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Do Land Market Restrictions Hinder Structural Change in a Rural Economy?

Evidence from Sri Lanka

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Abstract

This paper analyzes the effects of land market restrictions on structural change from agriculture to non-farm in a rural economy. This paper develops a theoretical model that focuses on higher migration costs due to restrictions on alienability, and identifies the possibility of a reverse structural change where the share of nonagricultural employment declines. The reverse structural change can occur under plausible conditions: if demand for the non-agricultural good is income-inelastic (assuming the non-farm good is non-tradable), or non-agriculture is less labor intensive relative to agriculture (assuming the non-farm good is tradable). For identification, this paper exploits a natural experiment in Sri Lanka where historical malaria played a unique role in land policy. The empirical evidence indicates significant adverse effects of land restrictions on manufacturing and services employment, rural wages, and per capita household consumption. The evidence on the disaggregated occupational choices suggests that land restrictions increase wage employment in agriculture, but reduce it in manufacturing and services, with no perceptible effects on self-employment in non-agriculture. The results are consistent with the migration costs model, but contradict two widely discussed alternative mechanisms: collateral effect and property rights insecurity. This paper also provides direct evidence in favor of the migration costs mechanism.

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Do Land Market Restrictions Hinder Structural Change in a Rural Economy? Evidence from Sri Lanka[1]

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1 Introduction

Economic transformation from a predominantly agrarian and rural economy to an industrialized and urban one is considered a cornerstone of economic development (Matsuyama (2008), Syrquin (1988), Chenery (1960), Kuznets (1966), Clark (1957)). The classic literature on structural change focused on the long-run evolution of sectoral shares of agriculture, manufacturing and services in employment and output.[1] Recent studies on structural change in developing countries provide robust evidence that labor productivity in non-agriculture is significantly higher than in agriculture. Macmillan and Rodrik (2011) report that the productivity ratio between manufacturing and agriculture is on average 2.3 in Africa, 2.8 in Latin America and 3.9 in Asia. Such large productivity gaps imply that a structural change in favor of non-farm activities (manufacturing and services) may be desirable for both efficiency and poverty alleviation (Macmillan and Rodrik, 2011; Gollin, Lagakos and Waugh, 2014; Vollrath, 2009).[2]

It is widely argued that misallocation of labor – reflected in persistent differences in labor productivity across areas and sectors – results from dualism in the labor market. The source of this dualism can be traced to formal and informal barriers/restrictions in the labor and land markets (for an excellent survey see Restuccia (2013)).[3] In many developing countries, especially in the 1950s and 1960s, dualistic policies were designed to ensure supply of labor at low wages to the urban manufacturing sector, and to relax the "savings constraint" and "wage goods constraint"; they include the so-called price-scissors (see Sah and Stiglitz (1987)), and geographic mobility restrictions such as Hukou in China and Ho Khau in Vietnam. Restrictions on sales of land were a salient feature of the policy regimes in many countries in Asia and Africa. This paper provides an analysis of the extent to which formal sales and rental restrictions on agricultural land can constrain employment transition from farm to non-farm activities in rural areas, and create policy-induced dualism within the rural economy. We provide evidence on the effects of land restrictions on disaggregated employment choices across five occupational categories. We also analyze whether land market restrictions affect per capita consumption in a village adversely.

[1] For recent surveys, please see Matsuyama (2008), Herrendorf et al. (2013).

[2] Gollin et al. (2014) show that the productivity gap between agriculture and nonagriculture cannot be explained by data quality and measurement problems.

[3] A partial list of contributions in this literature includes Adamopoulos and Restuccia (forthcoming), Hayami and Prescott (2008), Restuccia et al. (2008), Lewis (1954); Harris and Todaro (1970); Fei and Ranis (1964). In an interesting recent paper Gottlieb and Grobovsek (2015) find that policy distortions in land markets in Ethiopia may not have large effects on productivity because of strong endogenous price responses.

This paper is related to a growing literature in development economics that examines the economic effects of land market imperfections and restrictions.[4] The focus of most of the studies has been on two mechanisms through which the effects of land restrictions are mediated: insecurity of property rights (Demsez (1967)), and collateral constraint due to inalienability of land (Feder and Onchan (1987), Feder et al. (1988), de Soto (2000)). The evidence shows that insecure and nontransferable property rights in land adversely affects incentives to undertake irreversible long-term investment and productivity (Feder and Onchan (1987), Besely (1995), Iyer and Do (2008), Holden et al. (2007), Deininger and Jin (2006), Jacoby et al. (2002)).[5] A number of studies find that improvements in the security of property rights in land results in increased labor supply to market, as guard labor to protect contestable land rights is not needed anymore (see Field (2007) on Peru, Iyer and DO (2008) on Vietnam, Iyer et al. (2014) on China).[6] Feder et al. (1988), Feder and Feeny (1991) and de Soto(2000) argue that formalization of property rights in land can potentially create collateral value and improve access to credit, though empirical evidence on this channel is mixed, at best.[7] Hayashi and Prescott (2008), de Janvry et al. (forthcoming), and Emran and Shilpi (forthcoming, 2015) focus on the migration costs channel; they underscore that sales and rental restrictions can significantly increase the costs of migration as a household loses its claim to future income stream from land when it leaves the village.[8] Hayashi and Prescott (2008) show that in prewar Japan informal inheritance custom resulted in increased rural-urban migration costs, and the resulting misallocation of labor was an important factor in low productivity. Emran and Shilpi (forthcoming, 2015) find that as a result of the increased migration costs, land market restrictions increase female labor force

[4]For an excellent early survey, see Deininger and Feder (1998).

[5]Besley (1995) finds that, in Ghana, better land rights encourage investments in tree plantation; Jacoby et al. (2002) report that farmers invest more in organic fertilizers when the threat of expropriation (land reallocation) is lower in rural China; Iyer and Do (2008) Show that 1993 land law in Vietnam that made private land rights secure and transferable led to more investments in perennial industrial and fruit crops. Holden et al. (2007) find that, in Ethiopia, low cost land certification increases investment and productivity. Deininger and Jin (2006) report that transfer rights significantly enhance investment in Ethiopia.

[6]Iyer, Meng, Qian and Zhao (2014) find that transfer of property rights to private households and subsequent removal of sales restrictions on privately-owned houses in urban areas in China during the 1980s and 1990s increased supply of labor to private sector jobs. Field (2007) finds that land titling in Peruvian urban slums resulted in substantial increase in labor hours, a shift from work at home to work outside, and from child to adult labor.

[7]Besley (1995) does not find any evidence in Ghana to support the view that land rights enhance collateral value and access to credit, and Field (2007) fails to find any significant impact of land titling in Peruvian slums on access to or use of credit. Pender and Kerr (1999) report similar findings for south India, and Iyer and Do (2008) for Vietnam. In a series of papers on Thailand in late eighties and early nineties, Gershon Feder and his co-authors find that land titling increases access to credit and investment when formal banks are important source of credit. For a summary of the evidence and a critique of the de Soto view, see Haldar and Stiglitz (2013).

[8]Besley (1995) discusses "gains from trade" as a third mechanism for the effects of land rights, where better rights increase the land price. Although he does not discuss rural-urban migration, one can interpret part of the 'gains from trade' as 'gains from migration'.

participation and depress both male and female wages in the rural areas of Sri Lanka.

We analyze the effects of land market restrictions on structural change from farm to non-farm sector in rural areas as measured by employment shares.[9] We construct a two-sector model where agricultural production requires land and labor, and non-agricultural production uses labor and capital. The inalienability of land rights because of sales restrictions implies that, in the event of migration, households lose the net present value of future earnings from land. The increased migration costs lower migration, and thus lead to higher labor endowment in a location with higher incidence of land restrictions. However, the effects of land restrictions on the allocation of labor between agriculture and non-agriculture in equilibrium depends on the tradability of non-agricultural good.[10] When non-agriculture is traded beyond the village, the share of agricultural labor increases in response to imposition of land restrictions under the plausible assumption that agriculture is more labor intensive. With nontradable non-agricultural product (i.e, traded only in local market), the local demand plays an important role (Matsuyama (2008)), and a positive impact of land restrictions on agriculture's labor share is found if demand for non-agriculture is income inelastic. The assumption that demand for rural non-farm goods and services is inelastic seems plausible, because most of the goods and services produced in rural non-farm sector are of low quality, and a household substitutes away from them as income increases.[11] The sales restrictions in land markets may thus result in a form of "reverse structural change" in the rural economy where the labor share in agriculture increases.

It is important to appreciate that the prediction of a "reverse structural change" is not unique to the migration costs mechanism; it also holds true when the primary mediating channel for the effects of restrictions is instead collateral value of land. Restrictions on sales and alienability of land can destroy its collateral value, and restrict a household's access to capital, thus adversely affecting the non-farm entrepreneurship and expansion of non-farm businesses. An important advantage of our theoretical framework is its ability

[9]To the best of our knowledge, there is no analysis of the effects of policy restrictions in the land market on structural change in a rural economy that tries to disentangle three alternative causal mechanisms discussed above. In the context of Vietnam's land titling program, Iyer and Do (2008) find that households in provinces with more private land titles devote more time to non-farm activities, but they do not discuss whether this leads to structural change. Deininger et al. (2014) find that, in China, higher proportion of land certificates at the village level increases labor allocation to nonfarm activities. The focus in this paper is on the occupational choices (rather than time allocation); we estimate how the primary occupation of an individual changes in response to land market restrictions.

[10]We assume that agricultural good is internationally tradable, which seems appropriate given that rice is the most important crop in Sri Lanka. This implies that the endogeneous terms of trade responses highlighted by Gottlieb and Grobovsek (2015) may not be relevant for our application.

[11]Unfortunately, we are not aware of any credible estimates of income elasticity of demand for non-agricultural goods produced in the village. Part of the challenge is that most of the surveys do not identify the goods purchased by a household in terms of their location of production.

to distinguish between migration and other channels. If migration cost is the main mechanism driving the intersectoral reallocation of labor, then our model predicts lower wages in areas with higher incidence of land restrictions. In contrast, wage rate will remain largely unaffected by land restrictions if either credit constraint or insecurity of property requiring guard labor is important (assuming no change in migration costs, and no policy restrictions on rural-urban labor mobility).

To provide evidence on the effects of land market restrictions on occupational choices, we follow Emran and Shilpi (forthcoming, 2015) and exploit a historical quasi-experiment in Sri Lanka where the cross-section variations in the incidence of land restrictions across different sub-districts (i.e., proportion of land under policy restrictions) were primarily determined by historical malaria prevalence through its effects on 'crown land'.[12] Historical malaria caused exodus from the affected areas, and the land abandoned by the households was nationalized by the colonial government and designated as crown land. The crown land was later distributed through settlements after independence, and restrictions on sales, mortgage and rental were imposed by the government.[13] This created a positive correlation between historical malaria incidence and the incidence of land restrictions in an area through the availability of crown land. We use this positive correlation between historical malaria and incidence of land restrictions in a sub-district to identify the causal effects of land restrictions on structural change in the rural economy. As noted in Emran and Shilpi (forthcoming), the strength of the identification strategy derives partly from the following observations: (i) a successful nationwide malaria eradication program was implemented in Sri Lanka in 1947; malaria endemicity fell close to zero by 1950-51.[14] Historical malaria estimate from 1937-41 is used in this paper to identify the effects of land market restrictions in 2002 (survey year for our data), (ii) the timing of the malaria eradication program was determined by technological breakthrough (DDT), and thus can be treated as plausibly exogenous,[15] and (iii) most of the current population in a subdistrict devastated by high historical malaria were never exposed to historical malaria there, as they were resettled

[12]We use the data on malaria endemicity at the district level for 1937-41 compiled by Newman (1965). Emran and Shilpi (forthcoming, 2015) exploit historical malaria incidence to analyze the effects of land restrictions on male and female wages, and women's labor force participation.

[13]Over time, restrictions on mortgage have been relaxed when dealing with public banks.

[14]Malaria endemicity is measured by enlarged spleen rates. Reported malaria cases in Sri Lanka were reduced from about 3 million per year during pre-eradication era to only 29 in 1964 (Harrison, 1978). The number of malaria death cases were 30 in 2002 among a population of 21 million. The reported malaria death were 4 in 2003, and 0 in 2005.

[15]Although DDT was first synthesized in 1874, its insecticidal properties were discovered in 1939 by Swiss scientist Paul H Muller. It was widely used during second World War to control malaria and typhus, and after the war DDT was made available as an agricultural pesticide and for malaria eradication programs.

from other relatively malaria free regions;[16] (iv) unlike some countries such as China (Hukou), Vietnam (Ho Khau), and Ethiopia, there are no geographic mobility restrictions or other policies that tie households to a village in Sri Lanka. Sri Lanka is thus an excellent case study to understand the effects of sales and rental restrictions in the rural land market, as the effects of land market restrictions are not confounded by the effects of other related policies.[17] We discuss a number of potential objections to the identification strategy and provide evidence on its credibility in section 3 below.

The empirical results show that land restrictions reduce the probability of participation in non-agriculture significantly: a one percentage point increase in land restrictions reduces probability of participation in non-agriculture by 1.38 percent.[18] Land market restrictions thus hold back structural change from agriculture to non-agriculture in a rural economy. The more disaggregated analysis using a multinomial logit model indicates that land restrictions have negative and statistically significant effects on the probability of wage employment in manufacturing and services activities, with a particularly large effect on services employment.[19] In contrast, there is a positive effect on the probability of being employed in agricultural wage work. The probability of self-employment in the non-agricultural activity, however, does not change significantly with an increase in land under restrictions suggesting an absence of a collateral and credit effect. The instrumental variable regression for wages confirms that land restrictions affect local wage adversely, and lower per capita household consumption in a village. Earlier results in the context of Sri Lanka also show that land restrictions affect wages negatively, for both men and women (Emran and Shilpi (forthcoming, 2015)). The evidence thus rejects the collateral or insecure property rights channels, and highlights the importance of migration costs, as emphasized by Hayashi and Prescott (2008).

The rest of the paper is organized as follows. Section 2 provides the main conceptual framework for the empirical estimation. Section 3 discusses the empirical strategy adopted in the paper for identification

[16] This observation extremely important as it implies that health, education, attitude etc. of the current generation living in LDO lands (i.e., surveyed in 2002) cannot be affected by historical malaria in any significant way. For a complete discussion including direct evidence, please see section (3) below.

[17] For example, the Hukou system in China controlled rural urban migration with a web of restrictions that span employment, education, and marriage. Moving to urban areas without an urban Hukou would mean that a family would lose grain rations, the land, and the children will not be able to attend schools in urban areas. Although over the years, Hukou restrictions have been relaxed in some provinces, it is difficult to treat these changes as exogeneous, as they are partly determined by the economic changes in a province. A significant part of the existing literature focuses on China, Vietnam, and Ethiopia.

[18] The nonagriculture comprises of three categories of employment: self employment in non-farm, wage employment in manufacturing, and wage employment in services, and the agriculture includes both agricultural self-employment and wage employment.

[19] In multinomial logit model, we treat self-employment in agriculture as the omitted category. We use a control function approach to instrumental variables estimation for multinomial logit.

and estimation of the effects of land market restrictions on wages. Section 4 provides a description of the data. Empirical results are presented in section 5. The paper ends with a concise summary of the results in conclusion.

2 Conceptual framework

We set-up a simple model to understand the possible effects of land restrictions on employment pattern in a rural economy. To develop the intuitions regarding the impacts of land restrictions, we consider a two-sector economy producing food (f) and non-agriculture or non-farm good (n). We assume that food is internationally traded and its price is normalized to 1. The results depend on whether the non-farm good is tradable or nontradable (beyond the village), which is consistent with a substantial literature on the role of domestic demand in structural change when preference is non-homothetic (see the survey by Matsuyama (2008)). For simplicity, land is assumed to be used only in food production and capital in non-agricultural production. Both activities require labor in the production process. Denoting land as T and capital as K , the production functions for different activities are:

$$
\begin{aligned}
Q^f &= (T)^\alpha \left(L^f\right)^{1-\alpha} \text{ for agriculture} \\
Q^n_k &= (K)^\beta \left(L^n\right)^{1-\beta}) \text{ for nonagricultural goods and services}
\end{aligned}
$$

The utility function for a consumer is assumed to be of Stone-Geary form:

$$
U_k = U_k(C^f, C^n) = \tau_f \ln(C^f - \gamma^f) + \tau_n \ln(C^n - \gamma^n) \tag{1}
$$

where C^f and C^n are consumption of food and non-agricultural good respectively, and γ^f and γ^n are subsistence quantities of food and nonagricultural good respectively. This specification allows for Engel curve effect, with the homothetic utility function as a special case where $\gamma^f = \gamma^n = 0$. The budget constraint can be written as:

$$
C^f + P^n C^n = Y
$$

where P^n is the relative price of non-agricultural good and Y is the total income.

The demand functions for food and non-agriculture can be derived as:

$$C^j = \gamma^j + \frac{\tau_j}{P^j}[Y - \gamma^f - \gamma^n P^n] \text{ where } j = f, n, \text{ and } P^f = 1 \tag{2}$$

Households differ in terms of land; household h with $T_h \geq 0$ amount of land. Access to credit (capital) required for non-farm activities depend on the amount of land a household can offer as collateral. To keep the model simple, we assume that a household h can get $K_h = \lambda T_h$ amount of capital. Denote aggregate land and capital in the village by $\bar{T}$ and $\bar{K}$ respectively. To focus on the collateral role of land and credit constraint in the non-farm sector, we adopt the simplifying assumption that λ is low enough so that the household with the most land is credit constrained at the initial equilibrium without land restrictions[20] In this model, a collateral effect of land restrictions a la de Soto (1989) is represented as a lower value of λ. Labor is assumed to be mobile across locations and migration across locations is costly. We assume that households face heterogeneous migration costs $\varphi(d_h)$ in the absence of land restrictions, where d_h is the distance of household h from the urban labor market, and can be interpreted as a summary measure of all types of migration costs related to the geographic location of a household. Under land restrictions (transport costs, social costs, amenities etc.), a household with LDO land loses the rent on land when it decides to migrate out of the village. The migration costs are thus higher when a household is in LDO settlement, and has higher amount of land, and also lives in a village further from the urban center.

The important implication of higher migration costs due to land restrictions is that many households find it optimal not to migrate because of the loss of land, and the effective labor endowment in a village increases as a result. Under the assumption that land restrictions have no effect on credit availability (i.e., λ remains same after the imposition of restrictions), land restrictions lower equilibrium wage as more households stay back in the village and supply of labor increases. The equilibrium wage is determined by local labor market clearing coupled with a rural-urban migration condition (spatial arbitrage condition in the labor market). For a complete discussion on the wage determination and the effects of restrictions, please see Appendix

[20]This is similar to the assumption of Feder and Feeny (1991). We, however, note that the central results in this paper remain valid in a more general model where some households have enough land so that they are not credit constrained for their investment in the non-farm sector. For an alternative model where the collateral value affects the interest rate, please see Besley (1995).

A.[21] However, the intuition behind the result that land restrictions reduce equilibrium wage in local labor market can be seen readily from the migration equilibrium condition. Denote the indirect utility function is rural and urban areas as $V^r(.)$ and $V^u(.)$, we can write the migration equilibrium condition at the initial equilibrium (without restrictions):

$$V^u\left(T_{h^0}, w^u, 1, P^n\right) - \varphi\left(d_{h^0}\right) = V^r\left(T_{h^0}, w^0, 1, P^n\right) \tag{3}$$

So at the initial equilibrium all households located at a distance higher than a threshold $d_h > d_{h^0}$ decide not to migrate. The important point in equation (3) for our analysis is that the land endowment a household has as a rural resident appears also in the urban indirect utility function as they can sell the land and take the money with them. The land restrictions imply that the indirect utility for a migrant household becomes: $V^u\left(0, w^u, 1, P^n\right)$ where the land is set equal to zero because they lose the land. Thus we have $V^u\left(0, w^u, 1, P^n\right) < V^u\left(T_h, w^u, 1, P^n\right)$. This implies from equation (3) that all households at location $d_{h^0}$with positive land endowment will find it undesirable to migrate because they lose the land. The labor supply in the village market increases as a result and the equilibrium wage declines.

Case 1: Nonagricultural Sector Produces Tradable Goods

In this section, we consider the case when the non-farm good is also tradable along with the agricultural good, and thus the prices of both consumer goods can be normalized to 1. In this model, local demand (village income) does not play any role in the effects of land restrictions on intersectoral labor allocation. Denote the initial equilibrium wage rate (without restrictions) in the local labor market by w^0, and the wage rate following imposition of land restrictions as $w^* < w^0$.

Given a wage rate w, the relative share of labor in non-farm sector in the village can be expressed as:

$$S_{nf} = \frac{L^{n*}}{L^{f*}} = \left[\frac{(1-\beta)^{1/\beta}\bar{K}}{(1-\alpha)^{1/\alpha}\bar{T}}\right] w^{(\beta-\alpha)/\alpha\beta} \tag{4}$$

[21] In different models, Emran and Shilpi (2015, forthcoming) also show that land market restrictions lead to lower wages under plausible conditions when migration cost is the relevant channel.

Proposition 1: Under the assumptions that both food and non-food are traded goods, the imposition of land restrictions increases local labor supply and decreases local wage due to higher migration costs. The allocation of labor to food production increases relative to that to non-food production if labor's share in food production is higher than that in non-food production [(*1* $- \alpha$) > (*1* $- \beta$)].

The proof of the second part of proposition 1 follows directly from equation (4), by noting that $\frac{\partial S_{nf}}{\partial w} > 0$ if $\beta > \alpha$, and from the fact that land restrictions reduce the equilibrium wage in the village (see appendix A for the effects on equilibrium wage).[22] As labor share in agriculture is found to be much larger (more than 70%) compared to non-agricultural production (below 30%) in developing countries, the imposition of land restrictions is thus likely to lead to an increase in labor share in food/agricultural production, thus holding back the structural change.

Equation (4) can also be used to derive the implications of collateral constraints due to land restrictions. If restrictions on land sales reduce (or destroy) the collateral value resulting in a lower λ, the aggregate capital in the non-farm sector declines ($\bar{K}$ in equation (4) is lower). It follows from equation (4) that the ratio of employment in non-farm to agriculture decreases as a result even if there is no change in wages. Thus if one finds that land restrictions reduce employment in the no-farm sector, it does not shed light on whether the effects are due to increased migration costs or due to tightening credit constraints. However, it is possible to identify the channel which is primarily responsible for the observed impacts of land restrictions if one considers both employment pattern and wages. A decrease in capital due to land restrictions reduces the demand for labor in non-food production, resulting in a decline in the local wage. Such a wage decline will induce out-migration reducing the labor supply to the village labor market, and thus restoring the rural-urban labor market arbitrage condition. In the new equilibrium, the local wage rate will remain unaffected if migration costs are not affected by the imposition of land restrictions.

Note that insecurity of property rights will induce households to keep someone on the premise to guard against theft and expropriation, reducing the labor supply and increasing the local wage. Again if labor is mobile, then this will induce more in-migration and push the wage down to initial equilibrium wage. In contrast to both of these cases, an increase in migration costs (and no change in borrowing constraint) will

[22]That higher land restrictions result in lower wages in a village was established, both theoretically and empirically, in Emran and Shilpi (forthcoming). The theoretical model in Emran and Shilpi (forthcoming) is, however, substantially different from the model developed in this paper.

increase labor supply and reduce local wage. The empirical analysis in this paper relies on this insight to distinguish between migration, property right insecurity and borrowing channels.

Case 2: Nonfarm good is a Non-tradable

The theoretical model discussed above takes non-agricultural output as a tradable good, and as a result, the village income (local demand) does not play any role in the structural change. Most non-farm activities including some manufacturing firms in rural areas, however, cater to local demand, and their products and services are non-tradable beyond the confines of the village market. As noted before, when the nonagricultural good is a nontradable, the effects of higher migration costs on intersectoral labor allocation depends on the underlying assumptions about the income elasticity of demand. With Stone-Geary utility function, income elasticity of demand can range from inelastic to elastic by varying the subsistence quantities. For instance, if $\gamma^f = 0$ and $\gamma^n > 0$, then income elasticity of nonagricultural good is less than unity. If $\gamma^f = 0$ and $\gamma^n = 0$, then utility function represents homothetic preference and income elasticities of both goods are equal to unity. In contrast, $\gamma^f > 0$ and $\gamma^n = 0$ implies an elastic income response for nonagricultural goods. The equilibrium is now characterized by market clearing in nonagricultural goods market (n market), in addition to market clearing in the village market and the rural-urban migration equilibrium conditions (for details, please see appendix A).

We first consider the case where income elasticity of non-farm goods demand is greater than 1, and set $\gamma^f > 0$ and $\gamma^n = 0$ in the demand functions in equation (1). The share of non-farm employment in the village can be expressed as :

$$S_{nf} = \frac{\tau_n(1-\beta)}{(1-\tau_n)(1-\alpha)}\left[1 - \frac{\gamma^f w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}}\right] \tag{5}$$

Equation (5) can be used to examine the impact of land restrictions on the share of nonagricultural employment in the village. From equation (5), it is seen easily that $\frac{\partial S_{nf}}{\partial w} < 0$. Since land restrictions lower the equilibrium wage (equation (3) above), we thus expect that the share of non-farm employment will increase following imposition of restrictions in this case. Note also that if $\gamma^f = 0$, the utility function becomes homothetic (Cobb-Douglas in this case). In this homothetic case, it follows from equation (5) that

$\frac{\partial S_{nf}}{\partial w} = 0$. In other words, with homothetic preference, the relative employment shares of nonagriculture and agriculture remain unaffected by land restrictions.

When the demand for non-agricultural good is income inelastic, we set $.\gamma^f = 0$ and $\gamma^n > 0$ in the demand equations (1) above. In this case, it is not possible to derive a closed form solution for the share of non-farm employment in a village as a function of the equilibrium wage. As we show in appendix A, the response of the share of non-farm employment in the village can be expressed as below:

$$\frac{\partial S_{nf}}{\partial w} = \frac{(1-\beta)\, z_2}{\frac{(1-\alpha)}{(1-\beta)\beta} s_{nf} - z_1} > 0 \tag{6}$$

where $z_1 = \frac{P^n \gamma^n w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha} \bar{T}}$ and $z_2 = \dfrac{z_1 S_{nf}}{\alpha \beta w}$. It is easy to check that the denominator in the right hand side of equation (6) is positive (see appendix A). As we discussed before, land restrictions lower the equilibrium wage in the village, which implies from equation (6) that land restrictions result in a contraction of the non-farm share of employment. We collect the above results in proposition (2) below. For a more complete treatment of the results, please see appendix A.

Proposition 2: Under the assumption that agricultural good is tradable but nonagricultural good is non-tradable, the effects of imposition of land restrictions on labor allocation depends on the income elasticity of demand for nonagricultural good. The imposition of land restrictions leads to a decline (an increase) in the share of nonagriculture in total employment if the income elasticity of demand for non-farm good is smaller (greater) than 1.

3 Estimating Equations and Empirical Strategy

To test the hypotheses regarding the effect of land restrictions on employment pattern, we utilize a simple employment choice model at the individual level. Suppose an individual i can choose from a set of J different employment options ($j = 1, 2...J$) at location k. Individual i will choose activity j if:

$$U_{ijk} \geqq U_{imk} \text{ for all } j \neq m$$

where u_{ijk} is the utility derived from being engaged in activity j at location k. We further specify the utility function as:

$$U_{ijk} = U(y_{jk}, Z_{ijk}, \varepsilon_{ijk})$$

where y_{jk} is the wage/income in activity j and location k, Z_{ijk} is a vector of observed individual characteristics, and ε_{ijk} is the individual specific error term. Following McFadden(1981), we assume that utility u_{ijk} can be approximated by a linear specification:

$$U_{ijk} = \alpha_0 + \alpha_1 y_{jk} + \alpha_2 Z_{ijk} + \varepsilon_{ijk} \tag{7}$$

Utilizing the migration equilibrium condition in equation (3) above, wage/income can be specified as:

$$y_{jk} = \beta_0 + \beta_1 \theta_k + \beta_2 d_{ku} + Z_{ik}^{'} \Gamma + \nu_{ik} \tag{8}$$

Where ν_{ik} is the error term..

Substituting income from equation (8), we have:

$$U_{ijk} = \gamma_0 + \gamma_1 \theta_k + \gamma_2 d_{kh} + \gamma_5 Z_{ijk} + \eta_{ijk} \tag{9}$$

Under the assumption that η_{ijk} follows an iid logit distribution, equation (9) can be estimated using the standard multinomial logit formulation.

The main challenge in the estimation of equation (9) is that areas may differ systematically in observed and unobserved dimensions, and when the unobserved characteristics are correlated with the incidence of land restrictions, it may lead to significant omitted variables bias. The most important sources of omitted variables bias are unobserved labor and land productivity heterogeneity as they are first order determinants of labor demand and wages.

5.1 Sources of Bias

Many governments impose restrictions on sales and rental of land in settlement areas (as is the case in Sri Lanka), and settlement usually takes place in the low quality marginal land.[23] When we observe

[23]However, in many countries such as China, Vietnam, and Ethiopia the land policy is national in scope, and there is little if any variation in the restrictions imposed across regions.

land under private property rights to coexist with land under government restrictions, the land under restrictions is thus likely to be of lower quality.[24]

A second important issue is human capital heterogeneity. Since lands under policy restrictions in Sri Lanka are mainly settlement lands, one might worry that the people who had settled there are of lower productivity due to adverse human capital characteristics. Such negative labor productivity can give rise to a negative correlation between incidence of land restrictions and wages. Fortunately, Sri Lanka seems a non-standard case: all of the available evidence shows that land and labor productivity is *higher* in areas under land policy restrictions in Sri Lanka, as we discuss below.

To provide evidence on productivity differences, we analyze crop yield which is an excellent summary statistic for land and labor productivity of an area. Table 1 reports regressions of crop yields on proportion of land under restrictions while controlling for rainfall, slope and a dummy indicating whether a DSD is located within 5 kilometer (km) of a river. The yields are higher in land under restrictions for a number of crops including rice, the main crop in Sri Lanka (please see Table 1 for details).[25] Moreover, there is no evidence of unfavorable health conditions in areas with high land restrictions; if anything, the evidence indicates advantages in favor of the LDO areas in terms of health status of the population. The correlations of incidence of land restrictions with the incidence of chronic illness and disability and percentage of non-pregnant women suffering from different degrees of anemia are statistically insignificant (please see Table 2).[26] The observed parity in land and labor productivity in areas under land restrictions is the outcome of Sri Lanka government's heavy investment in irrigation development in resettled areas, and in health, education.

An additional source of bias in the OLS estimates is possible measurement error in the land restrictions variable. Both land and labor productivity advantages of land under restrictions and measurement errors

[24]A related issue is that land restrictions may affect incentives to invest in land improvements including irrigation, and if the high land restrictions areas are also areas with low irrigation, a negative effect of land restrictions may primarily reflect a lack of irrigation, rather than higher mobility costs. Note, however, that such differences in land investments is part of the causal effects of land restrictions, although the channel is different from the mobility costs arising from land restrictions emphasized in this paper. We provide evidence below that the concern that the high land restrictions areas suffer from low irrigation investment is unfounded.

[25]There is similar positive correlation between 'potential' yields of most major crops and proportion of land under restrictions, where potential yields are determined by soil quality, soil moisture, rainfall, temperature and other climatic variables assuming an ideal amount of labor is applied. The results from these regressions are omitted for the sake of brevity and are available from the authors.

[26]Only in the case of mild or moderate and severe anemia, the correlation is statistically significant at 10 percent level, but the sign is negative implying better health in subdistricts with more land under LDO restrictions.

imply that we would expect the OLS estimates of the effects of land restrictions on equilibrium wages to be biased toward zero. Thus simple OLS regressions may not be able to detect the effects of land restrictions (Emran and Shilpi (2015)).

5.2 Historical Malaria as a Natural Experiment

Malaria infestation played a unique role in the history of land policy in Sri Lanka. The spatial variations in historical malaria thus offers us an exogenous source of variations in the incidence of land restrictions. The areas affected by historical malaria endemicity from 13th century to early twentieth century witnessed exodus of population and large scale abandonment of land (De Silva, 1981). The abandoned land was taken over by the colonial government and designated as 'crown land'. The crown land was later distributed after the independence in 1948 under Land Development Ordinance of 1935, and restrictions on sales, mortgage, and rental were imposed (henceforth called LDO restrictions). Since the amount of crown land available in an area was historically determined by the intensity of malaria, the historical malaria incidence created exogenous variation in the incidence of land restrictions; the proportion of land under restrictions is higher in an area, the higher was the intensity of historical malaria prevalence during 13th-18th century.

In the regressions, we control for a number of area characteristics that can potentially be correlated with both malaria incidence and land productivity, resulting in omitted variables bias. For example, rainfall is among the most reliable predictors of incidence of malaria in Sri Lanka (Briet et al., 2008). Rainfall also has direct effect on land productivity.[27] In addition to rainfall, regressions control for slope (steeper slope means less standing water and less malaria), a dummy indicating whether the DSD is within 5 km of a river, and current incidence of malaria. We discuss additional evidence that provides strong support to the conclusion that the identification scheme developed based on historical malaria incidence is not compromised by land and labor productivity differences.

5.3 Potential Objections to Identification Strategy

As noted in the introduction, there are a number possible objections to our identification scheme that exploits the positive correlation between historical malaria incidence and land restrictions across different sub-districts in Sri Lanka.

Perhaps the most immediate objection that comes to the mind of a reader familiar with the large

[27]This is true even with well-developed irrigation, because the marginal cost of rain water is zero, while the marginal cost of irrigation water is positive and likely upward sloping.

literature on colonial institutions across countries following Acemoglu et al. (2001) is that historical malaria might have affected the quality of property rights institutions through its influence on settler mortality. However, the long-term effects of malaria on the quality of institutions emphasized in the literature are not relevant for our identification scheme as identification in our case comes from *variations across districts*. The relevant institutions such as legal system, labor market regulations, and enforcement of contracts and property rights, however, are determined at the national level, and Acemoglu et al. (2001) results also relevant for cross-country differences in institutional quality.[28] As a conservative strategy, we control for the proportion of Sinhalese population in a sub-district as a measure of ethno-linguistic fractionalization that can potentially affect public goods provision.[29] A related worry is that historical malaria might have affected human capital of current labor force adversely in our sample. There are good reasons to believe that this concern is unfounded. First, and probably the most important, is the fact that the settlement schemes brought in people from relatively malaria free regions to the subdistricts which were abandoned because of historical malaria starting from 13th century. As a result, the vast majority of the current population (or their parents) were never exposed to historical malaria in the sub-district of their current residence (i.e, residence in 2002). Second, we exclude the cohorts that were potentially exposed (in utero or post-natal) to historical malaria in Sri Lanka.[30] Thus our sample is not contaminated by the possibility that someone might have been exposed to historical malaria before his/her mother resettled in a subdistrict of high historical malaria incidence.[31]

The important implication of the above discussion is that historical malaria in a sub-district cannot be correlated with the health outcomes of most of the current population. Indeed evidence in Table 2 confirms that the historical malaria is *not* correlated with the current health conditions (measured by anemia and chronic illness/disability). To allay the concern that historical malaria might pick up the current malaria infections, we control for recent malaria incidence (both Plasmodium Vivax and Plasmodium Falciparum infection rates).

[28]It is important to appreciate that labor regulations are not likely to be relevant for our analysis, because we focus on rural labor markets where there is virtually no labor regulation in Sri Lanka.

[29]Note, however, that there is no evidence that ethnolinguistic fractionalization is correlated with the incidence of land restrictions across sub-districts in Sri Lanka. A regression of proportion of land under restrictions on a constant and share of Sinhalese population yields a coefficient close to zero (-0.002) with a very low t statistic (-0.33).

[30]Since malaria exposure in utero can have effects on adult health and education, we exclude cohorts born before 1950, even though nationwide malaria eradication was implemented in 1947.

[31]Note that the probability of such exposure is very low as malaria endemicity was much lower in the sub-districts from where the people were resettled.

5.4 A Control Function Approach

The dependent variables in our analysis are discrete indicating the probability of participation in non-agricultural activities. This requires an approach to instrument variable (IV) estimation which differs from the standard 2SLS regression. To correct for the omitted variable bias, we implement the control function approach developed for the discrete choice models (Blundell and Powell (2001), and Petrin and Train (2009) for multinomial logit model, and Rivers and Voung (1988) for probit model). According to this approach, the discrete choice instrumental variables model is estimated in two steps. First, the endogenous variable is regressed on exogenous instrument(s) and all other regressors in the model. Second, the multinomial logit/probit regression is then estimated by including the residual from the above first-stage regression as an extra variable (i.e., the control function term). To account for any non-linearity in the distribution, we follow the literature and include higher order terms of the residual from the first stage regression.

4 Data

The main data source for the estimation of employment choices is the Household Income and Expenditure Survey, 2002 (HIES, 2002). The survey collected information from a nationally representative sample of 16,924 households drawn from 1913 primary sampling units. The survey covered 17 of Sri Lanka's 25 districts, and 249 of its 322 Divisional Secretariat Divisions (DSDs).[32] The DSD identifier in the HIES (2002) allows us to examine the employment and wages pattern at a much disaggregated geographical level. To define our sample, we use three criteria: (i) we exclude age groups which may have been exposed to historical malaria that afflicted Sri Lanka before 1947; and (ii) we focus on the rural sample where at least 0.1 percent of total land are under restrictions; and (iii) we focus on the male sample to avoid complications from changes in labor force participation decision which is relevant for women. The number of adult male who are employed and were born after 1947 and are currently residing in rural areas is 9,702. Our main estimation is based on a conservative sample of 8,948 individuals who were born after 1950, though robustness checks were performed for the full sample of 9,702 individuals born after 1947.

A central variable for our analysis is the amount of land under LDO restrictions in a DSD. We draw

[32]Data collection in the North and Eastern provinces was not possible due to on-going civil conflicts at the time of survey field work.

this information from the Agricultural Census of 1998. We estimated percentage of agricultural land under LDO leases (including permits and grants). The geographic information including travel time from surveyed DSDs to major urban centers with population of 100,000 or more are drawn from the Geographical Information System (GIS) database. The travel time is estimated using the existing road network and allowing different travel speed on different types of roads.

The most important variable for our instrumental variables analysis is the historical district level malaria prevalence (endemicity) rate. The data on historical malaria prevalence are taken from Newman (1965). The measure for malaria prevalence used in this paper is called Gabaldon's endemicity index (see column 2 in Table 4, P.34, Newman, 1965). This index is based on the estimates of enlarged spleens in children due to malaria, and is a good indicator of the degree to which malaria is high and permanent in a district. The HIES 2002 has complete employment, education, age, gender, ethnicity and religion. Summary statistics of all variables used in the analysis are provided in appendix Table A.1.

In both farming and non-farm activities, a considerable share of those employed are self employed. In the empirical analysis, we thus make a distinction among five different types of employment. The main categories of employment considered are self-employment in agriculture (farming), wage employment in agriculture, self-employment in non-agriculture, wage employment in the manufacturing sector and wage employment in services. The analysis of HIES shows that among 8,948 adult males in the labor force, 14 percent are farmers, 18 percent are agricultural wage laborers, 17 percent employed in manufacturing, 37 percent in services and 14 percent as self-employed in non-agriculture (Table A.1).

5 Empirical Results

The empirical analysis of structural change focuses on two different dependent variables: one representing the broader choice of agriculture versus non-agriculture, and the other making a distinction among five different types of employment choices (self-employment in farming, and in non-agriculture, and wage employment in agriculture, manufacturing and services). The dependent variable in the more disaggregated employment choice regression is an unordered categorical variable indicating sector of employment of individual labor force participants (with self-employment in agriculture as the omitted category). The regressors include a set of individual and area characteristics. As individual level controls, we include education, age (as a measure of experience and cohort effect), marital status dummy and ethnicity/religion dummies.[33] The marital status may capture some differences in motivation and preference.

We control for travel time to large urban cities which is an important determinant of migration costs in a village. This variable also captures effects of differential access to infrastructure and services across areas.[34] As noted before, the area-specific regressors include controls for disease environment, and land and labor productivity indicators. All standard errors are clustered at sub-district (DSD) level. The disaggregated employment choices are estimated by a multinomial logit model.

5.1 Employment Choices and Land Restrictions: Preliminary Estimates

We start with the probit regression for employment choice between non-agriculture and agriculture. The dependent variable in this regression takes the value of one if an individual is employed in non-agricultural activities and zero otherwise. The probit coefficients and the marginal effects evaluated at the mean of all variables are presented in the first column of Table 3. The coefficient of proportion of land under restrictions has a negative sign and is statistically significant at the 1 percent. The estimate of marginal effect implies that a one percentage point increase in land under restrictions decreases the probability of participation in non-agricultural activities by about 0.53 percent. This is consistent with the prediction of proposition 1 when both farm and non-farm products are traded, and labor share in non-agricultural output is less than that in agriculture. This result is also consistent with prediction of proposition 2 when non-agricultural goods are non-traded and face an income inelastic demand. To see if the result for overall non-agricultural employment holds also at disaggregate level particularly when a distinction is made between traded and non-traded non-agricultural goods, we turn to results from multinomial logit regressions.

The results from simple multinomial logit regression (without control function terms) are reported in columns 2-5 of Table 3. The estimates of coefficients of land under restrictions show that an increase in proportion of land under restrictions decreases the log-odd of participation in all types of activities

[33]The omitted category for the ethnicity (religion) dummies is Sinhalese. About 84 percent of Sri Lanka's population are Sinhalese.

[34]Differential access to markets, other infrastructure and services is less of a concern in the context of Sri Lanka. The country invested heavily on providing equitable access to infrastructure and services to its citizens regardless of their residence. For instance, the country has 1 kilometer (km) of road per 1 square km of area and 5 km of road per 1000 inhabitants. There is a government run hospital within 5 km, and primary school within 3 km of a typical household.

including wage labor in agriculture. The logit coefficients are all statistically significant at 1 percent level. The logit results imply that the extra labor endowment in a village due to imposition of land restrictions is absorbed mostly in self employment in agriculture compared with all other activities.

The multinomial logit coefficients in Table 3 do not reflect the way probability of participation in an activity responds to a change in land restrictions. To measure the magnitudes of these responses, we estimate the marginal effects (evaluated at the mean) which are reported in the upper part of each panel in Table 3. The estimates of the marginal effects in the first panel show that the probability of participation in all types of non-farm activities decreases with an increase in the percentage of land under restrictions. Land restrictions have opposite effects on the probability of participation in self and wage employment in agriculture. The marginal effect is precisely estimated for different types of wage employment.

The main conclusion that emerges from Table 3 is that agriculture absorbs a disproportionate share of the increased labor supply resulting from land restrictions. The estimates of coefficients of land under restriction could however be biased if unobserved land and labor productivity heterogeneity is correlated with proportion of land under restrictions. In the following subsection, we tackle this omitted variable bias using a control function approach.

5.2 Employment Choices and Land Restrictions: Control Function Results

As noted earlier in the empirical strategy section (3), a control function is needed to be included to correct for the possible omitted variable biases in the probit and multinomial logit estimates. The control function term is defined from the residuals of the first stage regression where the endogenous variable (proportion of land under restrictions) is regressed on the instrument and rest of the regressors in the model. Panel A of Table 4 shows the control function estimates of the effects of land restrictions on employment probabilities, and panel B reports the first stage regressions and diagnostics for the strength of the instrument. According to the first stage result in panel B, the instrument has the expected positive sign and is statistically significant at the 1 percent level; higher historical malaria in 1935 led to higher incidence of land restrictions. The Kleibergen-Paap F statistics is 24.50 which is more than twice the Stock-Yogo critical value of 9.08 for 10 percent maximal relative IV bias. This confirms that the instrument has excellent strength to identify the effects of land restrictions. The control function term is defined from the residual of this first stage regression. To account for possible nonlinearity, we also include square and cubic terms of the residual as additional control function terms.

The estimates from the control function approach are reported in the upper panel of Table 4. The control function terms are jointly significant at the 5 percent or less level in probit regression as well as for all options in the multinomial logit regression. This result confirms that unobserved heterogeneity is important in determining the substitution probability across alternative activities. The linear term of the control function has a positive sign in all regressions and options. The positive sign of the control function term justifies the worry that unobserved heterogeneity causes downward bias in the simple probit/multinomial logit coefficients.

The inclusion of the control function terms in the probit regression of propensity to participate in non-agricultural activities increases the magnitude of land restrictions coefficient (column 1, Table 4). The estimated coefficient indicates that an increase in land under restrictions reduces the log-odds of participation in non-farm activities. The marginal effect estimated at the mean of all explanatory variables shows that an increase in land under restrictions by 1 percentage point decreases the probability of participation in non-farm activities by 1.38 percent.

The control function estimation results for the multinomial logit model are reported in columns 2-5 of Table 4. The results confirm the negative impact of land restrictions on the log-odds of participation in all types of non-farm activities, and in agricultural wage labor (compared to selfemployment in agriculture). However, the effects on log-odds do not lend to easy interpretation, thus we look at the marginal effects. The marginal effects estimated at the mean of all explanatory variables show that land restrictions have negative and statistically significant effects on the probability of being employed in the manufacturing and services activities. The employment share of services declines the most: a percentage point increase in land under restrictions decreases probability of being employed in services by 0.71 percent. However, the marginal effect estimate suggests a positive effect of land restrictions on the probability of being employed in agricultural wage work. The probability of being employed as a self-employed in the non-agricultural activity on the other hand does not change significantly with an increase in land under restrictions at least at the mean of all explanatory variables.

Among different types of non-agricultural activities, manufacturing is more likely to be traded and services non-traded. The multinomial logit results indicate that employment shares of both of these activities decline in areas with higher proportion of land under restrictions. As the labor share of manufacturing output is significantly lower than that of agriculture, the decline in relative share of manufacturing employment is consistent with the prediction of proposition 1. The result for services on the other hand is consistent with the case where demand for local services is income inelastic (proposition 2). The result for self employment in non-agriculture however casts doubt on the credit channel through which land restrictions may affect employment. Access to capital is more important for self employment and should reduce in employment in this category. Contrary to this expectation, we find no significant impact of land restrictions on the propensity to become self-employment. We checked if the estimated marginal effects of land restriction on probability of self-employment is statistically significant along the entire distribution of proportion of land under restrictions.[35] The estimates confirm the result estimated at the mean value of land restrictions: none of the marginal effects are statistically significant for self employment.

5.3 Robustness Checks for the Effects on the Pattern of Employment

In this subsection, we present results from several robustness checks. The results from the robustness checks are reported in Table 5. To reduce clutter, we report only the marginal effects evaluated at the mean of all explanatory variables in Table 5.

The main results reported in Tables 3-4 are based on samples that excluded cohorts born before 1950. This strategy is motivated by a conservative approach to make sure that our results are not contaminated by possible long-run effects of historical malaria on health and education. One can, however, plausibly argue that the appropriate cut-off is 1947 if the goal is to exclude the cohorts that were potentially exposed to historical malaria. To see if our results are sensitive to the cut-off year chosen, we report the estimates using 1947 sample. The results in first panel in Table 5 show that estimated coefficient of land under restrictions for different employment options are nearly identical to those reported in Table 4.

The non-agricultural activities are observed to concentrate around urban areas and areas with better market access (Fachamps and Shilpi, 2003 and 2005). This spatial concentration is captured by the travel time to urban center variable. Emran and Shilpi (2011) shows that size of the market is an important

[35]The marginal effects are estimated for 10 different quantiles of proportion of land under restrictions starting at lowest 10th quantile.

determinant of location of activities in addition to just access to market. To see if our results are driven by heterogeneity in the market size, we include an additional control: log of urban population within 5 hours of travel time from the sub-district. The estimated effects of land restrictions on employment pattern is nearly unchanged (panel 2, Table 5).

In the third robustness check, we explore if the results are driven by inclusion of the Western province in our sample. A disproportionate share of manufacturing and skilled services activities are located in the Western province where Sri Lanka's main city Colombo is located. This province also has much less land under restrictions. The estimation results that included a dummy for the western province are shown in the third panel of Table 5. A comparison with results in Table 4 shows clearly that though the estimated coefficients are somewhat smaller for all different options of employment, the overall results on the effect of land restrictions on employment pattern remains remarkably unchanged in this case.

Finally, we check the sensitivity of the estimates with respect to the inclusion/exclusion of sub-districts with very high incidence of land restrictions; are the estimates driven by a few outliers in the right tail? The last panel of Table 5 reports the estimates from a sample that excludes sub-districts with proportion of land under restrictions more than 30 percent. The restricted sample has 212 subdistricts and thus loses 30 out of a 242 subdistricts in the full sample. The estimated effects are similar to those in Table 4 with one exception: the marginal effect (at mean values of all explanatory variables) of land restrictions on probability of employment in manufacturing is no longer statistically significant though it still has the negative sign.

The results from the robustness checks are reassuring. Though the magnitudes of estimated coefficients change slightly, the overall conclusions about the impacts of land restrictions on the probability of participation in non-agricultural activities remains unchanged.

5.4 Effects on Wage and Per Capita Consumption

In this section, we test whether the lower wages that result from land restrictions also affect household consumption adversely. Table 6 reports the estimated effects of land market restrictions on log annual wage and log per capita consumption. The IV results are based on a regression specification similar to those reported in Tables 3-5. The estimate for wage shows clearly that the impact of land restrictions on local wage is negative and statistically significant at the 5 percent level. The elasticity estimate in column 2 (at mean) suggest that a one percentage point higher incidence of land restrictions from the mean level (0.13) reduces wages by about 0.125 percent.[36]

The results on per capita consumption are consistent with a priori expectation that a lower wage rate caused by higher incidence of land restrictions is likely to increase poverty. The coefficient and the elasticity estimates from IV results are reported in columns (3) and (4) respectively in Table 6. The estimated coefficient on the land restriction in per capita consumption regression is negative and significant at the 10 percent level. The elasticity estimate implies that a 1 percentage point increase in the land under restrictions starting from the mean (13 percent) reduces per capita consumption in a village by 0.084 percent. The results thus suggest that restrictions on sales and rental of land not only constrain the structural change from agriculture to non-farm sector, they also reduce per capita household consumption, and thus are likely to increase poverty.

5.5 Alternative Mechanisms: Collateral Value, Insecure Property Rights, or Migration Costs?

The results presented and discussed above provide strong evidence that the land market restrictions adversely affect participation in different types of non-agricultural activities significantly. The effects of land restrictions on the pattern of labor allocation are, however, consistent with not only the theoretical analysis in section (3) above that identifies rural-urban migration as the main channel through which the restrictions work. As noted earlier in section (3), a declining share of non-farm employment following land restrictions is also consistent with the predictions from the two other alternative mechanisms: credit channel and guard labor requirements arising from insecurity of property rights. The theoretical analysis in section 3 provides a simple but convincing way to discriminate among different mechanisms at work. When labor is not geographically mobile, a decline in access to credit due to lower collateral value of land will reduce local wage, as the demand for labor is lower. On the other hand, insecurity of property rights implies a negative shock to labor supply to the market as people stay at home to guard the property, leading to higher local

[36]These estimates, although based on a different sample, are similar to the earlier estimates in Emran and Shilpi (forthcoming).

wages. However, with mobile labor, local wages will remain nearly unchanged in either of these cases. Before presenting the test based on local wages, we note that there is no direct restriction on mobility of labor within Sri Lanka in contrast to some other countries such as China (Hukou) and Vietnam (Ho Khau). Indeed, labor market in Sri Lanka has been characterized by considerable geographic mobility with a large proportion of workers employed in manufacturing in cities (e.g. garments) coming from rural areas. Data from census 2012 show that about 27 percent of population in two prominent destination districts (Colombo and Gampaha) are migrants. About a fifth of Sri Lanka's total population are migrants. This evidence suggests substantial mobility of population within Sri Lanka.

With geographically mobile workers, the impact of land restrictions on local wage provides a credible test of whether migration channel is indeed driving the results reported earlier: there should be a negative effect on wages only if migration cost is the relevant channel. As discussed above, the estimates in Table 6 provide clear evidence of a negative effect of land restrictions on local wage, thus strongly reject both the rights insecurity and collateral mechanisms. Note that a lower wage is not consistent with a labor supply response to property rights insecurity, even when geographic mobility of labor is limited.

There are two additional pieces of evidence that contradicts the existence of a significant collateral and credit constraint effect. First, the evidence that land restrictions do not affect self-employment in non-agriculture contradicts a collateral and credit effect, because a reduction in credit availability should affect non-farm self-employment (entrepreneurship) the most. The second piece of evidence comes from the fact that manufacturing is the most capital intensive among all of the sectors in our context. Thus a decline in capital due to tighter credit constraint would mean a larger negative impact of land restrictions on employment probability in manufacturing than services. The evidence discussed above suggests the opposite, probability of employment in services declines by a larger magnitude than that in manufacturing. Direct evidence on outmigration from villages also supports the conclusion that migration costs is the mechanism mediating the effects of land restrictions. Taking advantage of the recently released migration data at the district level from the 2012 population census, we plot outmigration rates against proportion of land under restrictions in Figure 1. It is clear from Figure 1 that propensity of outmigration is lower in areas with higher proportion of land under restrictions.

Taken together, all of the different pieces of evidence above suggest strongly that the negative effects of land restrictions on employment propensities in nonagricultural activities are driven primarily by increasing migration costs due to restrictions on sales and rental of land in LDO areas.

6 Conclusions

This paper investigates the effects of sales and rental restrictions in the land market on structural change from agriculture to non-farm employment in a rural economy. A growing body of economic literature analyzes the effects of land restrictions on labor supply, focusing on property rights insecurity and credit access as the relevant mechanisms. This paper on the other hand emphasizes the increased migration costs due to a ban on land sales. We develop a theoretical model that identifies the conditions under which imposition of land restrictions can lead to an increase in the relative share of employment in agriculture, resulting in a reverse structural change in the rural economy. Equally important, and consistent with the analysis of Emran and Shilpi (forthcoming, 2015), the theoretical model shows that local wages will be lower in a location with a higher proportion of land under restrictions if migration cost is the main causal channel. In contrast, if either access to credit or insecurity of property rights is the channel through which land restrictions operate, then local wages should remain relatively unaffected by land restrictions due to labor mobility.

To provide estimates of the causal effect of land restrictions, we exploit a historical quasi-experiment in Sri Lanka that allows us to use variations in malaria prevalence across districts from the 13th to early 20th centuries to identify the effects of land restrictions on employment choices. The marginal effects from the control function approach to logit and multionomial logit models show that land restrictions reduce the probability of participation in non-agriculture significantly: a one percentage point increase in land restrictions reduces the probability of participation in non-agriculture by 1.38 percent.

The more disaggregated empirical results show that land restrictions have negative and statistically significant effects on the probability of wage employment in manufacturing and services activities, but a positive effect on the probability of being employed in agricultural wage work. The negative effect is particularly large for services: a percentage point increase in land under restrictions decreases the probability of wage employment in services by 0.71 percent. We present evidence that land market restrictions reduce wages in the village labor market, and result in lower household per capita consumption. A 1 percentage point increase in land under restrictions reduces per capita household consumption by 0.084 percent. We provide a rich array of evidence that rejects collateral value and insecurity of property rights as relevant mechanisms through which land restrictions affect employment choices and household consumption. The evidence is strong that the effects are driven primarily by higher migration costs due to restrictions on sales and rental of land. The theoretical and empirical analysis in this paper thus has two important contributions: (i) it shows that land market restrictions can result in a reverse structural change in the rural economy, and lower per capita household consumption, and (ii) it allows us to discriminate among three possible causal mechanisms, and provides clear evidence in favor of the migration costs channel which has been relatively neglected in the recent development literature.

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Online Appendix A: Not for Publication

Wage Determination

We first consider the case where the nonagricultural good is tradable. This simplifies the algebra and help focus on the role played by the migration costs in determining the wage in village labor market. The wage determination in this case is the outcome of the local labor market clearing and the migration equilibrium condition. We normalize the prices to 1, and ignore transport and other costs involved in trading at the international market (we consider transaction costs below when nonagricultural good is assumed to be nontradable).

Given a wage rate w, the demand for labor in different activities can be expressed as:

$$\frac{\partial Q^f}{\partial L^f} = w \quad \Rightarrow L^f = \bar{T}\left(\frac{1-\alpha}{w}\right)^{\frac{1}{\alpha}}$$
$$\frac{\partial Q^n}{\partial L^n} = w \quad \Rightarrow L^n = \bar{K}\left(\frac{1-\beta}{w}\right)^{\frac{1}{\beta}}$$

Denote the equilibrium wage at the initial equilibrium by w^0. Then the local labor market clearing implies:

$$L^S\left(w^0\right) = \int_{h^0}^{1} dF(h) = \bar{T}\left(\frac{1-\alpha}{w^0}\right)^{\frac{1}{\alpha}} + \bar{K}\left(\frac{1-\beta}{w^0}\right)^{\frac{1}{\beta}} \tag{10}$$

Where $L^S(.)$ is the labor supply function to the village labor market, and h^0 is the threshold value such that all households h with distance to the urban center $d_h < d_{h^0}$ decides to migrate. Denoting the indirect utility functions in rural and urban areas by $V^r(.)$ and $V^u(.)$ respectively, the cut-off distance is determined from the following migration equilibrium condition:

$$V^u\left(T_{h^0}, w^u, 1, 1\right) - \varphi\left(d_{h^0}\right) = V^r\left(T_{h^0}, w^0, 1, 1\right) \tag{11}$$

where migration cost faced by a household is heterogeneous and increasing in its distance to the urban market d_h and land owned by the household is an argument in the indirect utility function at urban location because a household can sell the land take the money with them when migrating in the absence of any sales restrictions. Equations (10) and (11) jointly determine the equilibrium values of w^0 and h^0 in the initial equilibrium without any land restrictions.

We are now ready to consider the effects of land restrictions on the local labor market equilibrium. The first thing to note is that given the initial equilibrium wage w^0, for a household $d_h > d_{h^0}$ the imposition of land restrictions has no effect on its migration decision, as they do not migrate irrespective of whether there are restrictions on land sales. However, under land restrictions, a household with $d_h \leq d_{h^0}$ might find it no longer profitable to migrate, because they lose the land when migrating. Note that the households at a given distance differ in terms of land endowment and thus the costs they incur when migrating is heterogeneous. We can thus consider the migration decisions conditional on distance for the subset of households that finds migration desirable without land restrictions. Fix a distance $\hat{d_h} < d_{h^0}$. The migration equilibrium condition given $\hat{d_h}$ is given as follows:

$$V^u\left(0, w^u, 1, 1\right) - \varphi\left(\hat{d_h}\right) = V^r\left(\left(T_{h^*} \mid \hat{d_h}\right), w^0, 1, 1\right) \tag{12}$$

What equation (12) shows is that even though all households at distance $\hat{d_h}$ would choose to migrate without land restrictions, but those with land more than a threshold $T_h > T_{h^*}$ stay back in the village after the imposition of land restrictions. In fact at any given distance $d_h < d_{h^0}$ there will be some households with enough land that they find it undesirable to migrate after the imposition of land restrictions. This increases the supply of labor in the local labor market and drives down the equilibrium wage.

When the nonagricultural good is a nontradable, the equilibrium is characterized by local market clearing of nonagricultural good along with the labor market clearing and migration equilibrium condition. We discuss the market clearing in non-tradable nonagriculture below. A simple way to model nontradability of the nonfarm good is that its equilibrium price in the local market falls in the transaction cost band. We assume that nonagricultural good is imported to the urban area from world market and its price in the urban market is normalized to 1. A unit of nonagricultral good requires μ^n transport costs to bring it to the village market.[37] The nonagriculture is nontradable because its local price $P^n \in (1, 1 + \mu^n)$. Denoting the equilibrium prices before ad after land restrictions as P^{n0} and P^{n*} respectively, we have the following migration equilibrium conditions:

$$V^u\left(T_{h^0}, w^u, 1, 1\right) - \varphi\left(d_{h^0}\right) = V^r\left(T_{h^0}, w^0, 1, P^{n0}\right)$$

$$V^u\left(0, w^u, 1, 1\right) - \varphi\left(\hat{d_h}\right) = V^r\left(\left(T_{h^*} \mid \hat{d_h}\right), w^0, 1, P^{n*}\right)$$

[37]For simplicity, we ignore trading costs for agriculture. Alternatively, one can normalize the trading costs inclusive price to 1.

Employment Choices and Labor Allocation Under Land Restrictions: The Case of Nontradable Non-farm Good

(a) Labor Allocation with Income elasticity of demand for nonagricultural goods greater than or equal to unity:

We consider the case where wage is given w and$\gamma^f \geqslant 0$ and $\gamma^n = 0$. The demand for nonagricultural good can be defined as:

$$C^n = \frac{\tau_n}{P^n}[Y - \gamma^f]$$

Total village income consists of agricultural income ($Y^f = Q^f$) and non-agricultural income ($Y^n = P^n Q^n$). Setting demand for non-agricultural good to its supply:

$$\begin{aligned} P^n C^n &= P^n Q^n = \tau_n[Y - \gamma^f] = \tau_n[Q^f + P^n Q^n - \gamma^f] \\ P^n Q^n &= \frac{\tau_n[Q^f - \gamma^f]}{1 - \tau_n} \Longrightarrow (P^n)^{1/\beta} = \frac{\tau_n[\bar{T}(1-\alpha)^{(1-\alpha)/\alpha} w^{(\alpha-\beta)/\alpha\beta} - \gamma^f w^{(1-\beta)/\beta}]}{(1-\tau_n)\bar{K}(1-\beta)^{(1-\beta)/\beta}} \end{aligned}$$

The employment ratio can be derived as:

$$C^n = \frac{\tau_n}{P^n}[Y - \gamma^f]$$

$$S_{nf} = \frac{(1-\beta)^{1/\beta}\bar{K}w^{(\beta-\alpha)/\alpha\beta}(P^n)^{1/\beta}}{(1-\alpha)^{1/\alpha}\bar{T}} = \frac{\tau_n(1-\beta)}{(1-\tau_n)(1-\alpha)}[1 - \frac{\gamma^f w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}}] \tag{13}$$

Equation (13) can be used to examine the impact of land restrictions on the ratio of nonagriculture to agricultural employment. From equation (13), it is seen easily that $\frac{\partial s_{nf}}{\partial w} < 0$.

Note also that if $\gamma^f = 0$, the utility function becomes homothetic (Cobb-Douglas in this case). In this homothetic case, it follows from equation (13) that $\frac{\partial s_{nf}}{\partial w} = 0$. In other words, with homothetic preference, the relative employment shares of nonagriculture and agriculture remains unaffected by land restrictions (proposition 2(ii)).

(b) Labor Allocation with Income inelasticity of demand for nonagricultural goods:

This case is represented by the following restrictions on subsistence quantities: $\gamma^f = 0$ and $\gamma^n \geqslant 0$. The demand for nonagricultural good can now be described as:

$$C^n = (1-\tau_n)\gamma^n + \frac{\tau_n Y_k}{P^n}$$

The equilibrium price for nonagricultural good can be determined from the following equilibrium condition:

$$\begin{aligned} (1-\tau_n)P^n[Q^n - \gamma^n] &= \tau_n Q^f \\ \frac{P^{n(1/\beta)}\bar{K}(1-\beta)^{(1-\beta)/\beta} w^{(\beta-\alpha)/\alpha\beta}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}} - \frac{P^n \gamma^n w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}} &= \frac{\tau_n}{(1-\tau_n)} \end{aligned} \tag{14}$$

Since equation (14) is non-linear in P^n, it cannot be solved directly. Instead define the following system of equations:

$$\begin{aligned} S_{nf} - \frac{(1-\beta)^{1/\beta}\bar{K} w^{(\beta-\alpha)/\alpha\beta}(P^n)^{1/\beta}}{(1-\alpha)^{1/\alpha}\bar{T}} &= 0 \\ \frac{(1-\alpha)}{(1-\beta)} S_{nf} - \frac{P^n \gamma^n w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}} &= \frac{\tau_n}{(1-\tau_n)} \end{aligned}$$

Taking partial derivatives and expressing it in matrix form, we have the following system which can be solved to find out how s_{nf} and P^n change in response to wage:

$$\begin{bmatrix} 1 & -\frac{s_{nf}}{\beta P^n} \\ \frac{(1-\alpha)}{(1-\beta)} & -\frac{z_1}{P^n} \end{bmatrix} \begin{bmatrix} \partial s_{nf} \\ \partial P^n \end{bmatrix} = \begin{bmatrix} \frac{(\beta-\alpha)s_{nf}}{\alpha\beta w} \\ \frac{(1-\alpha)z_1}{\alpha w} \end{bmatrix} \partial w$$

We define $z_1 = \frac{P^n \gamma^n w^{(1-\alpha)/\alpha}}{(1-\alpha)^{1-\alpha/\alpha}\bar{T}}$ and $z_2 = \dfrac{z_1 S_{nf}}{\alpha\beta w}$. The response of employment ratio to wage can be solved as:

$$\frac{\partial S_{nf}}{\partial w} = \frac{(1-\beta)\, z_2}{\frac{(1-\alpha)}{(1-\beta)\beta} S_{nf} - z_1} > 0$$

The numerator is unambiguously positive. The denominator is positive because $\frac{(1-\alpha)}{(1-\beta)\beta}S_{nf} - z_1 > \frac{(1-\alpha)}{(1-\beta)}S_{nf} - z_1 = \frac{\tau_n}{(1-\tau_n)} > 0$.Thus ratio of employment in nonagriculture to agriculture changes positively with wage and hence negatively with the proportion of land under restrictions.

Online Appendix B: Not for Publication

Land Market Restrictions in Sri Lanka: Historical Background

The present day land tenure system in Sri Lanka is largely an outcome of colonial laws and its subsequent amendments. During the early colonial period, the Crown Lands Encroachment Ordinance of 1840 transferred all lands without private title—unoccupied or uncultivated land (abandoned due to malaria), forests and waste land– to the state. As a result of this Ordinance, the British Crown became the owner of nearly all lands, as landownership in Sri Lanka was governed by local customs and few in the peasantry possessed clear formal titles (De Silva,1981; Peebles, 2006). Between 1840 and 1870, Crown land suitable for coffee plantation were purchased rapidly by British officials and investors as well as some wealthy Sri Lankans.[38] After the complete demise of coffee crop due to leaf disease by 1875, plantations diversified into other crops such as tea, rubber etc. The expansion of plantations on the basis of Crown lands subsided by the 1920s.[39] The point to emphasize here is that purchase of Crown lands by private individuals/plantation owners during the late 19th and early 20th century was driven by suitability of land for coffee production, a crop which had virtually disappeared from Sri Lanka.[40]

The Land Development Ordinance (LDO for short) of 1935 initiated a program of making Government-owned agricultural land available for private household use. The state introduced a system of protected tenure under which recipients of LDO land had the right to occupy and cultivate the land in perpetuity subject to restrictions imposed on sale, leasing and mortgaging, and conditions related to abandoning or failing to cultivate the land. While subsequent amendments have weakened some conditions on mortgage

[38]Peebles (2006) states that land in Kandyan hills were particularly suitable for coffee plantation. This land was reclaimed from the Kandyan peasantry regardless of the status of their titles and was to be sold to plantation owners later on.

[39]The larger plantations were nationalized during the early 1970s, and are now run by private companies under long-term lease arrangements with the government.The importance of plantation crops in Sri Lankan economy today has also declined substantially with an increasing share of land going to paddy and other field crops. Indeed, the estate/plantation sector now accounts only for 5.5 percent of Sri Lanka's population in 2006. Only 8.6 percent of our sample comes from estate/plantation areas. We use a dummy for estate sector in our regressions.

[40]The coffee land (hilly land) are not necessarily considered as particularly suitable for paddy and other field crops which are now the mainstay of Sri Lanka's smallholder agriculture.

(now allowed for loans from public banks) and limited transferability (with special permission from the government which was uncommon during the period covered by the survey), the basic provisions of unitary succession and ban on subdivision and sales of plots and land rental remain largely intact (see Peebles, 2006; De Silva, 1981; World Bank, 2008 for the history of land reforms). Distribution of LDO land took place mainly after Sri Lanka's Independence in 1948 and much of the land was distributed under various settlement schemes. The settlement schemes brought landless Sinhalese people from the relatively malaria-free south to the historically malaria infested DSDs in the dry zone. The LDO leases today coexist with complete private holding in the same location (World Bank, 2008). The share of land under LDO leases varies significantly across areas in our sample (the maximum is 63 percent and minimum 0.2 percent) which is critical for estimation of the effect of LDO restrictions on equilibrium wages.

Table 1: Relationship between historical malaria and current productivity (yield)

	Rice	Banana	Ground Nut	Other Oilseeds
Proportion of Area Under LDO	1,589**	-1,645	843.5***	-314.1
	(2.336)	(-1.095)	(3.007)	(-0.236)
R-squared	0.199	0.080	0.285	0.069
Historical Malaria Incidence	-2.546	-18.03	-0.00790	-15.65
	(-0.350)	(-1.658)	(-0.00254)	(-1.432)
R-squared	0.170	0.088	0.100	0.088
Observations	118	98	57	101

Regressions control for average rainfall, slope, current malaria incidence and dummy for within 5km of a river. Robust t statistics in parentheses: * significant at 10%; ** significant at 5%; *** significant at 1%. Source: Emran and Shilpi (forthcoming)

Table 2: Land under Restrictions, Historical Malaria and Health Status

	Anemia among non-pregnant women			% suffering Chronic Illness/disability		
	Mild/Moderate	Severe	Any	Male	Female	All
Proportion of Area Under LDO	-34.726	0.459	-34.164	-7.264	7.000	-3.421
	(1.79)*	(0.38)	(1.81)*	(1.39)	(0.51)	(0.47)
Historical Malaria Incidence	-1.181	-0.045	-1.203	-0.134	0.071	-0.085
	(1.21)	(0.77)	(1.26)	(0.51)	(0.11)	(0.24)

Robust z statistics in parentheses. Standard errors corrected for clustering at village level
* significant at 10%; ** significant at 5%; *** significant at 1%, Source: Emran and Shilpi (forthcoming)

Table 3: Employment Choice and Land Restrictions: Probit and Multinomial Logit Results

	Non-Farm Employment	Self-employ. Non-Agriculture	Wage employment in Agriculture	Manufacturing	Services
	(Marginal Effects)		(Marginal effects)		
Proportion of Area Under LDO	-0.527***	-0.0819	0.148*	-0.174***	-0.205*
	(-5.174)	(-1.285)	(1.750)	(-3.460)	(-1.749)
	(Probit Coefficients)		(Multinomial Logit Coefficients)		
Proportion of Area Under LDO	-1.451***	-3.006***	-1.592***	-3.511***	-2.937***
	(-5.174)	(-4.342)	(-2.821)	(-6.245)	(-5.111)
Controls					
Individual characteristics	Yes	Yes	Yes	Yes	Yes
Household characteristics	Yes	Yes	Yes	Yes	Yes
Area characteristics	Yes	Yes	Yes	Yes	Yes
Observations	8,948	8,948	8,948	8,948	8,948

All regressions include individual's age, education, marital status, and dummies for household's religion/ethnicity and a full set of area controls including travel time to cities, disease environment, land and labor productivity indicators and ethnic composition.
Robust t statistics in parentheses. Standard errors corrected for clustering at the sub-district level (DSD)
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 4: Employment Choice and Land Restrictions: Control functions Results

	Non-Farm Employment	Self-employ. Non-Agriculture	Wage employment in Agriculture	Manufacturing	Services
	(Marginal Effects)		(Marginal effects)		
Proportion of Area Under LDO	-1.382***	-0.147	0.594***	-0.424**	-0.713***
	(-5.094)	(-0.950)	(2.593)	(-2.450)	(-2.589)
	(Probit Coefficients)		(Multinomial Logit Coefficients)		
Proportion of Area Under LDO	-3.817***	-6.430***	-2.037	-8.032***	-7.185***
	(-5.094)	(-3.721)	(-1.197)	(-4.183)	(-4.821)
First stage Regressions					
Malaria			0.00354***		
			(4.950)		
Relevance of Instruments					
Kleibergen-Paap/Angrist-Pischke F			24.5		
Stock-Yogo 10% max. rel. IV bias			9.08		
Controls					
Individual characteristics	Yes	Yes	Yes	Yes	Yes
Household characteristics	Yes	Yes	Yes	Yes	Yes
Area characteristics	Yes	Yes	Yes	Yes	Yes
Observations	8,948	8,948	8,948	8,948	8,948

All regressions include individual's age, education, marital status, and dummies for household's religion/ethnicity and a full set of area controls including travel time to cities, disease environment, land and labor productivity indicators and ethnic composition.
Robust t statistics in parentheses. Standard errors corrected for clustering at the sub-district level (DSD)
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 5: Employment Choice and Land Restrictions: Robustness of Control functions Results

	Non-Farm Employment	Self-employ. Non-Agri.	Wage employment in Agriculture	Manufacturing	Services
1947+ sample			(Marginal effects)		
Proportion of Area Under LDO	-1.436***	-0.164	0.577***	-0.421**	-0.760***
	(-5.314)	(-1.070)	(2.581)	(-2.510)	(-2.811)
Population within 5 hrs of travel time					
Proportion of Area Under LDO	-1.510***	-0.132	0.657***	-0.488***	-0.791***
	(-5.512)	(-0.951)	(2.891)	(-3.214)	(-2.924)
Controlling for Western province					
Proportion of Area Under LDO	-1.225***	-0.141	0.442*	-0.429**	-0.558**
	(-4.478)	(-0.940)	(1.947)	(-2.310)	(-1.976)
Restricted sample: proportion LDO<0.3					
Proportion of Area Under LDO	-2.904***	-0.464	1.535**	-0.409	-1.821**
	(-4.069)	(-1.330)	(2.526)	(-0.744)	(-2.526)

All regressions include individual's age, education, marital status, and dummies for household's religion/ethnicity and a full set of area controls including travel time to cities, disease environment, land and labor productivity indicators and ethnic composition.
Robust t statistics in parentheses. Standard errors corrected for clustering at the sub-district level (DSD)
* significant at 10%; ** significant at 5%; *** significant at 1%

Table 6: Land market restrictions and wages

	Log(Real Annual wage)	Log (real per capita household expenditure)
Proportion of Area Under LDO	-0.964**	-0.644*
	(-2.164)	(-1.794)
First stage- regressions		
Historical Malaria	0.00307***	0.00346***
	(4.420)	(4.795)
Relevance of Instruments		
Kleibergen-Paap/Angrist-Pischke F	19.72	22.99
Stock-Yogo 10% max. rel. IV bias	9.08	9.08
Individual/head’s characteristics	Yes	Yes
Household characteristics	Yes	Yes
Area characteristics	Yes	Yes
Observations	6,428	11,091

All regressions include individual's age, education, marital status, and dummies for household's religion/ethnicity and a full set of area controls including travel time to cities, disease environment, land and labor productivity indicators and ethnic composition.
Robust t statistics in parentheses. Standard errors corrected for clustering at the sub-district level (DSD)
* significant at 10%; ** significant at 5%; *** significant at 1%

Figure 1: Land Market Restrictions and Out-migration

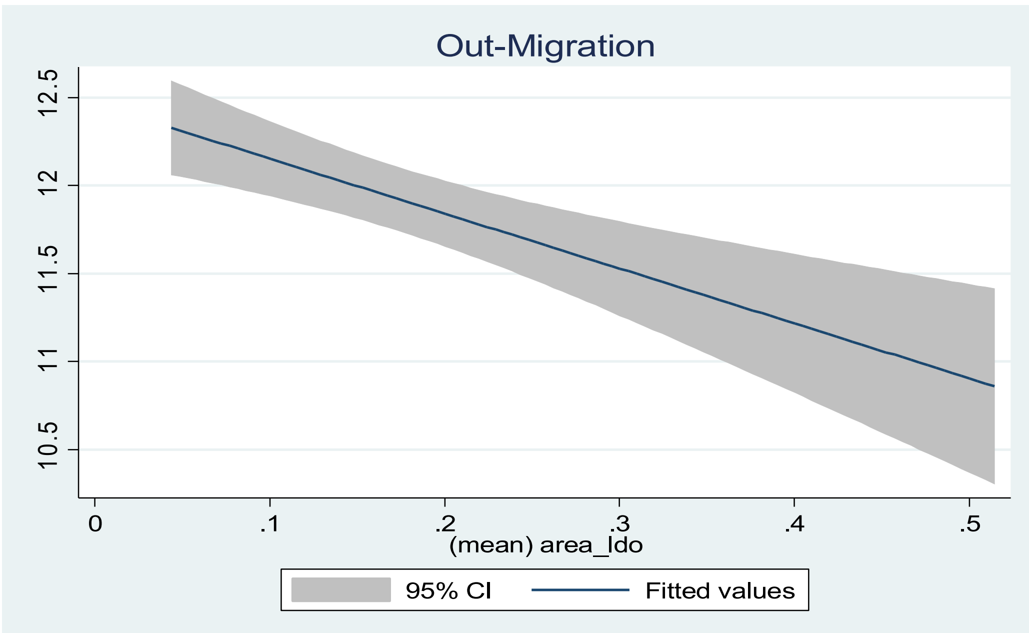

Table A.1: Summary Statistics

Variable	Mean	Median	Std. Dev.
Proportion of Area Under LDO Leases	0.13	0.07	0.14
Travel Time to Large City (hour)	2.65	2.00	2.43
Female Wage (annual in rupees)	60578	49263	51670
Malaria incidence (spleen rate)	23.23	12.20	20.97
P.Vivax (1000)	0.05	0.01	0.12
P.Fac. (1000)	0.02	0.00	0.03
Share of Sinhalese in population	0.86	0.92	0.19
Rainfall (milimetre)	2411	2269	819
Slope (%)	10.65	7.23	9.03
River within 5 km (yes=1)	0.34	0.00	0.47
Age (Year)	34.62	34.00	9.12
Education Level (year)	8.35	10.00	3.45
Married (yes=1)	0.71	1.00	0.45
Christian (yes=1)	0.05	0.00	0.22
Muslim (yes=1)	0.05	0.00	0.21
Buddist (yes=1)	0.82	1.00	0.38
Moor (yes=1)	0.05	0.00	0.21
Tamil (yes=1)	0.09	0.00	0.29
Estate (yes=1)	0.09	0.00	0.29