.32	1.05	-1.50	4.60
Investment ratio	479	23.90	6.16	4.84	54.42
Human Capital index	479	2.38	0.65	1.08	3.72
Openness index	479	63.28	31.09	9.50	205.54
Terms of trade	479	1.01	0.25	0.21	3.04
Foreign direct investment	471	2.59	3.76	-3.13	40.41
Government spending	469	15.47	5.18	4.14	46.89
Excess Money growth	389	0.09	0.26	-1.43	1.47
Rule of Law indicator	479	0.21	0.99	-1.74	1.96
Control of Corruption indicator	479	0.28	1.09	-1.37	2.51

WDI. It is further proxied by the non-agriculture commodity export price index[33], constructed using a similar methodology to Deaton et al. (1995), Dehn (2000), Collier and Goderis (2008). Whereas the former index captures the impact of both price and quantity variations on the windfall income, the latter one represents only price variation. To construct the composite commodity export price index, I first collect data on world prices of 16 non-agriculture commodities [34] as well as commodity export and import values for available countries data over the period

[31] The Estimated Household Income Inequality dataset (EHII) is derived from the econometric relationship between UTIP-UNIDO and the World Bank's Deininger and Squire (1996) dataset (i.e. Galbraith and Kum (2005)).

[32] The real effective exchange rate $(REER)$ is calculated as $REER = \frac{NEER.CPI^{domestic}}{CPI^{foreign}}$, where $NEER$ is the nominal effective exchange rate of the country under study.

[33] Adopting the methodology defined in WDI, the concept of reserves does not apply to many commodities including agricultural products. Hence, I only construct the non-agricultural commodity price for as many countries as data availability allowed to make proxies be consistent together.

[34] The non-agricultural commodities are aluminium, cobalt, lead, Oil crude, tin, coal, natural gas, phosphates, platinum, zinc, copper, iron ore, nickel, silver, uranium, wood. Points: 1) a normalized average prices (US dollar) of hard and soft swan-woods is given as the price of wood (2010=100), 2) a normalized average prices (US dollar) of coal for Australia and South Africa bases is given as the price of coal (2010=100), 3) a normalized average prices (US dollar) of natural gas for USA and Europe bases is given as the price of gas (2010=100), and *APSP* crude oil price that is weighted average of three crude oil spot prices(west Texas intermediate, dated Brent and Dubai Fateh) is considered as oil crude price.

1990-2014. Further, commodity export and import data for each country are collected from the $UNCTSD$ (United Nations Commodity Trade Statistics Database) database [35], while data for commodity price indices are sourced from the $IMF-IFS$ International Financial Statistics database. Then a country-commodity specific weight for each country in 1990 is constructed by dividing the individual 1990 net export values for each commodity to the total net export value of all commodities in 1990 [36]. The weight is held fixed over periods to prevent the index from arising possible endogeneity problems due to supply responses to world prices. More specifically, the geometrically-weighted index of commodity export prices for country i in year t (PC_{it}) is constructed as follows:

$$PC_{it} = \prod_{j=1}^{J} PC_{jt}^{w_{j90}}, \tag{19}$$

where PC_{jt} represents the international market prices for commodity j in year t and w_{j90} is country-commodity specific weight in 1990. Finally, the log of the geometrically weighted index is weighted by the 1990 share of net commodity exports in a country's GDP (see.Collier and Goderis (2008)) to allow the effect of commodity export prices to be larger for countries with higher commodity exports.

Following the recent empirical studies, I include a number of control variables in the regression equations. Commonly used control variables are Population growth [37], Investment ratio [38], Human Capital Index [39], Openness ratio [40], Terms of trade [41], Foreign direct investment [42] and Government spending [43]. Additionally, the difference between the growth rate of the Broad money $(M2)$ and Economic growth is defined as Excess Money growth [44]. The average data of Rule of Law and Control of Corruption indicators [45] are utilized to construct a normalized index,

[35] It reports dollar values of exports and imports according to the $SITC1$ system.

[36] For countries with missing 1990 data for commodity export and import values, I use the value available in the year closest to 1990.

[37] The average value for the 5-year period is sourced from WDI.

[38] Gross fixed capital formation, in the percentage of GDP, is used to proxy the investment ratio. Observed values as averages for the 5-year period are derived from WDI.

[39] Our measure of human capital is an index constructed by $Penn\ World\ Table$. This index is based on Barro and Lee (2013) database for the average years of schooling and an estimated rate of return for primary, secondary, and tertiary education, introduced by Caselli (2005). Following the recent work of Barro (2008), I observe the human capital index at the start of each period.

[40] I collect data of trade to GDP ratio, a proxy for openness level, from WDI database. The ratio is observed as averages for 5-year period.

[41] The net barter terms of trade index (2000=100) are calculated as the percentage ratio of the export unit value indexes to the import unit value indexes. The average value for the 5-year period is derived from WDI and $OECD$ database.

[42] It is defined as the net inflows of investment divided by GDP. The average value for the 5-year periods is collected from WDI.

[43] It refers to general government final consumption expenditure divided by GDP. The data is an average value for the 5-year period, sourced from WDI.

[44] The average growth rate of $M2$ for each period is calculated as $GM_t = Ln\left(\frac{M2_t}{M2_{t-1}}\right)$. The data source is WDI.

[45] Both indicators proposed by $World\ Bank's\ Governance\ Indicators\ Project$ are in the range of approximately -2.5 (weak) to $+2.5$ (strong). The data are available form 1996 to 2014. For preceding periods, I assume that the indicators are equal to the earliest value.

representing the quality of institution (Institution index) across countries.
In order to verify the theoretical findings, dynamic panel data models are applied. The general form of a dynamic model is specified as follow:

$$y_{i,t} = \alpha + \delta y_{i,t-1} + X_{i,t}^{'}\beta + \mu_i + \varepsilon_{i,t}. \tag{20}$$

Where the subscripts $i = 1,...N$ and $t = 1,...T$ index the countries and periods in the panel, respectively. $y_{i,t}$ and $X_{i,t}^{'}$ denote respectively the dependent variable and a vector of independent variables. α is a constant term, μ_i denotes the unobserved country-fixed effect and $\varepsilon_{i,t}$ is the error term.
Several econometric problems may arise in the estimation of Equations 20. The country-fixed effects may be correlated with the explanatory variables. Also, the inclusion of the lagged dependent variable among regressors gives rise to autocorrelation. The last problem arises when explanatory variables are predetermined or endogenous [46] instead of strictly exogenous. In this paper, I apply first-differenced GMM estimator developed by Arellano and Bond (1991) and the system GMM estimator proposed by Blundell and Bond (1998) to cope with these problems. In the former method, the lagged level variables instrument the explanatory variables, while in the latter method, both lagged level and lagged differences are used.
The levels of the dependent variables and the endogenous variables are instrumented by lagged two or more periods, while those for the levels of the pre-determined variables are instrumented by lagged one or more periods. Further, the valid instruments for the levels of exogenous variables are simply current or lagged periods. The quality of the estimated dynamic models must be verified to make sure that the results are valid. The utilization of GMM estimator requires two conditions: the serial autocorrelation of errors and the proliferation of instruments which causes overidentification in the estimation model. First states that the error term must be serially uncorrelated to verify the lagged variables as valid instruments. The condition can be avoided by using the *Arellano-Bond* test. The null hypothesis of the test is that autocorrelation doesn't exist. The joint validity of full instruments is also checked by using the *Hansen* test of over-identifying restrictions [47]. The null hypothesis of the test states that the instruments, as a group, are uncorrelated with the error term. Moreover, the rule of thumb is to keep the number of instruments less than the number of country groups. Therefore, I only use certain lags, instead of all available lags, as an instrument to avoid an overidentification (Roodman, 2009*b*). Finally, I use a two-step system GMM (first-differenced GMM) with Windmeijer (2005) robust correction procedure [48].

[46] A predetermined variable is correlated with past observation-specific disturbance while an endogenous variable is correlated with past and current observation-specific disturbance.
[47] *Hansen* test is adequate when the estimation considers a *heteroscedastic* weight matrix.
[48] All estimations and statistical tests described in continue are carried out using *xtabond2* command proposed by Roodman (2009*a*).

3.2 Impact of the windfall income on the real exchange rate

The aim of this section is to evaluate the effect of a natural resource boom on the real exchange rate. Following the proposed theory and the recent study (Torvik, 2001), the real exchange rate appreciates along the transition path as the windfall income increases. Further, regarding the assumption that the unskilled workers with respect to the skilled workers allocate a larger expenditure share on non-traded goods, falling income inequality tends the real exchange rate to appreciate more (see. Figure 3).

The dependent variable of the dynamic panel data model is the real effective exchange rate and the explanatory variables of interest are the windfall income proxy and *Gini* index. I also include a number of control variables, namely the level of *GDP* per capita, Excess Money growth, Government spending, Terms of trade, Openness index and Foreign direct investment. Both dependent and explanatory variables are log-transformed variables.

The empirical results for regression models are reported in Table 2. Columns (1) and (2) represent, respectively, *OLS* and System GMM estimators for the baseline specification model. Although the *OLS* estimation results may not be informative, the results may still be interesting as a benchmark. The estimated positive and significant coefficient for the correlation between the windfall income (lagged) and the real effective exchange rate confirms the theory's prediction that a natural resource boom appreciates the real exchange rate.

I include the Gini index in the baseline model to reveal the impact of income inequality on the real exchange rate. The results are presented in column (3). The estimated coefficient is significant at 1% and enters with a negative sign. It states that falling income inequality is associated with rising the real exchange rate (the relative factor price). This supports the theory's prediction if one is assumed that the rich (skilled workers) relative to the poor (unskilled workers) spend a larger share of their income on traded goods (i.e. Proposition 3). The finding contrasts with Min et al. (2015), where under non-homothetic preferences between different income groups more inequality causes an increased demand for non-traded goods. Since the high-income group has a higher elasticity of demand for non-traded goods than does the low-income group (i.e. Engel's law). Accordingly, the relative price of non-traded goods increases and so the real exchange rate applications.

The reported small estimated coefficient on the windfall income proxy in column (3) seems to imply a heterogeneity across countries. To discuss this issue, I include an interaction term between income level and the windfall income proxy in the baseline model. Column (4) reports the results. It demonstrates that in responding to an increase in the windfall income proxy, the real exchange rate appreciates more for developing countries than for developed countries. In other words, the real exchange rate appreciation attenuates as GDP per capita increases [49].

To test the robustness of the results, I restrict the sample to developing countries, introduced by the International Monetary Fund's World Economic Outlook Database. The results reported in column (5) show that the coefficient on the windfall income remains close to the baseline's

[49] The real exchange rate will never depreciate on the dataset. Since the largest per capita GDP (Ln) observation is equal to 11.41 and so $\frac{dP}{dR} \approx 0$.

Table 2: Estimation results for the real effective exchange rate

Explanatory variable	(1) OLS	(2) Sys.GMM	(3) Sys.GMM	(4) Sys.GMM	(5) Sys.GMM	(6) Sys.GMM	(7) Sys.GMM	(8) Sys.GMM	(9) Sys.GMM	(10) Sys.GMM
	Full Sample	Full Sample	Full Sample	Full Sample	Developing	Poor Instit	Non Europe	IMF	Recession	Commodity
Real effective exchange rate (lagged)	0.574***	0.483***	0.592***	0.519***	0.584***	0.537***	0.594***	0.494***	0.543***	0.499***
	(0.0338)	(0.135)	(0.0754)	(0.0727)	(0.0614)	(0.168)	(0.0761)	(0.128)	(0.123)	(0.158)
Windfall income (lagged)	0.0227**	0.0261***	0.0424***	0.173**	0.0407***	0.0268**	0.0331***	0.0337*	0.0609***	
	(0.00942)	(0.00927)	(0.0150)	(0.0777)	(0.0134)	(0.0130)	(0.0109)	(0.0192)	(0.0223)	
Gini index			-0.950***	-1.017***	-0.888**		-0.839**	-1.288*	-1.080**	
			(0.360)	(0.374)	(0.438)		(0.375)	(0.672)	(0.460)	
Windfall income *GDP per capita				-0.0152*						
				(0.00822)						
Commodity Price index (lagged)										0.108**
										(0.0537)
GDP per capita (lagged)	0.00407	-0.00926	-0.0904**	-0.0542	-0.0594	0.000497	-0.0635	-0.0680	-0.0895*	-0.0242
	(0.0141)	(0.0166)	(0.0414)	(0.0377)	(0.0502)	(0.0231)	(0.0508)	(0.0414)	(0.0499)	(0.0161)
Excess Money growth (lagged)	0.0790*	0.0732*	0.122**	0.114**	0.136**	0.0918**	0.118**	0.100*	0.133**	0.0651*
	(0.0433)	(0.0388)	(0.0582)	(0.0571)	(0.0651)	(0.0463)	(0.0520)	(0.0547)	(0.0604)	(0.0393)
Government spending	0.0426	0.0902**	0.0580	-0.120*	0.124*	0.0923*	0.0737*	-0.206**	0.0528	0.0484
	(0.0400)	(0.0376)	(0.0574)	(0.0614)	(0.0668)	(0.0495)	(0.0429)	(0.100)	(0.0827)	(0.0438)
Terms of trade	0.132**	0.115**	0.0502	0.0849	0.0215	0.0898	0.0690	0.159**	0.111	0.0180
	(0.0590)	(0.0576)	(0.0657)	(0.0672)	(0.0726)	(0.0594)	(0.0503)	(0.0665)	(0.0942)	(0.0677)
Openness index	-0.0532*	-0.0940	-0.171**	-0.153**	-0.173**	-0.0864	-0.163**	-0.0842	-0.206**	-0.0705***
	(0.0278)	(0.0719)	(0.0791)	(0.0728)	(0.0836)	(0.0706)	(0.0724)	(0.0716)	(0.0891)	(0.0241)
Foreign direct investment	-0.00673	-0.00336	0.0125	0.00436	0.0109	0.00265	0.00139	-0.0205	0.00683	0.0220
	(0.0105)	(0.0112)	(0.0147)	(0.0136)	(0.0164)	(0.0165)	(0.0143)	(0.0198)	(0.0128)	(0.0186)
Time dummies	YES	YES	YES	YES	YES	YES	YES	YES	YES	
Observations	307	307	307	307	244	249	274	195	221	120
Number of Countries	63	63	63	63	52	54	54	40	62	36
Arellano-Bond test for AR(2) (p-value)	-	0.584	0.672	0.709	0.494	0.462	0.584	0.100	0.276	0.339
Hansen OID test (p-value)	-	0.422	0.187	0.501	0.290	0.529	0.377	0.661	0.342	0.313
R-squared	0.621	-	-	-	-	-	-	-	-	-

Note: The dependent variable is the real effective exchange rate derived by *Bruegel*. The Gini index is instrumented by second lag level and further in first differences equation. Robust standard errors are reported in parentheses. ***, ** and * denote significance at 1, 5 and 10 percent levels.
The null hypothesis of *Arellano-Bond* test is that autocorrelation doesn't exist. The null hypothesis of *Hansen* test is that the instruments, as a group, are uncorrelated with the error term.

coefficient. While the coefficient on the Gini index is less negative than the baseline's coefficient. The latter may suggest that the intensity of a change in income inequality on the real exchange rate is stronger in developed countries than in developing countries. Following the proposed theory, it may arise because the relative consumption expenditure of non-traded to traded goods for the poor is larger in developed countries than in developing countries.

The theory predicts that the real exchange rate appreciates more in low-quality-institution countries than in high-quality-institution countries. Hence, I run the regression model for a sample including countries with a normalized institution index less than 0.7, specifying poor-institution countries. I also exclude the Gini index from the regression model to prevent the real exchange rate from the indirect effect of the windfall income through income inequality change. The results reported in column (6) reveal a larger value of the coefficient on the windfall income (i.e. 0.0268) relative to the value estimated on the full sample (i.e. 0.0261). It seems to confirm the theory anticipated.

Further robustness checks are included in Column (7) and (8) to test the consistency of the results, respectively, in terms of the currency union's effects and different data sources. I run the baseline regression model for a sample of Non-Europe union countries and for the real effective exchange rate sourced by *IMF* [50]. The estimated results reported in column (6) show that the empirical findings are still confirmed and excluding Europe union countries has no effect on our findings. Furthermore, column (7) demonstrates that these findings are independent of a database used for the real effective exchange rate, however, the coefficients are more significant in the estimation on the *Bruegel* database than on the *IMF* database.

The global recession 2008-2013 has been the worst postwar recessions, both in terms of the number of countries affected and the decline in real World GDP per capita. Given that falling the interest rate over the recession period may depreciate the real exchange rate, it seems worthy to check the consistency of the results on non-recession periods. Hence, I run the regression model for a sample including the period 1975-2004. The results presented in column (9) show that the coefficients on the variable of interests are significant and qualitatively the same as the full sample. In other words, the great recession effect affects the results only quantitatively.

The next question that seems worthy to be discussed is the response of the real exchange rate to resource price variations. In this respect, the impact of the commodity price index rather than the windfall income index on the real exchange rate is investigated. Column (10) reports the empirical results. The coefficient on the commodity price index is significant and expresses that a one percent increase in the commodity price index appreciates the real effective exchange rate by about 0.1%. It suggests that the hypothesis is confirmed in terms of both measurement approaches estimating a resource boom. The empirical finding is also in line with recent studies in the commodity prices–exchange rate nexus. Among many research, Jahan-Parvar and Mohammadi (2011) using monthly observation data for 15 oil-exporting countries on the period 1970-2007, Bodart et al. (2012) using monthly observation data for 68 developing countries on

[50] The main departure of *Bruegel* database from *IMF* database is in the calculation of the geometrically weighted average of *CPI* indices of trading partners.

the period 1980-2009 [51], and Ricci et al. (2013) using annual observation data for 27 developing & 21 developed countries on the period 1980-2004 confirm the existence of a long-run positive relationship between commodity prices and the real exchange rate.

I also check the sensitivity of the coefficient on the windfall income to a change in the composition of the samples. Following Mihasonirina and Kangni (2011), I select randomly 98% of the observations (without replacement) and run the baseline regression model (column 4). This process, repeated 250 times, gives values for the coefficient on the variable of interest. I used the same procedure to select randomly 95%, 90% and 85% of observations. The mean and standard deviation of the coefficient for each selected sample are reported in Table 3. The test shows when the regression runs on different sample sizes, the coefficients of the windfall income don't change very much (less than 13%) while their distributions have heavier tails as the sample size shrinks. This may suggest that the coefficient becomes less significant in decreasing sample size.

Table 3: Robustness test for the sensitivity of coefficient of interest

Explanatory variable		98%	95%	90%	85%
Windfall income (lagged)	Coefficient	0.03835	0.03606	0.03383	0.03355
	Standard Deviation	0.00435	0.00637	0.00807	0.00875

Further, an additional robustness check to test the stability of the coefficient on the windfall income is conducted. I first rank the observations in increasing order of the windfall income [52]. Then the regression is run for a sample with the lowest order of windfall income (i.e. less than 5%). In the next step, I add the subsequent observations with larger windfall income and

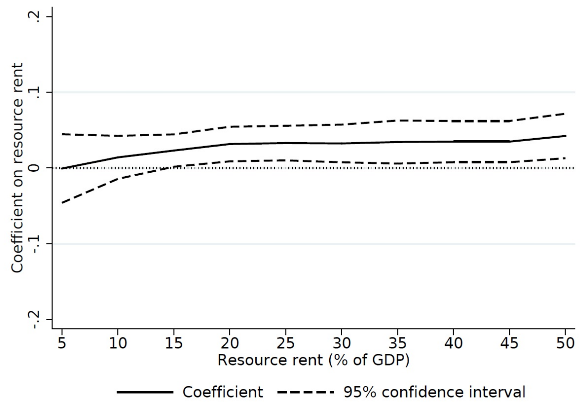


Figure 6: Recursive estimation on the coefficient of interest.

[51] They show that the dominant commodity's price affects the real exchange rate significantly in the long run when the leading commodity export share is at least 20 percent in the country's total exports of merchandises. Further their study show that the larger the share, the stronger the impact.

[52] I ranked the observations in 10 orders: less than 5%, between 5% and 10%, between 10% and 15%, and so on.

rerun the regression. The results, shown in Figure 6, remain positive and significant. They also represent that the coefficient increases slightly as the share of resource rent in GDP becomes larger. It, in turn, suggests that the real exchange rate appreciation is more intense in the natural resource-rich countries than natural resource-poor countries.

3.3 Income inequality response

In this section, I examine the impact of the windfall income on income inequality. Gini index proxies total income inequality and so it is the dependent variable of the dynamic panel data model. The proposed theory clarifies that income inequality is affected by the distribution of resource rent benefit (subsides) between the rich and the poor (skilled and unskilled workers). Furthermore, a natural resource boom induces income inequality indirectly through the real exchange rate appreciation (see. Figure 5). Therefore, the explanatory variables of interest are the windfall income and the real exchange rate. I also include per capita GDP (Ln), Investment ratio, Human Capital index, Openness index, and Institution index as the control variables.

Table 4 displays a set of results. The first column corresponds to the *OLS* estimation. The results for the indirect effect of the windfall income on the Gini index are reported at column (2). The estimated coefficient of interest is negative and significant. This implies that the real exchange rate appreciation is associated with falling income inequality. Following Proposition 1, it intuitively refers to the assumption that the traded sector is relatively more intensive in skilled workers. Since the relative price appreciation increases the factor price used intensively in the non-traded sector and decreases the other, the market-based income inequality falls if the non-traded sector is relatively unskilled worker intensive (i.e. $Stolper-Samuelson$ theorem).

In order to reveal the total direct and indirect effects of a natural resource boom on income inequality, I include the windfall income index in the regression model. Column (3) reports the estimated results. The coefficient enters with a negative sign and is significant at 1 percent. It shows that a one-standard-deviation increase in the windfall income is associated with a decrease in income inequality level by about 6%. Accordingly, as in Goderis and Malone (2011) and, Kim and Lin (2018)[53], the finding suggests that a natural resource boom, on average, tends to reduce income inequality.

Now it is worth studying whether there is a difference between the response of income inequality to a resource boom in a resource-poor economy and in a resource-rich economy. To address this issue, I introduce a dummy variable for the natural resource-poor economics, equal to one when total natural resource rent is less than 5 percent of GDP. Column (4) reports the estimated coefficients of the dummy variable and the interaction between the dummy variable and the windfall. The results conclude that the correlation between the windfall income and income inequality is positive in the resource-poor economies while it is negative in other. From a theoretical viewpoint, Proposition 5 suggests when the distribution of the resource income benefit is less fairly than the distribution of the non-resource income, income inequality rises

[53] Using correlated effects pooled mean group methodology for a sample of developed and developing countries, they find that both oil abundance and oil dependence reduce income inequality.

Table 4: Estimation results for Gini index

Explanatory variable	(1) OLS	(2) Sys.GMM	(3) Sys.GMM	(4) Sys.GMM	(5) Sys.GMM	(6) Sys.GMM	(7) Sys.GMM
	Full Sample	Full Sample	Full Sample	Full Sample	Developing	Poor Instit	Non Europe
Gini index (lagged)	0.841***	0.781***	0.790***	0.821***	0.756***	0.786***	0.771***
	(0.0237)	(0.0500)	(0.0426)	(0.0543)	(0.0747)	(0.0637)	(0.0678)
Real effective exchange rate	-0.159	-0.215***	-0.186***	-0.235***	-0.249***	-0.197***	-0.238***
	(0.124)	(0.0551)	(0.0491)	(0.0827)	(0.0861)	(0.0754)	(0.0900)
Windfall income (lagged)			-0.0592***	-0.177***	-0.0633**	-0.0631***	-0.0639**
			(0.0221)	(0.0481)	(0.0264)	(0.0234)	(0.0250)
Dummy Resource Poor countries				-2.533***			
				(0.656)			
Dummy* Windfall income				0.208***			
				(0.0573)			
GDP per capita (Ln)	-0.305**	-0.393**	-0.359***	-0.119	-0.169	-0.176	-0.207
	(0.137)	(0.162)	(0.128)	(0.170)	(0.213)	(0.147)	(0.194)
Investment ratio	-0.0391**	-0.0470*	-0.0471	0.0180	-0.0656	-0.0329	-0.0709*
	(0.0162)	(0.0279)	(0.0319)	(0.0252)	(0.0449)	(0.0418)	(0.0417)
Human capital index	-0.417	-0.441	-0.632*	-0.736**	-1.033**	-1.102**	-1.027**
	(0.268)	(0.298)	(0.342)	(0.371)	(0.511)	(0.437)	(0.413)
Openness index	-0.00938***	-0.0129***	-0.0126**	-0.0143***	-0.0128	-0.0106	-0.0114
	(0.00316)	(0.00482)	(0.00537)	(0.00462)	(0.00929)	(0.00769)	(0.00836)
Institution index	-0.555	-1.035	-1.075	-0.929	0.484	-1.744	-0.0989
	(0.611)	(0.798)	(0.819)	(0.811)	(1.020)	(1.105)	(0.770)
Time dummies	YES	YES	YES	YES	YES	YES	YES
Observations	396	396	396	396	265	286	291
Number of Countries	78	78	78	78	55	60	56
Arellano-Bond test for AR(2) (p-value)	-	0.330	0.328	0.306	0.132	0.158	0.151
Hansen OID test (p-value)	-	0.740	0.674	0.447	0.507	0.450	0.501
R-squared	0.919	-	-	-	-	-	-

Note: The dependent variable is the Gini index of *EHII* data set. The real effective exchange rate and windfall income (lagged) are instrumented by first lag level and second lag level in first differences equation, respectively. The control variables except the Institution index are instrumented by first lag level. Robust standard errors are reported in parentheses. ***, ** and * denote significance at 1, 5 and 10 percent levels.
The null hypothesis of *Arellano-Bond* test is that autocorrelation doesn't exist. The null hypothesis of *Hansen* test is that the instruments, as a group, is uncorrelated with the error term.

as the windfall income increases. Hence, the empirical findings seem to be consistent with the theoretical prediction. Since one is more likely that in the natural resource-poor economies, a group of elites appropriates a larger share of resource income benefits or private companies, exploiting the natural resource, distribute their income among a small group of shareholders. Thus, the assumption that, in the natural resource-poor economies, the distribution of resource income is more unevenly than the distribution of non-resource (market-based) income seems to be reliable. While in the natural resource-rich economies, direct subsidies or the spillover effects of investing the resource income on education, infrastructure or health care seem to make more evenly the distribution of resource income benefit than the distribution of non-resource income. Column (5) trough (7) at Table 4 contain robustness checks of the baseline model for heterogeneity across countries. I first run the specification model for the developing economies. Further, a robustness analysis of the baseline model is examined for heterogeneity in terms of institutional quality. Given that the quality of institution is more likely to affect income inequality, I split countries with a high normalized institution index, larger than 0.7, and run the regression. For both analysis, there is a significantly negative impact on income inequality from both the real exchange rate and the windfall income, the same as the results for the baseline model (i.e. column (3)). The value of the coefficient on the windfall income is more negative in the developing economies and the poor institution countries' samples (equal -0.0633 and -0.0631 respectively) than the full sample (equal -0.0592). This may indicate a larger gap in expenditure shares in developing or poor-institution countries than others. Since, regarding the theory, a natural resource boom leads to a stronger real exchange rate appreciation in countries with a larger gap in expenditure shares (i.e. $\theta_L - \theta_S$) which, in turn, results in more falling in income inequality. In addition, column (7) represents the estimation results for non-European countries. The results are qualitatively similar to those for the full sample. It indicates currency union effects plays no role on the relationship between the variables of interest and the dependent variable. Given that the unskilled labor intensity in the non-traded sector is larger in a sample excluding European countries than in the full sample, a more negative value of the coefficient on both variables for the restricted sample seems to be explainable.

I further examine whether the estimation results are robust in terms of different income inequality databases. The first measure of the Gini coefficient is taken from the Standardized World Income Inequality Database (*SWIID*) (Solt, 2016) [54]. The results represent in column (1) of Table 5. Those show that the value for both explanatory variables of interest are negative, however, it is significant only for the coefficient on the windfall income. Also, the coefficient on the windfall income is more intensive when the *SWIID* database is used than when the *EHII* database is done. The second database is the World Bank's *PovcalNet* database published in the World Development Indicators *WDI*. Column (2) reports the results. Although both coefficients on the real effective exchange rate and the windfall income have negative signs, they are less significant. It may arise because *WDI* database publishes both consumption and income measures of inequality. The difference in measures would be very large and so the significance

[54] I use inequality in market (pre-tax, pre-transfer) income.

Table 5: Robustness checks for Gini index

Explanatory variable	(1) Sys.GMM	(2) Sys.GMM	(3) Sys.GMM	(4) Sys.GMM	(5) Sys.GMM	(6) Sys.GMM
	SWIID	*WDI*	*Commodity*	*Commodity*	Full sample	*Commodity*
Gini index (lagged)	0.851***	0.880***	0.673***	0.630***	0.772***	0.768***
	(0.0443)	(0.0454)	(0.115)	(0.127)	(0.0965)	(0.101)
Gini index (Second lagged)					-0.0310	-0.0712
					(0.0731)	(0.0859)
Real effective exchange rate	-0.0990	-0.575*	-0.0675	0.545	-0.296*	0.535
	(0.0983)	(0.314)	(0.908)	(1.212)	(0.159)	(1.087)
Windfall income (lagged)	-0.0823***	-0.00250			-0.0691***	
	(0.0274)	(0.0772)			(0.0196)	
Commodity Price index (lagged)			-0.574**	-1.631***		-0.900**
			(0.278)	(0.336)		(0.454)
Dummy Resource Poor countries				-2.917***		
				(0.855)		
Dummy* Commodity Price				1.270*		
				(0.656)		
GDP per capita (Ln)	0.335	0.468	-0.636	-0.413	-0.370	-0.358
	(0.236)	(0.554)	(0.390)	(0.362)	(0.237)	(0.307)
Investment ratio	-0.0195	-0.0278	-0.0311	0.00107	-0.0331	-0.0775
	(0.0393)	(0.0702)	(0.0528)	(0.0768)	(0.0359)	(0.0756)
Human capital index	-0.977*	-1.501	-1.434**	-1.863**	-0.953***	-1.885***
	(0.508)	(1.190)	(0.568)	(0.904)	(0.296)	(0.683)
Openness index	-0.00469	0.00225	-0.0177*	-0.0151**	-0.0114***	-0.0141**
	(0.00443)	(0.00975)	(0.00925)	(0.00672)	(0.00414)	(0.00627)
Institution index	1.378*	0.207	-0.644	-0.918	-1.824*	-0.975
	(0.789)	(2.255)	(1.357)	(1.283)	(0.978)	(1.680)
Time dummies	YES	YES	YES	YES	YES	YES
Observations	359	211	139	139	318	130
Number of Countries	77	55	42	42	78	42
Arellano-Bond test for AR(2) (p-value)	0.487*	0.953	0.384	0.431	0.173	0.474
Hansen OID test (p-value)	0.238	0.492	0.171	0.318	0.499	0.281

Note: The dependent variables for estimation results, reported in columns (1) and (2), are the Gini index of *SWIID* database and *WDI* database, respectively. While it is the Gini index of *EHII* database in the rest of columns. The real effective exchange rate and windfall income (lagged) are instrumented by first and deeper lags level and second lag level in first differences equation, respectively. The control variables except the Institution index are instrumented by first lag level. Robust standard errors are reported in parentheses. ***, ** and * denote significance at 1, 5 and 10 percent levels.
The null hypothesis of *Arellano-Bond* test is that autocorrelation doesn't exist. The estimation results for *SWIID* dataset, marked by *, report the AR(3) instead of AR(2). The null hypothesis of *Hansen* test is that the instruments, as a group, is uncorrelated with the error term.

of coefficients may reduce.
Column (3) reports the results when replacing the measurement of a resource boom by the Commodity price index. The coefficient on the commodity price index enters negative and is statistically significant at 5 percent. The effect is substantial and it can be translated as a function of a country's dependence on price variation of the commodity exports. The finding shows that income inequality's response to a resource boom is qualitatively independent of how to measure this boom. Further, similar to what discussed in column (4) of Table 4, I check the heterogeneity between countries' groups in terms of the dependency on the commodity exports. Column (4) of Table 5 represents the results when adding the interaction term between the dummy for resource-poor countries and the commodity price index to the specification of column (3). This shows that an increase in the commodity price index decreases the income inequality in both groups of countries and the effect is larger in resource-rich countries than resource-poor countries. This is contrary to the previous finding, reported in column (4) of Table 4, that a resource revenue boom makes the income distribution be worse.
Using panel cointegration methodology for a sample of 90 countries between 1965 and 1999, Goderis and Malone (2011) analyze the short and long-run effects of commodity prices on the *Gini* index. Their estimates show that income inequality falls in the short run immediately after a resource boom and then rises steadily over time until the initial impact of the resource boom disappears. The difference in the adopted empirical approach in this paper from their approach is to apply the five-year average of variables in order to ignore business cycle fluctuations. Although the approach applied in this paper doesn't have well efficient to capture the short-run effects of the resource measurement on the income inequality index, I address this question by including the second lagged *Gini* index in the regression model. These first and second lags of the dependent variables can be translated as the intermediate and long-run effects. The results of the regression equation including one of the proxies for the resource rent are respectively reported in columns (5) and (6). The coefficients of the resource measurement index are also similar to before. Those also show that for both resource measurement proxies the coefficients on the first and second lagged *Gini* index enter with positive and negative signs respectively but they are statically significant at 1% only for the first lagged. The statistical insignificance of the coefficients on the second-lagged dependent variables may originate from a small number of periods for each country which, in turn, may undermine the reliability of the empirical results. Nevertheless, calculations seem to confirm the long-term effects (i.e. the second lagged) is smaller than the intermediate-term (i.e. the first lagged) [55]. More precisely, the intermediate and long terms effects are equal to -0.303 and -0.267 in a regression model including the windfall income index. While they are equal to -3.879 and -2.968 in a model including the commodity price index. Those may clearly reflect that falling income inequality is moderated over time, similar to Goderis and Malone (2011) finding.
I also check the sensitivity of coefficients of interest to a change in sample size the same pro-

[55] For the following regression model: $x_t = ax_{t-1} + bx_{t-2} + cz_t$, the intermediate and long terms effects are reflected by $\frac{dx}{dz} = \frac{c}{1-a}$ and $\frac{dx}{dz} = \frac{c}{1-(a+b)}$, respectively.

cedure as before. The average coefficient and standard deviation for each selected sample are reported in Table 6. The analysis shows that reducing the random sample size has less effect on the average coefficient for the real effective exchange rate (less than 20%) however, the base of the normal distribution widens. In contrast to this, the estimated coefficient on the windfall income weakens rapidly as the sample size shrinks. This suggests that the coefficient estimated for the baseline model (i.e. column (3) at Table 4) may hide significant country heterogeneity. It is partly clarified by opposite impacts of the windfall income on income inequity in the resource-poor and resource-rich economies.

An additional robustness check is a recursive estimation to test the stability of the coefficients

Table 6: Robustness test for the sensitivity of coefficients of interest

Explanatory variable		98%	95%	90%	85%
Real effective exchange rate	Coefficient	-0.19326	-0.19353	-0.17653	-0.15712
	Standard Deviation	0.03671	0.07515	0.14097	0.23469
Windfall income (lagged)	Coefficient	-0.05208	-0.04611	-0.03875	-0.02380
	Standard Deviation	0.01601	0.02071	0.02899	0.03603

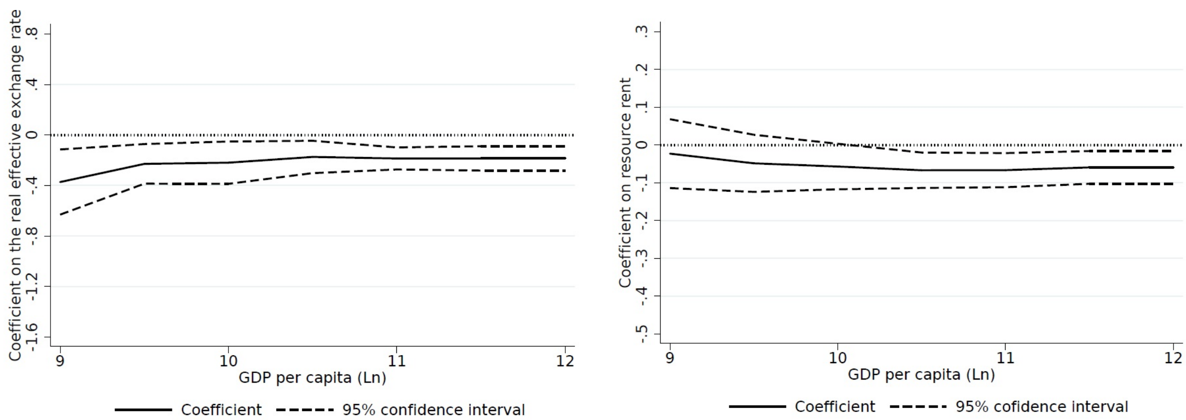


Figure 7: Recursive estimation on the coefficients of interest.

of interest for cross-country heterogeneity. The observations are ranked in increasing GDP per capita (Ln). The regression is run for a sample having GDP per capita (Ln) less than 9 [56]. Subsequent observations with larger GDP per capita are added and the regression is rerun. The same procedure is replicated as long as the full sample is covered. The results show in Figure 7. The coefficients on both the resource rent and the real exchange rate are negative and significant at each level of GDP per capita. This demonstrates that the negative correlation between variables of interest and income inequality is stable across country income level.

[56] It covers about 50% of total observations.

3.4 The intensity of the Dutch disease

The third question of interest is the impact of a natural resource boom on the relative productivity ratio. The lack of enough data for the productivity level at the manufacturing and service sectors, proxies for the traded and non-traded sectors, motivates us to search a proxy for the relative productivity ratio. Following the theory, the labor allocation of skilled and unskilled workers return to their initial level along the transition path so that they would be constant at steady-state (i.e. Equation 15). Accordingly, the relative sectoral output, at steady-state, will be a function of the steady-state relative productivity ratio ϕ^* [57]. Hence, the relative sectoral output seems to be a reliable proxy to capture the impact of a natural resource boom on the relative productivity levels.

The income level in the manufacturing sector to that in the service sector (in constant price) is introduced as the dependent variable in the regression model. The windfall income proxy and Gini index are the explanatory variables of interest. Since the theory predicts that the resource income through the channels of income inequality can affect the relative productivity ratio. In addition, per capita GDP (Ln), Investment ratio, Human Capital index, Openness index, Government spending, and Institution index are included to control the regression model.

The estimated results are shown in Table 7. The results of *OLS* regression are reported in column (1) as a benchmark. Column (2) corresponds to the baseline model to test the impact of the windfall income on the dependent variable. The coefficient on the windfall income has a negative sign and is significant. The finding is consistent with the empirical literature, as in Amiri et al. (2019) [58], and the theoretical prediction. A natural resource boom shifts the skilled labor force away from the traded sector and into the non-traded sector. The productivity level shrinks faster in the traded sector than in the non-traded sector [59], leading to a decrease in the relative productivity ratio, to induce a countervailing movement of labor which in turn gradually brings back the allocation of labor to a constant long-run equilibrium level.

The estimations further show that there is a negative correlation between income inequality and the relative sectoral output. More precisely, the dependent variable, on average, increases by about 2% per year as one-standard-deviation decreases in the Gini index. The finding is not following the proposed theory. Since the theory predicts that falling income inequality due to an increase in the resource rent tends to decrease more the relative productivity ratio (i.e. exacerbating the Dutch disease). This inconsistent result may reflect the weak efficiency of the relative sectoral output level in being a proxy of the relative productivity level. It also may arise because of country heterogeneity in terms of income equality. The latter issue is more clarified when an interaction term between variables of interest is included in the baseline model.

57 $\frac{X_T}{X_N} = \frac{A_T S_T^{\beta} L_T^{1-\beta}}{A_N S_N^{\alpha} L_N^{1-\alpha}} = \chi \phi^*$, where $\chi = \chi\,(S_N^*, L_N^*)$.

58 They evaluate the impacts of natural resource rents on the relative sectoral output in terms of the quality of institutions. Under the panel data model of a sample including data of 28 natural resource-rich countries for the period of 2000-2016, they find that the relative output level decreases as the natural resource rents increase. Further, their estimation results show that the higher the level of institutional quality, the smaller the size of the impact.

59 An alternative is that the productivity level grows faster in the non-traded sector than in the traded sector. Nevertheless, the empirical findings don't support this interpretation.

Table 7: Estimation results for the relative sectoral output

Explanatory variable	(1) OLS	(2) Sys.GMM	(3) Sys.GMM	(4) Sys.GMM	(5) Sys.GMM	(6) Sys.GMM	(7) Sys.GMM	(8) Sys.GMM
	Full Sample	Full Sample	Full Sample	Full Sample	Developing	Poor Instit	Commodity	Commodity
Relative sectoral output (Ln) (lagged)	0.906***	0.741***	0.857***	0.849***	0.694***	0.737***	0.842***	0.996***
	(0.0194)	(0.0581)	(0.0684)	(0.0636)	(0.107)	(0.0608)	(0.106)	(0.121)
Windfall income (lagged)	-0.00196*	-0.00495**	-0.00644	-0.0134**	-0.00438**	-0.00494**		
	(0.00115)	(0.00203)	(0.0201)	(0.00601)	(0.00199)	(0.00220)		
Gini index	-0.00837***	-0.0209***	-0.0211***	-0.0191***	-0.0187***	-0.0164***	-0.0185***	-0.0210**
	(0.00217)	(0.00532)	(0.00618)	(0.00551)	(0.00684)	(0.00607)	(0.00691)	(0.00920)
Windfall income*Gini index			0.000085					
			(0.000438)					
Dummy Resource Poor countries				-0.157*				-0.302**
				(0.0838)				(0.139)
Dummy* Resource measurement				0.0106**				0.159**
				(0.00518)				(0.0729)
Commodity Price index (lagged)							-0.0432**	-0.190***
							(0.0216)	(0.0676)
GDP per capita (Ln) (lagged)	-0.0145	-0.0385*	-0.0385*	0.000170	-0.0244	-0.0165	-0.0338	0.0655
	(0.0110)	(0.0228)	(0.0213)	(0.0331)	(0.0252)	(0.0255)	(0.0288)	(0.0706)
Investment ratio	0.00212	-0.000548	0.00126	0.00185	-0.00110	-0.000702	-0.00117	0.000112
	(0.00140)	(0.00251)	(0.00331)	(0.00285)	(0.00313)	(0.00258)	(0.00292)	(0.00331)
Human capital index	-0.0222	-0.0862**	-0.0591*	-0.0924**	-0.0776**	-0.0684	-0.0665	-0.140**
	(0.0209)	(0.0421)	(0.0317)	(0.0399)	(0.0359)	(0.0500)	(0.0589)	(0.0688)
Openness index	-0.000074	0.000403	-0.000134	0.000124	0.000671	0.000686	0.000076	-0.00003
	(0.000256)	(0.000622)	(0.000643)	(0.000571)	(0.000525)	(0.000582)	(0.000569)	(0.00100)
Government spending	-0.00587***	-0.0158***	-0.0116**	-0.0117***	-0.0138***	-0.0128***	-0.0118	-0.00828
	(0.00212)	(0.00366)	(0.00475)	(0.00337)	(0.00394)	(0.00330)	(0.00742)	(0.00851)
Institution index	-0.0840*	-0.117	-0.105	-0.163*	0.000611	-0.128	-0.0717	-0.285*
	(0.0478)	(0.0915)	(0.0933)	(0.0949)	(0.105)	(0.153)	(0.100)	0.0655
Time dummies	YES	YES	YES	YES	YES	YES	YES	YES
Observations	313	313	313	313	213	226	128	128
Number of Countries	68	68	68	68	46	51	40	40
Arellano-Bond test for AR(2) (p-value)	-	0.859	0.843	0.842	0.954	0.960	0.264	0.06
Hansen OID test (p-value)	-	0.331	0.164	0.336	0.624	0.614	0.121	0.127
R-squared	0.914	-	-	-	-	-	-	-

Note: The dependent variable is the relative sectoral output (in *Ln* forms). The windfall income (lagged) and GDP per capita (lagged) are instrumented by second lag level, while the Gini index is instrumented by first lag level in first differences equation. Robust standard errors are reported in parentheses. ***, ** and * denote significance at 1, 5 and 10 percent levels.
The null hypothesis of *Arellano-Bond* test is that autocorrelation doesn't exist. The null hypothesis of *Hansen* test is that the instruments, as a group, is uncorrelated with the error term.

Column (3) shows the estimated results for this regression. The coefficients on the windfall income and the interaction term are insignificant. However they, respectively, enter with negative and positive signs, meaning that falling income inequality intensifies the negative effect of the windfall income on the relative sectoral output (i.e. the relative sectoral productivity), the same as the theory's prediction.
In the preceding sub-section, I found that the response of income inequality to a natural resource boom is different in terms of the dependency on the natural resource (i.e. country heterogeneity). Hence, study the impact of a resource-dependent proxy on the relative productivity ratio would be a useful evaluation to make sure that the windfall income through the transmission channel of income inequality induces the intensity of the Dutch disease. To address this issue, like the preceding section, I include a dummy variable for the natural resource-poor economics as well as an interaction term between the dummy variable and the windfall income. Column (4) reports the results. The estimated results show that the adverse effect of the windfall income on the relative sectoral output is less intensive in the resource-poor economies. It is more likely to be because of rising income inequality in the resource-poor economies which attenuates the negative value of coefficient on the windfall income (i.e. Proposition 6).
As the first robustness check, I test the baseline model (i.e. column (2)) for developing economies and countries with low quality of the institution. The results are reported in columns (5) and (6), respectively. The main variable of interest (windfall income) enters with a negative sign and is significant, the same as the baseline model however, the coefficient is slightly smaller. This generally suggests that the correlation between the explanatory variable of interest and the dependent variable is probably robust in terms of country heterogeneity.
The results in the previous subsections imply the commodity price index being important in explaining the resource curse. Hence, I replace the windfall income with the commodity price index and run the regression models to analyze firstly the relationship between the commodity price and the relative sectoral output and secondly the effect of the resource dependency level on this nexus. The results reported in columns (7) and (8) are similar to before. Those suggest that notwithstanding the resource measurement index, a resource boom decreases the relative sectoral output level and this negative effect is less intensive in resource-poor countries than resource-rich countries.
I further evaluate the sensitivity of the variable of interest to a change in the sample size. A repeated process for randomly select of 98%, 95%, 90% and 85% of observations gives values of the coefficient on the resource rent. Table 8 represents the values. When the baseline model is run on different sample sizes, the values of the coefficient remain close to full sample coefficients (reduce at most 23%), while their normal distributions have heavier tails. In other words, the significance of the coefficient weakens as the sub-sample size decreases.

Following the same procedure proposed in section 3.2 and 3.3, I also conduct a recursive estimation to evaluate the accuracy of the results. I estimate the coefficient on the resource rent for sample of observation ranked in increasing order of GDP per capita. Figure 8 shows the

Table 8: Robustness test for the sensitivity of coefficient of interest

Explanatory variable		98%	95%	90%	85%
Windfall income (lagged)	Coefficient	-0.00481	-0.00456	-0.00399	-0.00372
	Standard Deviation	0.00102	0.00147	0.00216	0.00247

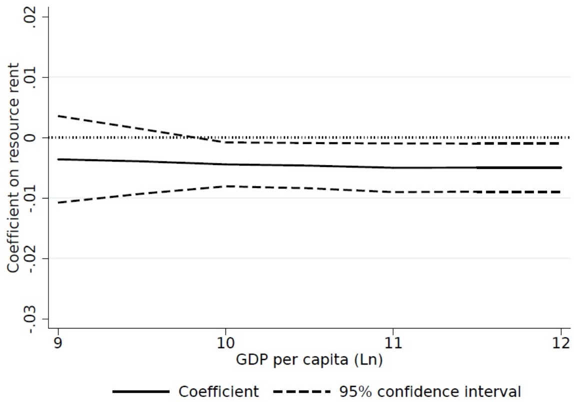


Figure 8: Recursive estimation on the coefficient of interest.

result. It may demonstrate the significance and stability of negative coefficients on the windfall income across countries in terms of income level.

3.5 The sectoral growth

The conceptual mechanism of the model (Figure 5) illustrates that the windfall income directly through a change in the real exchange rate induces the sectoral economic growth. It can also motivate growth indirectly through the channel of income inequality. Therefore, an evaluation of the sectoral economic growth's response to the total effects of a natural resource boom seems to be useful to figure out the mechanism of the model. Hence, the main contribution of this sub-section is to test the interaction between the windfall income and income inequality on sectoral economic growth.

Sectoral GDP per capita level in constant price is the dependent variable and the Gini index, as well as the windfall income proxy, are the explanatory variables of interest. I also include Population Growth, Investment ratio, Human capital index, Openness index, Government spending and Rule of Law indicator to control the regression model. In line with Forbes (2000), I use the difference GMM to estimate the regressions. Since the System GMM estimator yields less clear and significant results.

Table 9 reports the results. To save space, the results estimated by *OLS* are not reported. Columns (1) and (3) represent, respectively, the baseline specifications' results for the manufacturing and the service sectors' regression models. The estimated coefficients for the resource-dependence index enter with positive signs while those for the Gini index enter with opposite signs. Neverthe-

Table 9: Estimation results for the sectoral growth

Explanatory variable	(1) Diff.GMM	(2) Diff.GMM	(3) Diff.GMM	(4) Diff.GMM	(5) Diff.GMM	(6) Diff.GMM	(7) Diff.GMM	(8) Diff.GMM
	Full Sample	Full Sample	Full Sample	Full Sample	Commodity	Commodity	Commodity	Commodity
	Manufacturing	Manufacturing	Service	Service	Manufacturing	Manufacturing	Service	Service
Sector income level per capita (Ln) (lagged)	0.686***	0.630***	0.506***	0.624***	0.631***	0.586***	0.653***	0.648***
	(0.0994)	(0.129)	(0.104)	(0.130)	(0.106)	(0.151)	(0.0805)	(0.0944))
Windfall income (lagged)	0.00368	-0.0515**	0.00363	-0.0626***				
	(0.00712)	(0.0215)	(0.00617)	(0.0240)				
Commodity Price index (lagged)					-0.263**	-0.754***	0.133**	-0.225
					(0.130)	(0.282)	(0.0658)	(0.450)
Gini index	-0.00859	-0.0194***	0.00744	-0.0247**	-0.0153**	-0.0384**	0.00764	-0.00842
	(0.00903)	(0.00738)	(0.00823)	(0.0116)	(0.00741)	(0.0153)	(0.00799)	(0.0156)
Resource measurement*Gini index		0.00114**		0.00150***		0.0180**		0.00922
		(0.000477)		(0.000536)		(0.00735)		(0.0147)
Population Growth	-0.0516*	-0.0246	-0.00584	-0.0212	-0.0168	-0.0486	-0.0268	-0.0483
	(0.0268)	(0.0254)	(0.0269)	(0.0237)	(0.0409)	(0.0394)	(0.0476)	(0.0509)
Investment ratio	0.0134***	0.0138***	0.0104***	0.00991***	0.0103**	0.00997***	0.0138***	0.0141***
	(0.00322)	(0.00260)	(0.00219)	(0.00294)	(0.00465)	(0.00380)	(0.00214)	(0.00316)
Human capital index	0.000926	-0.0162	0.192	-0.0515	0.0922	0.0663	0.0929	0.0633
	(0.180)	(0.202)	(0.129)	(0.129)	(0.139)	(0.254)	(0.127)	(0.160)
Openness index	0.00165	0.000548	-0.000027	-0.00120	-0.000918	0.000279	-0.00190	-0.00161
	(0.00144)	(0.00141)	(0.000888)	(0.00115)	(0.00277)	(0.00164)	(0.00144)	(0.00109)
Government spending	-0.0177	-0.0267*	-0.0186	-0.0364**	-0.0158	-0.0247	-0.0189	-0.0182
	(0.0137)	(0.0146)	(0.0166)	(0.0146)	(0.0185)	(0.0186)	(0.0138)	(0.0183)
Rule of Law indicator	0.0970	0.0479	0.0657	0.0821	-0.176	0.0196	0.0381	-0.0287
	(0.197)	(0.185)	(0.0491)	(0.0677)	(0.125)	(0.180)	(0.0560)	(0.137)
Time dummies	YES	YES	YES	YES	YES	YES	YES	YES
Observations	255	255	258	258	88	88	94	94
Number of Countries	68	68	69	69	39	39	41	41
Arellano-Bond test for AR(2) (p-value)	0.898	0.558	0.628	0.229	0.496	0.360	0.391	0.287
Hansen OID test (p-value)	0.332	0.579	0.146	0.261	0.372	0.232	0.376	0.124

Note: The dependent variable is GDP per capita level in each sector (in constant price). In the estimation of the manufacturing sector regression model, the Gini index and the windfall income (lagged) are instrumented by first and second lags level in first differences equation, respectively. While to estimate the service sector regression, they are instrumented by second lags level in first differences equation.Robust standard errors are reported in parentheses. ***, ** and * denote significance at 1, 5 and 10 percent levels.

The null hypothesis of *Arellano-Bond* test is that autocorrelation doesn't exist. The null hypothesis of *Hansen* test is that the instruments, as a group, are uncorrelated with the error term.

less, both are insignificant. These results are neither consistent with the recent empirical studies (e.g. Sachs and Warner (1995), Gylfason and Zoega (2002) and Goderis and Malone (2011)) nor the proposed theory. These may reflect a strong interaction between the variables of interest.

To address this problem, I include an interaction term between the windfall income index and the Gini index. The results for the manufacturing and service sectors are respectively reported at columns (2) and (4). For both regression models, the estimated coefficients on the windfall income and the interaction term are significant and enter with negative and positive signs, respectively. Consistently with the theory (Proposition 9), the estimations imply that the adverse effect of the windfall income on both sectoral economic growth rate attenuates as the Gini index increases. These support the theoretical finding that falling inequality intensifies the adverse effects of a natural resource boom on sectoral growth rate. Further, following Proposition 9, this may indirectly confirm the fundamental assumption that skilled workers allocate a larger expenditure share on traded goods than the unskilled workers do.

In addition, the regression coefficients on a natural resource boom indicate that the non-traded sector (proxied by the service sector) shrinks as the windfall income level goes up. It may demonstrate that the learning generated in the traded sector and spilled over to the non-traded sector is the dominant driving force of productivity growth in the non-traded sector (i.e. Proposition 8).

One more thing that seems to be worth studying is the dependency of the results in terms of the heterogeneity across the resource measurement indices. Columns (5) through (8) illustrate the estimation results when regression models include the commodity price index. The coefficients on the commodity price index are significant at 5% for both sectoral regression models, while their signs are different (see. columns (5) and (6)). The latter point indicates that resource price appreciation tends to shrink the manufacturing sector similar to the case related to the windfall income, however, it tends to expand the service sector, contrary to the case related to the windfall income. These reveal that the reaction of the service sector differs in terms of the driving forces of booming resource rent. Increasing resource dependency decelerates the expansion of the service sector while rising the resource price accelerates it. In following the theory, this may point out that productivity improvement in the service sector due to making more expensive the energy price along with the commodity price appreciation, at least in resource-poor countries, may lead to being stronger the direct effect of learning generated by the non-traded sector relative to the spillover effect. Therefore, high-skilled labor reallocation due to the real exchange rate appreciation tends to expand the service sector. Nevertheless, the magnitude of the coefficient is larger for the manufacturing sector than for the service sector. Hence, an appreciation in the commodity price causes the relative sectoral output level to go down, similar to the main case and the presented theory.

Columns (6) and (8) show the effect of income inequality on the response of each sectoral level to a resource price variation. The significant results for the manufacturing sector's regression model are also similar to before. Falling income inequality exacerbates the productivity growth deceleration in the manufacturing sector. On the contrary, the results for the service sector's regression model are insignificant. Regarding Proposition 9, the latter statically insignificant result may be because of heterogeneity across countries in terms of productivity improvement in the service sector due to commodity price appreciation.

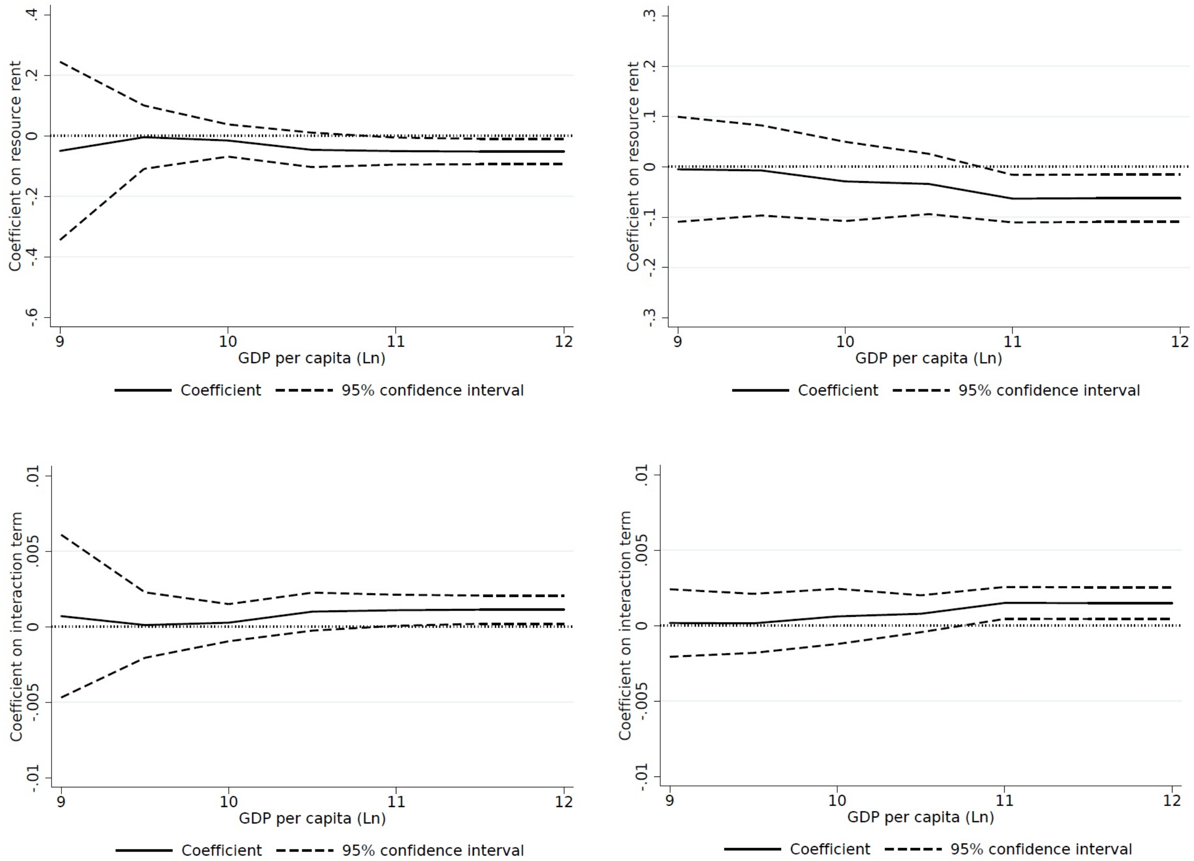
The first robustness test is a sensitivity analysis of a change in the sample size. In the same procedure described in the preceding sub-sections, I estimate the values and standard-deviations of the explanatory coefficients for randomly select of 98%, 95%, 90% and 85% of observations. Table 10 shows the results for the selected samples. When the regressions run on 98% of the sample, the estimated coefficients remain close to the full sample coefficients. While the analysis shows when the regression runs on other sample sizes, the coefficients decrease and their distributions widen slightly. These may state less significance of the coefficients as the sample size shrinks.

The final robustness test is included to test the stability of coefficients on the direct and indirect

Table 10: Robustness test for the sensitivity of coefficients of interest

	Manufacturing				
Explanatory variable		98%	95%	90%	85%
Windfall income (lagged)	Coefficient	-0.05145	-0.04808	-0.04454	-0.03795
	Standard Deviation	0.0111169	0.01794	0.02438	0.03362
Windfall income*Gini	Coefficient	0.00115	0.00110	0.00103	0.00086
	Standard Deviation	0.0002508	0.00039	0.00054	0.00076
	Service				
Explanatory variable		98%	95%	90%	85%
Windfall income (lagged)	Coefficient	-0.05343	-0.04426	-0.03553	-0.02519
	Standard Deviation	0.01531	0.02106	0.02686	0.03074
Windfall income*Gini	Coefficient	0.00131	0.00111	0.00090	0.00066
	Standard Deviation	0.00033	0.00046	0.00059	0.00069

effects of the windfall income on the sectoral economic growth rates. The observations are firstly ranked in increasing order of GDP per capita (Ln). Then the process is as follows: start with a sample of low GDP per capita and run the regressions, add subsequent observation with larger GDP per capita to the sample and rerun the regression, continue the process as long as total observations are covered. The results show that the coefficients on the resource rent and the interaction term remain significant and are respectively, negative and positive in both regression models for the manufacturing and service sectors (Figure 9). In addition, it seems to be useful to investigate whether the direct and indirect effects of the windfall income on the sectoral economic growth rate are different in the developing and developed economies. The negative coefficient on the resource rent slightly decreases, while the positive coefficient on the interaction term slightly increases in both sectors. It states that both direct and indirect effects of a natural resource boom on sectoral growth rate are stronger in the developed economies than in the developing economies.



(c) Manufacturing sector (d) Service sector

Figure 9: Recursive estimation on the coefficients of interest.

4 Conclusion

The paper acknowledges theoretically and empirically that income inequality plays a key role in the economic performance of the natural resource-dependent countries. The main contribution of this paper is to analyze how income inequality responds to a natural resource boom and how a combination of income inequality and resource rent motivates the intensity of the Dutch disease. In theory, I develop a two-sector growth model in which each sector employs skilled and unskilled workers. Workers' groups allocate different consumption expenditure shares on traded and non-traded goods. The gap in expenditure shares captures the feedback of a change in income inequality on economic growth. I have analyzed the model in short-run under a comparative static and in long-run under a dynamic approach driven by learning-by-doing (LBD) model.
The present study yields a number of theoretical findings. In the short-run study, a permanent rise in the windfall income leads respectively to an appreciation in the real exchange rate, a

reallocation in the factor inputs, a shrinkage in the traded sector and a deceleration in economic growth. By the way, a change in inequality depends on the factor intensity and the distribution of the resource rent benefits (subsidies) between workers' groups. Income inequality falls if the traded sector is relatively intensive in skilled workers and the resource rent benefits are distributed more evenly than the real wage between workers' groups. In the long-run study, a change in income inequality is only driven by a resource boom. In response to a windfall income boom, income inequality falls (rises) if the relative resource rent benefit of skilled workers is smaller (larger) than the relative real wage. In addition, falling income inequality deepens the Dutch disease if the skilled workers, with respect to unskilled workers, allocate a larger expenditure share for traded goods.

Consistently with the theory, I estimate a panel data study to evaluate the theoretical predictions. In this respect, the data for 79 countries over the period 1975-2014 are collected. I apply the first-differenced and the system GMM approaches to estimate dynamic panel data regressions. The impact of a natural resource boom on the real exchange rate and then the response of income inequality to a change in the windfall income are examined. Further, I estimate the effect of interaction between a natural resource boom and income inequality on the intensity of the Dutch disease as well as the sectoral economic growth. These empirical studies represent some clear evidence in supporting the crucial role of income inequality in economic performance of the resource-dependent countries. A natural resource boom reduces income inequality and falling income inequality is associated with a more intensive natural resource curse.

Appendix A Proof of Proposition 3

The proposition summarizes the short-term effect of the gap in expenditure shares between workers' groups on the real exchange rate. To find the vertical shift of *NN*-curve, I need to compute to what extent the real exchange rate appreciates in the response of a natural resource boom, $\frac{dP}{dR}$ of Equations 10b. The result will be:

$$\left[1-\frac{P}{\sigma\Psi}\frac{d\Psi}{dP}\right]\frac{dP}{dR}=\frac{P}{\sigma\left(S_T^{\beta}L_T^{1-\beta}+R\right)}+\frac{P}{\sigma\Psi}\frac{d\Psi}{dI}\frac{dI}{dR}. \qquad A.1$$

Where $\frac{dI}{dR}$ follows the Proposition 1. In addition, $\frac{d\Psi}{dP}=\frac{\Psi}{P}\left(\sigma-\sigma_c\right)$ and $\frac{d\Psi}{dI}=\frac{\theta_S e_S^{\sigma-1}-\theta_L e_L^{\sigma-1}}{(1-\Psi)^2}$. The latter result indicates that $\frac{d\Psi}{dI}>0$ if $\theta_S>\theta_L$ and $\frac{d\Psi}{dI}<0$ if $\theta_S<\theta_L$. Replacing the results of $\frac{d\Psi}{dP}$ and $\frac{d\Psi}{dI}$ in Equation $A.1$ gives:

$$\frac{dP}{dR}=\frac{P}{\sigma_c\left(S_T^{\beta}L_T^{1-\beta}+R\right)}+\frac{P}{\sigma_c\Psi}\frac{\theta_S e_S^{\sigma-1}-\theta_L e_L^{\sigma-1}}{(1-\Psi)^2}\frac{dI}{dR}. \qquad A.2$$

It verifies when $\theta_L=\theta_S\Rightarrow e_L=e_S$ and $\sigma_c=\sigma\Rightarrow\frac{d\Psi}{dI}=0\Rightarrow\left(\frac{dP}{dR}\right)_{\theta_L=\theta_S}=\frac{P}{\sigma\left(S_T^{\beta}L_T^{1-\beta}+R\right)}$. So the size of a vertical shift in *NN*-curve depends on the sign of $\frac{dI}{dR}$ and the gap in expenditure shares on non-traded goods (i.e. θ_L and θ_S).

Appendix B Proof of Proposition 4

The gap in the price elasticity of demand for traded and non-traded goods is defined as: $\sigma_c\equiv\varepsilon_{T,P}-\varepsilon_{N,P}=\sigma+(\gamma_T-\gamma_N)\left(\varepsilon_{N,P}^L-\varepsilon_{N,P}^S\right)$, where $\gamma_J\equiv\frac{C_J^L}{C_J}$; $J=T,N$ is the share of good J allocated to unskilled workers and $\varepsilon_{N,P}^i\equiv\frac{P}{C_N^i}\frac{\partial C_N^i}{\partial P}=-\left(\frac{\sigma(1-\theta_i)+\theta_i P^{1-\sigma}}{1-\theta_i+\theta_i P^{1-\sigma}}\right)$; $i=L,S$ is the price elasticity of non-traded goods in terms of i workers. Thus $\sigma_c<\sigma$ if $(\gamma_T-\gamma_N)\left(\varepsilon_{N,P}^L-\varepsilon_{N,P}^S\right)<0$. Let me first investigate the sign of the gap in the price elasticity of non-traded goods between workers' groups (i.e. $\varepsilon_{N,P}^L-\varepsilon_{N,P}^S$). The computation reveals that $\varepsilon_{N,P}^L-\varepsilon_{N,P}^S>0\Leftrightarrow\sigma<1$. To satisfy $\sigma_c<\sigma$, now I need to find a condition so that $\gamma_T-\gamma_N<0$.

$$\gamma_T-\gamma_N<0\;\Leftrightarrow\;\begin{cases}a)\;\frac{C_T^L}{C_T}<\frac{C_N^L}{C_N}\\ b)\;\frac{C_T^S}{C_T}>\frac{C_N^S}{C_N}\end{cases}\;\Leftrightarrow\;\frac{C_N^S}{C_T^S}<\frac{C_N}{C_T}<\frac{C_N^L}{C_T^L}\;\Leftrightarrow\;\theta_S<\frac{1}{1+\Psi}<\theta_L. \qquad B.1$$

Where $\Psi\equiv\frac{P^{\sigma}C_N}{C_T}$. By applying Equation 5 and 6, the inequality a of Equation $B.1$ is rewritten as:

$$\frac{PC_N}{C_T}<\frac{\theta_L P^{1-\sigma}}{1-\theta_L}\Rightarrow\frac{C-C_T}{C_T}<\frac{\theta_L P^{1-\sigma}}{1-\theta_L}\Rightarrow\frac{C}{C_T}<1+\frac{\theta_L P^{1-\sigma}}{1-\theta_L}\Rightarrow\frac{C_T}{C}>\frac{1-\theta_L}{1-\theta_L+\theta_L P^{1-\sigma}}.$$

It is easy to know that $\frac{1-\theta_L}{1-\theta_L+\theta_L P^{1-\sigma}}=\frac{C_T^L}{C^L}$. It gives:

$$\frac{C_T}{C}>\frac{C_T^L}{C^L}\Rightarrow\frac{C^L}{C}>\frac{C_T^L}{C_T}\Rightarrow\frac{C-C^S}{C}>\frac{C_T-C_T^S}{C_T}\Rightarrow\frac{C^S}{C}<\frac{C_T^S}{C_T}\Rightarrow I<I_T,$$

where $I_T \equiv \frac{C_T^S}{C_T}$ denotes the consumption inequality on traded goods. In the same way for the inequality b of Equation $B.1$, we have $I_N < I$, where $I_N \equiv \frac{C_N^S}{C_N}$ is the consumption inequality on non-traded goods. To sum up, the final result derived from Equation $B.1$ is equal to $I_N < I < I_T$.

Appendix C Proof of Proposition 7

A steady-state combination of Equation 10a and 10b gives

$$(P^*)^{1-\sigma} = M^*(R)\ \Psi(P^*, I^*(R)). \qquad C.1$$

Where $M^*(R) \equiv \frac{\alpha}{\beta} \frac{S_T^* + \left(\frac{S_T^*}{L_T^*}\right)^{1-\beta} R}{S_N^*} \Rightarrow \frac{dM^*}{dR} > 0$. The steady-state response of the real exchange rate to increased R is then given by

$$-\left(\frac{1}{\Psi}\frac{d\Psi}{dP^*} + \frac{1-\sigma}{P^*}\right)\frac{dP^*}{dR} = \frac{1}{M^*}\frac{dM^*}{dR} + \frac{1}{\Psi}\frac{d\Psi}{dI^*}\frac{dI^*}{dR}. \qquad C.2$$

We can find easily that $\frac{1}{\Psi}\frac{d\Psi}{dP^*} = \frac{\sigma - \sigma_c}{P^*}$. Thus Equation $C.2$ is rewritten as following.

$$\frac{dP^*}{dR} = -\frac{1}{1-\sigma_c}\frac{P^*}{M^*}\frac{dM^*}{dR} - \frac{1}{1-\sigma_c}\frac{P^*}{\Psi}\frac{d\Psi}{dI^*}\frac{dI^*}{dR}. \qquad C.3$$

It indicates when $\theta_S = \theta_L,\ \sigma = \sigma_c \Rightarrow \frac{1}{\Psi}\frac{d\Psi}{dI^*} = 0$. So that the steady-state real exchange rate response to an increase in R equals $\left(\frac{dP^*}{dR}\right)_{\theta_S=\theta_L} = -\frac{1}{1-\sigma}\frac{P^*}{M^*}\frac{dM^*}{dR} < 0$. Now regarding Proposition 4, $\theta_S < \theta_L \Rightarrow \frac{1}{\Psi}\frac{d\Psi}{dI^*} < 0$ (see. Appendix A). Hence the size of $\frac{dP^*}{dR}$ depends on the steady-state income inequality response to increased R (i.e. $\frac{dI^*}{dR}$).

Appendix D Countries included in the sample database

Table 11: List of Countries

	Period	No. Period		Period	No. Period
Albania	1995-2014	4	Italy	1990-2009	4
Algeria	1975-1999	5	Jamaica	1995-2009	3
Argentina*	1975-2014	8	Japan*	1980-2009	6
Australia*	1975-2014	8	Kazakhstan*	1995-2014	4
Austria	1975-2009	7	Korea south	1975-2009	7
Azerbaijan	1995-2009	3	Kyrgyzstan*	1995-2014	4
Bangladesh	1975-1999	5	Latvia*	1995-2009	3
Belgium	1995-2009	3	Malaysia*	1975-2009	7
Bolivia*	1980-2009	6	Mexico*	1975-2014	8
Botswana	1980-2009	6	Moldova	1995-2009	3
Brazil*	1975-2014	8	Mongolia	1990-2009	4
Bulgaria*	2000-2014	3	Morocco*	1980-2009	6
Cameron*	1975-2009	7	Nepal	1985-2004	4
Canada*	1975-2014	8	Netherlands*	1975-2009	7
Chile*	1975-2014	8	New Zealand*	1980-2014	7
China*	1975-2014	8	Norway*	1975-2014	8
Colombia*	1975-2014	8	Pakistan	1975-2014	8
Costa Rica	1985-2014	6	Paraguay	1995-2014	4
Cote d′Ivoire	1975-1999	5	Peru*	1990-2014	5
Croatia	1995-2009	3	Philippine*	1975-2014	8
Cyprus	1980-2009	6	Poland*	1995-2009	3
Czech*	1995-2009	3	Portugal	1995-2009	3
Denmark*	1975-2009	7	Romania	1995-2014	4
Dominican Rep.	1980-2014	7	Russia*	2000-2014	3
Ecuador*	1975-2014	8	South Africa	1975-2014	8
Egypt*	1975-2009	7	Spain*	1995-2009	3
El Salvador	1975-1995	5	Sri Lanka	1980-2009	6
Eritrea	1995-2009	3	Sweden*	1980-2004	5
Ethiopia	1990-2009	4	Tanzania	1995-2009	3
Finland*	1975-2009	7	Thailand	1975-2014	8
France*	1975-2009	7	Trinidad and Tobago	1985-2004	4
Ghana	1975-1999	5	Turkey	1975-2009	7
Greece*	1995-2009	3	Ukraine*	1995-2014	4
Honduras*	1975-2014	8	United Kingdom*	1990-2014	5
Hungary*	1995-2009	3	Uruguay	1985-2009	5
India*	1975-2009	7	Venezuela*	1975-2014	8
Indonesia*	1990-2014	5	Yemen*	1995-2009	3
Iran*	1975-2014	8	Zimbabwe	1975-1999	5
Ireland	1995-2009	3			

Note: countries marked by (∗) are included in the sample of the commodity price index.

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