Creating a Market: Land Reallocation in Vietnam's Agrarian Transition

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Abstract: While liberalizing key factor markets is a crucial step in the transition from a socialist control-economy to a market economy, the process can be stalled by imperfect information, high transaction costs and covert resistance from entrenched interests. The paper studies land-market adjustment in the wake of Vietnam's reforms aiming to establish a free market in land-use rights following de-collectivization. Inefficiencies in the initial administrative allocation are measured against an explicit counter-factual market solution. Our tests using a farm-household panel data set spanning the reforms suggest that land allocation responded positively but slowly to the inefficiencies of the administrative allocation. We find no sign that the transition favored the land rich or that it was thwarted by the continuing power over land held by local officials.

Key words: Land reform, decentralization, land markets, Vietnam

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1. Introduction

Vietnam's agrarian transition has closely followed a now classic policy scenario for economies in transition. First one privatizes the main productive assets — in this case agricultural land-use rights — then one legalizes their free exchange. In the first step, the de-collectivization of agriculture meant that the land that had previously been farmed collectively was to be allocated by administrative means across households within each commune. Without a market mechanism to guide the process, it can be expected that inefficiencies in land allocation would remain, with some households having too much land relative to a competitive market allocation and some too little. Ravallion and van de Walle (2004) document evidence of such inefficiencies in the initial administrative allocation at the time of de-collectivization.

The second step was reforming land laws so as to create the framework for a free market in agricultural land-use rights. While land remained the property of the state, Vietnam reformed land laws in 1993 to introduce official land titles and permit land transactions for the first time. Having removed legal obstacles to buying and selling land-use rights, the expectation was that land would be re-allocated to eliminate the initial inefficiencies in the administrative assignment.

However, the outcomes are far from clear on *a priori* grounds. Land was not the only input for which the market was missing or imperfect. Indeed, as a stylized fact, other factor markets are still poorly developed in rural areas, which is likely to limit the efficiency gains from freeing up land transactions on their own. Pervasive market failures fuelled by imperfect information and high transaction costs could well have stalled the process of efficiency-enhancing land re-allocations during the transition.

The local state continued to play an active role, though it is unclear whether the continuing exercise of communal control over land was synergistic with market forces or opposed to them. Possibly the local political economy operated to encourage otherwise sluggish land re-allocation to more efficient users.[2] Or it may have worked against an efficient agrarian transition, given risk-market failures and limitations on the set of redistributive instruments. Resistance to the transition on the part of local cadres may then be interpreted as a form of social protection, recognizing the welfare risks that a free market in land might entail. Or it might be argued that the frictions to agrarian transition stemming from the local political economy worked against both greater equity and efficiency; while socialism may have left in-grained preferences for distributive justice, the new possibilities for capture by budding local elites — well connected to the local state authorities — would not presumably have gone unnoticed.

The *ex post* outcomes of this reform in Vietnam are also of interest to neighboring China, which is embarking on a similar process of liberalizing the exchange of agricultural land-use rights (McGregor and Kynge, 2002). As in Vietnam, the hope is that land will be reallocated to more efficient users, and that inefficient farmers will switch to (rural or urban) nonfarm activities. And, as in Vietnam, there are concerns in China that local officials and elites will subvert the process.

This paper offers what we believe to be the first test of whether the classic policy scenario of privatization followed by liberalized exchange has actually worked in a developing transition economy. In particular, the paper assesses whether the post-reform allocation of annual agricultural land-use rights in Vietnam redressed the inefficiencies of the initial administrative allocation. We first measure the extent of inefficiencies in the pre-reform administrative allocation, judged relative to an explicit counter factual. We then see to what extent those inefficiencies can explain the subsequent land re-allocations in a panel of farm

[2] In the context of rural China, Benjamin and Brandt (2002b) argue that administrative land re-allocations served an efficiency role given other market failures.

households spanning the change in land laws, with controls for other "non-market" factors bearing on land allocation.

The following section describes key features of the setting. Section 3 describes our approach to testing whether the post-reform land re-allocation responded to the household-specific efficiency losses from the pre-reform administrative allocation. Our data are described in section 4. We then present and interpret our results in section 5. Section 6 concludes.

2. Land allocation in Vietnam's agrarian transition

In the late 1980s, Vietnam abandoned socialist agriculture, whereby rural workers had been organized into "brigades" that jointly farmed the commune's land. The central government gave local authorities the power to allocate to individual households the agricultural land that had been farmed collectively. De-collectivization was followed in 1993 by a new land law that introduced official land titles in the form of certificates and permitted land transactions for the first time since communist rule began. Land remained the property of the state, but usage rights could be legally transferred and exchanged, mortgaged and inherited (Cuc and Sikor, 1998).

The central government's explicit aim in introducing this new land law was to promote greater efficiency in production by creating a market in land-use rights (see, for example, de Mauny and Vu, 1998).[3] The expectation was that, after these legal changes, land would be re-allocated to assure higher agricultural output, taking account of such factors as farmers' abilities, supervision costs of hiring labor and the micro-geographic organization of land plots.

Despite the center's aim of creating a free market in land-use rights, local authorities retained a degree of both formal and informal power over land. Local cadres oversee titling,

3 This was one element of a set of reforms to increase agricultural output. Other reforms include relaxing trade restrictions, which improved farmers' terms of trade; see Benjamin and Brandt (2002a).

land-use restrictions and land appropriation for infrastructure projects. Sikor and Truong (2000) describe well how the reforms were mediated by village institutions in Son La, a northern uplands province:

> "Local cadres were located at the intersection of the state and villages. A large majority of them came from local villages and maintained close ties with their kin and fellow villagers. The close ties between local cadres and villagers influenced the activities of the local state. Local cadres attempted to accommodate villagers' interests, sometimes even when they contradicted national policy." (Sikor and Truong, 2000, p.33).

In these circumstances, it would be wrong to view the land-market reform as necessarily undermining the power of the local state over land allocation. Indeed, staff of one NGO argued that the pro-market reforms enhanced the power of the state over land usage (Smith and Binh, 1994). Although both the 1988 and 1993 land laws extended land use rights for "stable and long-term use" it is widely believed that many local authorities continue to re-allocate land periodically by administrative means (particularly in the north), such as in response to demographic changes and new family formations.

There is anecdotal evidence that the continuing power of the local state stalled the reforms in some parts of Vietnam. Writing a few years after the 1993 Land Law, Smith (1997) reports that in one northern province (Ha Tinh) the major commercial bank that lent for agricultural purposes had not yet accepted a single land-usage certificate as collateral for a loan. The resistance of local officials to have the land sold to an outsider was one of the reasons given by the bank; another was that the bank was unsure it would ever find a buyer for the land should it foreclose on the loan. However, this should not be generalized; indeed, the same study reported cases of land certificates being accepted as collateral in another province.

Just how much the local state has inhibited the development of a land market is unclear. It appears that land transactions can by-pass state control. There have been reports of land transactions without titles (Smith, 1997; de Mauny and Vu, 1998). Possibly a quasi-market has emerged despite the continuing intrusions of the local state.

There have also been concerns about rising landlessness stemming from the reforms. A report by ActionAid staff exemplifies these concerns; while presenting no supportive evidence, the report predicted that the reforms would lead to:

> "..a greater concentration of land ownership, a greater disparity in wealth throughout the rural community and a possible increase in the phenomenon of landlessness and full-time agricultural wage labour." (Smith and Binh, 1994, p.17.)

There have been anecdotal reports of rising landlessness, notably in the south's Mekong Delta region (de Mauny and Vu, 1998; Lam, 2001b).

Some of the efforts made to avoid rising landlessness may well have had perverse effects. There are reports that, in response to central Communist Party concerns about rising landlessness in the late 1990s, some local officials in the Mekong Delta tried to stop poor families selling their land (de Mauny and Vu, 1998). Whether this would be in the interests of such families is a moot point. The consequent devaluation of their main non-labor asset could make the poor worse off, depending on whether any compensation is provided locally to those prevented from selling their land as a response to some negative shock. It is likely that transfers still happened despite such policies, though the transactions would become informal, and possibly on less favorable terms for those forced to sell their land because of adverse shocks.

There were regional differences that are likely to have mattered to the pace of the agrarian transition. After re-unification in the mid-1970s, farmers in the Mekong Delta (in the south) had resisted collectivization, and by the time the country de-collectivized 13 years later, less than 10 percent of all of the region's farmers had been organized into collectives. By contrast, virtually all of the crop land in the north and the south's Central Coastal provinces was collectivized by that time (Pingali and Xuan 1992; Ngo 1993).

The market economy was thus more developed in the Mekong Delta at the beginning of the transition. Rural per capita income growth was also higher in the south over this period, fuelled in part by improvements in farmers' terms of trade arising from external trade reforms; Benjamin and Brandt (2002a) report a 95% increase in real income per person in the south over 1993-98, versus 55% in the north. Such rapid growth in real incomes may well have dampened the pressure to secure the efficiency gains from land re-allocation in the south.

There were other pre-reform regional differences. The distribution of land was more equal in the north.[4] The collectivization of agriculture in the north over roughly a generation fostered a more equitable allocation at the time of de-collectivization. In the south, the fall back position was the land allocation pre-unification, and the realized allocation was more unequal than in the north (Ravallion and van de Walle, 2004). Lower inequality in the north may well have made it easier to achieve cooperative outcomes, including more efficient assignments of land-use rights.[5]

Another regional difference that could well have bearing on land allocation can be found in the performance of (formal and informal) institutions that deal with risk. The safety net in rural areas of Vietnam is largely community-based; central and provincial programs tend to have very limited coverage (van de Walle, 2002). It is widely believed that villages in the north are better organized socially than in the south, so that when a farm household in the north suffers a negative shock (such as crop damage or ill-health) it will almost never need to sell land to cope. For example, writing about Son La province, Smith reports that:

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[4] This difference shows up in the results from the VLSS of 1992/93. The coefficient of variation in the log of allocated annual agricultural land was 8.3% in the North's Red River Delta, versus 15.3% in the south's Mekong Delta (Ravallion and van de Walle 2004). (Among the five regions for which the sample size was deemed adequate, these were the regions with lowest and highest land inequality respectively.)

[5] For an excellent review of the theoretical arguments as to why high inequality can impede efficiency see Bardhan et al., (1999).

> "..there is a tendency for the local authorities to seek to protect households from the dangers of a market in land, despite the provisions of the 1993 Law. This constitutes an attempt to protect poor households who may be tempted to sell their land for short term gain and lose their principal means of subsistence." (Smith, 1997, p.11.)

By contrast, an Oxfam team in the province of Tra Vinh in the Mekong Delta (in which the NGO had been working for a few years) reported that:

> "The crucial problem is that there are no safety nets for helping households who encounter temporary crises. … It is no surprise that many families resort to transferring or mortgaging their land, discounting the future to cope with the current crisis" (de Mauny and Vu, 1998, p.23).

This difference between the north and the south is no doubt a legacy of the longer period of collective organization in the north. However, the more equal land allocation in the north after breaking up the collectives could well have facilitated this, by making it easier to continue to achieve quasi-cooperative arrangements within communities. Better insurance in the north is likely to have also made it easier for land transactions to be made on efficiency grounds. Land re-allocations in the south, by contrast, are likely to have been less flexible, since land would be more likely to be held as insurance than in the north.

These observations suggest that it would be naïve to think that simply legislating the pre-requisites for a competitive land market in this setting would make it happen. The reality is more complex and uncertain, given the institutional/historical context. The continuing power of local cadres could have served to either undermine the expected efficiency gains from the center's reforms (to assure that other distributional goals were achieved) or to help secure those gains. This will depend in large part on the resolution of the likely power struggle at local level between potential gainers and losers.[6] In the rest of this paper we will study the outcomes of this process of post-reform land re-allocation, given its institutional and historical context.

[6] This echoes recent analyses of the case for community-based welfare programs (Bardhan and Mookherjee, 2000; Galasso and Ravallion, 2004).

3. Modeling land allocation

The main hypothesis to be tested is that land re-allocation during Vietnam's agrarian transition helped offset prior inefficiencies in the administrative allocation. To test this, we need to explicitly characterize the extent of inefficiency in the initial allocation. Then we will see how subsequent re-allocations of land responded to the measured inefficiencies.

3.1 Gainers and losers from the initial administrative allocation

The administrative allocation at de-collectivization gives L_i^A of land to household *i* where *i*=1,..,*n*. The administrative allocation need not be efficient in the specific sense of maximizing aggregate output or consumption. Suppose that holding L_i of land yields an output of $F(L_i, X_i)$ for household *i* where X_i is a vector of exogenous household characteristics. We assume that the function F is increasing and strictly concave in L_i. The household also has (positive or negative) non-farm income, $Y(X_i)$.[7] The household consumes its current income:[8]

$$C_i = C(L_i, X_i) = F(L_i, X_i) + Y(X_i) \tag{1}$$

Our efficiency counterfactual is the allocation that maximizes the commune's aggregate current consumption, as given by:

$$(L_1^*,..,L_n^*) = \arg\max\left[\sum_{i=1}^{n} C(L_i, X_i) \,\middle|\, \sum_{i=1}^{n} L_i = n\bar{L}\right] \tag{2}$$

where $\bar{L}$ is mean land availability in the commune. The solution equates $F_L(L_i^*, X_i)$ with the multiplier λ on aggregate land in (2), giving:

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[7] To the extent that non-farm income depends on landholding, this can be interpreted instead as that component of income that is not dependent on landholding.

[8] We ignore saving/dissaving and borrowing/lending; incorporating these features would complicate the model in unimportant ways for our purposes.

$$L_i^* = L(X_i, \lambda) \ (i=1,..,n) \tag{3}$$

We call this the "consumption-efficient allocation." This is also the competitive equilibrium assuming that utility depends solely on consumption and allowing costless re-contraction in each state of nature. In the market allocation, each household's consumption will be $F(L_i, X_i) + Y(X_i) - \lambda L_i$ where λ is the market rental price for land. Demands then equate $F_L(L_i, X_i) = \lambda$ over all *i*, which is the allocation that maximizes aggregate consumption.

In our empirical implementation, we assume that (1) takes a specific parametric form:

$$\ln C_i = a + b \ln L_i + X_i c + \varepsilon_i \tag{4}$$

where *a*, *b* and *c* are parameters and ε_i is a white noise error process. Given data on *X*, and estimates of the parameters and error term, we then calculate the consumption efficient allocation to each household. For 0<*b*<1 the solution is $L_i^* = \exp[(\ln(b/\lambda) + X_i c + \nu_i)/(1-b)]$.

We can postulate a general measure of the efficiency loss from the administrative allocation of the form $\tau_i = \tau(L_i^*, L_i^A)$ where the function τ is strictly increasing in L_i^* and strictly decreasing in L_i^A. Thus τ_i measures household *i*'s land shortfall in the administrative assignment relative to the efficient allocation of land to that household. Naturally we want the function τ to have the property that $\tau(L, L) = 0$. We assure this by adopting the functional form:

$$\tau(L_i^*, L_i^A) = \phi(L_i^*) - \phi(L_i^A) \tag{5}$$

for some strictly increasing function ϕ.

We can embrace a reasonably wide range of possible empirical measures of the efficiency loss by restricting attention to the class of parametric functions: $\phi(L) = (L^\eta - 1)/\eta$

where $\eta \in [0,1]$. The two extreme cases are (i) proportionate differences, in which $\eta = 0$, implying that $\tau_i = \ln(L_i^* / L_i^A)$ (noting that $\lim_{\eta \to 0}(L^\eta - 1)/\eta = \ln L$); and (ii) absolute differences ($\eta = 1$) whereby $\tau_i = L_i^* - L_i^A$.

3.2 Modeling the post-reform land re-allocation

We only observe a single time interval in the process of land re-allocation after legalizing market transactions and we should not, of course, assume that the process has reached its long-run solution by the end of the period of observation. However, we do assume that the dynamic process will eventually converge to a unique long-run equilibrium, which depends on the competitive market allocation of land to that household but can also be influenced by the household's power in local decision making about the allocation of use rights. That power could be exercised through the market or through political processes.

The new allocation observed at a date after the reform is $(L_1^R, L_2^R, ..., L_n^R)$. Let $\rho_i = \rho(L_i^R, L_i^A)$ denote a measure of the extent of land re-allocation.[9] We clearly want $\rho(L_i^R, L_i^A)$ to be strictly increasing in L_i^R and decreasing in L_i^A with $\rho(L, L) = 0$. We also want to assure that if $\rho(L_i^R, L_i^A) = \tau(L_i^*, L_i^A)$ then $L_i^R = L_i^*$; if the land re-allocation to household *i* exactly matches the initial efficiency loss then the household must have reached the market solution. These conditions require that ρ and τ have the same functional form i.e., $\rho_i = \phi(L_i^R) - \phi(L_i^A)$.

9 We do not assume that $\sum L_i^R - L_i^A = 0$. Thus land "re-allocation" can come with higher total acreage.

To see how land allocation responded to initial inefficiencies in the administrative assignment we begin by studying the non-parametric regression of ρ on τ:

$$\rho_i = f(\tau_i) + \nu_i \tag{6}$$

where $f(\tau_i) \equiv E[\rho_i | \tau_i]$ in which the expectation is formed over the distribution of the random error term ν. In the special case with $f(0) = 0$ and $f'(\tau_i) = 1$ for all τ, there are no systematic non-market constraints on land re-allocation, so $L_i^R = L_i^*$ in expectation. Adjustment to the market solution is then complete within the period of observation. More generally one can allow $0 \leq f'(\tau_i) \leq 1$ in which case we have a nonlinear partial adjustment model by which land holdings adjust to any discrepancies between the administrative allocation and the market solution, but in which the process need not be complete in the period of observation. With repeated observations, L_i^* will be reached whatever the initial start value of the process (in this case, the administrative allocation at de-collectivization). The slope, $f'(\tau_i)$, is the "partial adjustment coefficient" for household *i*, giving the speed at which initial inefficiencies are eliminated.

The partial adjustment model described above is questionable from a number of points of view. One concern is the possibility of measurement error in the data for the initial land allocation. Classical measurement error in L_i^A will bias the Ordinary Least Squares (OLS) estimate of the linear partial adjustment coefficient, though the direction of bias is ambiguous in this case; the usual attenuation bias will be at least partly offset by the fact that the measurement error also appears positively in the dependent variable.[10] However, land allocation appears to be

[10] With an extra pre-reform survey round one could correct for this problem using an Instrumental Variables Estimator, but that is not an option in our case given that we only have two survey rounds.

well known at farm-household and commune level. So we do not expect sizeable bias for this reason.

A second concern is that the process may not be homogeneous in that the initial land allocation may influence land re-allocation independently of the gains and losses from the initial administrative allocation. Imposing homogeneity when it does not hold will bias upward (downward) the OLS partial adjustment coefficient if there is convergence (divergence) at a given land deficit relative to the efficient allocation. By adding L_i^A as an additional regressor, we can test homogeneity. Again, any measurement error in L_i^A may induce some bias, which will tend towards showing convergence.

A third concern is that the efficient allocation of land may have changed over time. For example, demographic shocks will no doubt shift the consumption-efficient allocation. This can be thought of as measurement error in our estimate of the loss from the administrative allocation. We address this issue by adding controls for observed changes in household characteristics that are likely to influence the efficient allocation. Latent measurement error will leave some bias.

A final concern is that the local political economy may influence land re-allocation, as discussed in section 2. To deal with this concern we can postulate instead a solution, L_i^{R*}, such that the higher $\tau(L_i^{R*}, L_i^*)$, the higher the weight that a given household has in local decision making about land. This allows some households to acquire more land in the long run than implied by the efficient solution. Thus $\tau(L_i^{R*}, L_i^*)$ can be thought of a measure of the household's (market or non-market) power over land allocation. We assume that L_i^{R*} depends on assets (education and other types of land), connections (such as having a government job and being a long-standing resident) and possible discriminating variables (such as gender of head and ethnicity). We then augment the partial adjustment model for these household characteristics. Notice that the initial administrative allocation may itself be one such factor; if a higher initial administrative allocation gives one the power to acquire more land then we will see signs in the data of a divergent (non-stationary) process.

Combining these considerations, we shall also estimate a parametric model:

$$\rho_i = \alpha + \beta\tau_i + \gamma \ln L_i^A + \pi Z_i + \nu_i \tag{7}$$

in which Z_i denotes a vector of controls for other (market and non-market) factors influencing $\tau(L_i^{R*}, L_i^*)$. It is readily verified that the long-run solution to (7) (when $L^R = L^A = L_i^{R*}$) is:

$$L_i^{R*} = \phi^{-1}[\phi(L_i^*) + \frac{\alpha}{\beta} + \frac{\gamma}{\beta}\ln L_i^A + \frac{\pi}{\beta} Z_i + \frac{\nu_i}{\beta}] \tag{8}$$

We can also allow the partial regression coefficient of ρ_i on τ_i to vary between individuals according to their characteristics, by testing for appropriate interaction terms to equation (7).

In augmenting the unconditional partial adjustment model for these controls, we will not be able to cleanly separate "competitive" from "non- competitive" forces on land allocation. A non-zero element of the parameter vector π could reflect that characteristic's influence over how the competitive market allocation has changed over time or it could reflect its bearing on the ability of a household to distort the market in its favor, by exercising its (market or non-market) power. In this setting it is hard to imagine any household characteristic that could be unambiguously interpreted as one rather than the other. For example, finding a significant effect of gender or ethnicity is suggestive of a non-competitive force at work, but we cannot know in which market it operates; possibly the discrimination is in access to credit rather than land.

However, we will be able to see whether the controls reinforce or offset the adjustment process. We will say that the controls are "cooperant" ("noncooperant") with competitive market forces arising from inefficiencies in the initial administrative allocation if the unconditional adjustment coefficient (setting $\gamma = \pi = 0$) is found to be biased upward (downward).

4. Data

We use the household panel data from the 1992/93 and 1997/98 Vietnam Living Standard Surveys (VLSS). The first survey preceded the change in the land laws in 1993. These are nationally representative, high quality surveys with comprehensive and carefully collected data on a wide range of household characteristics including consumption expenditures, production and land holdings (World Bank 1995 and 2000). The surveys contain a balanced panel of 4308 households. We limit our sample to the 2559 rural farming households in the panel who had allocated annual agricultural land in 1993. The 1992/93 VLSS is self weighted so that expansion factors are not needed. Both surveys spanned 12 months.

Perennial, forest and water surface land have also been allocated to households. However, we focus on allocated annual agricultural land because of its importance in production and total area, and because its allocation began earlier and has progressed more rapidly than for other land types.[11] (Annual agricultural land is for annual crops such as rice or groundnuts.)

Annual agricultural land can be irrigated or non-irrigated. To facilitate the analysis we convert all allocated annual agricultural land into an allocated irrigated land equivalent amount for each household. Non-irrigated land amounts are weighted by the ratio of the coefficients on non-irrigated to that of irrigated land estimated from region-specific regressions of farm profits on allocated irrigated and non-irrigated annual land and all other land cultivated by households,

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[11] We will hereafter refer to allocated annual agricultural land simply as allocated land.

household characteristics and commune dummies. The weights are estimated using the 1992/93 VLSS and used to create the allocated irrigated land equivalents in both 1992/93 and 1997/98. [12]

A household's cultivated land can differ from its allocated land. Rural households typically have their own private residential land with its garden area. We consider this type of land as being a well-known and longstanding asset associated with each household and hence we control for it in our analysis. The rental market is thin. Rented-in land represented 6.2% of annual crop land in 1993 and 5.1% in 1998. A more active rental market has clearly not emerged since the reforms. Our impression is that rentals tend to be temporary arrangements, such as when a family worker is sick or temporarily absent. There is also a small amount of "auction land" that is effectively rented from the commune. (This accounted for 2.1% of all cultivated land in 1993, and 2.2% in 1998.) We do not control for land obtained though rental arrangements, given the possible endogeneity concerns.

The land situation has been evolving during the 1990s — reflecting changing official attitudes towards the market economy and the role of land, and consequent policy and legal reforms. This is apparent in the surveys. There were some changes in land categories and definitions between the 1992/93 and 1997/98 VLSS. Our aim here is to study changes in the allocated annual land amounts over time. Fortunately, this is straightforward. In 1993, our allocated land variable comprises the questionnaire categories "allocated" and "long-term-use" annual land. (Both categories refer to land allocated to households for long-term use. They differ only in that the allocation terms are slightly different with the first arrangement more common in the north and the second more so in the south.) By 1998, this distinction is no longer enforced. The 1997/98 VLSS refers to allocated land as either long-term-use or 'contract' land. The latter

[12] See Ravallion and van de Walle (2004) on construction of the allocated land equivalent.

is also allocated to households for long-term and stable use, but its land-use title is held by a state managed farm or enterprise rather than the household. This category of land was subsumed in either allocated or long-term use land in the 1992/93 survey. We consider this to be part of the allocated land category in 1998. Finally, in contrast to the 1992/93 VLSS where allocated annual land amounts include any area that was rented out, the latter is recorded separately in 1998 and so must be added in to determine the amount of the household's total allocated annual land.

The measure of consumption in 1992/93 (used to estimate the consumption-efficient land allocation) includes the value of consumption from own production, imputed housing expenditures and the use value of consumer durables (World Bank 1995). It also takes account of temporal price variation across the survey year as well as spatial price differentials. The determinants of initial (1993) consumption and land allocation were household demographics, the dependency ratio (1-ratio of working age household members to all members), disability incidence, age and age-squared of the head, education attainments, having a government job or a job in a state-owned enterprise, private land by type and land quality variables (Ravallion and van de Walle, 2004). As far as feasible, these variables were lagged five years, to better reflect circumstances at the time of the 1988 land law; this mainly affected the demographics and workforce data, though these were clearly key factors in land allocation. The regressions for initial consumption and land allocation are reported in Ravallion and van de Walle (2004).

Vietnam is commonly divided into seven regions that are relatively homogeneous. We estimate our regressions nationally as well as for the five regions for which there was sufficient data, namely the Northern Uplands, the Red River, North Coast (these three are in the north) and the Central Coast and Mekong Delta (the south). In addition, the augmented model includes a full set of commune dummy variables to capture geographic differences in prices and possibly institutional differences.

In our augmented model for land re-allocation we control for exogenous household level variables that describe the household's initial 1993 situation in terms of assets, connections and possible discriminating variables. These include the years of education of the head and of other adults; dummy variables for his/her religion (1 if the head practices the Christian or Buddhist religion, 0 otherwise), ethnicity (1 if the head belongs to an ethnic group other than the majority Kinh or relatively wealthy Chinese minority) and whether born locally; dummies for whether the household contains one or more handicapped adult members, members who work for the government or for a state owned enterprise, and whether the household is a recipient of social insurance fund transfers. The latter are given to war heroes or martyrs and their families; such households are often singled out for preferential policy treatment by the authorities and the fact of receiving the transfer is the only way of identifying them in our data. We run the model with and without the dummy variable for receipt of social fund transfers, given the possible endogeneity concerns. We also control for the household's private land (discussed above), whether it cultivates swidden land or not, and the share of its irrigated and non-irrigated land that is considered of good quality.

In addition, we include variables that capture exogenous changes in the household's characteristics that are likely to shift the consumption efficient allocation — namely the change in the number of disabled adult members, the change in the number of able bodied working age members, the number of new members aged between 8 and 99 in 1998, and whether an adult or elderly member died between the two surveys.

However, there are limits to how many control variables we can add to the partial adjustment model. We cannot include all the postulated determinants of initial consumption (the full vector *X* in equation 4) as well as initial land allocation since doing so would create a singularity (given that the log efficiency loss is linear in log initial land and *X*). We must thus impose exclusion restrictions. We follow common practice in panel data econometrics in relying on lagged values to help in identification. In our augmented model based on equation (7) the excluded variables from the model for initial consumption are the lagged values (lagged five years prior to 1993) for the demographics (notably household size and the dependency ratio), and the presence of a disabled adult in 1993. While these variables influence consumption, they are assumed to be irrelevant to the post 1993 land re-allocation conditional on the initial efficiency loss, initial land holding and other control variables.

Table 1 provides summary statistics. Notice that there was a net increase in total allocated land for the panel sample, reflecting new land brought under cultivation.

5. Results

Recall that in measuring land re-allocation and the initial efficiency loss we assume that $\phi(L) = (L^\eta - 1)/\eta$ where $\eta \in [0,1]$. To choose a value of η we regressed ρ_i on τ_i across the entire data set for alternative values of η at 0.1 intervals over the [0,1] interval. The best fit (measured by the t-ratio on the partial adjustment coefficient) was obtained at $\eta = 0$, which gave a partial adjustment coefficient for proportionate differences of 0.33 with a t-ratio of 9.8.[13] The coefficient for absolute differences ($\eta = 1$) was 0.17 and between the two, the t-ratio declined monotonically. So we chose the proportionate (log difference) specification in all further work.

However, this specification has the drawback that we lose some observations with zero land allocation in 1997/98 (since we cannot take the log of zero); this applies to less than 8% of the sample.[14] Later in this section we will study this sub-sample with zero allocated land in the second survey more closely, and test for sample selection bias. For the present discussion we confine attention to the proportionate case.

For the national sample, Figure 1 plots the proportionate changes (log differences) in land allocation against our measure of the initial loss relative to the efficient allocation, measured by $\ln(L_i^* / L_i^A)$. The empirical relationship suggests a tendency for land re-allocation to respond positively to the initial inefficiency in the administrative allocation. As already noted, the linear regression coefficient is 0.33, indicating that one third of the initial disparity between the administrative allocation and the market allocation was eliminated over this five year period. Figure 1 also gives the nonparametric regression function (using Cleveland's, 1979, local regression method). The slope is positive but less than unity throughout, though it is clear that $f(0) \neq 0$, reflecting an overall expansion in allocated annual land area over this period.

Figure 1 is suggestive of partial adjustment toward the market allocation, though still leaving two-thirds of the initial mean proportionate efficiency loss after five years. However, as noted in the previous section, there are a number of concerns about bias, which could go in either direction. One concern is that the relationship might not be homogeneous. On adding $\ln L_i^A$ to the regression of $\ln(L_i^R / L_i^A)$ on $\ln(L_i^* / L_i^A)$, we could convincingly reject the null hypothesis

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[13] All t-ratios in this paper are based on standard errors corrected for both heteroskedasticity and clustering.

[14] We also tried defining the proportionate difference as the percentage change rather than log difference, thus allowing us to keep these observations; the results were similar, though (again) the log difference specification gave a better fit.

implied by homogeneity. The regression coefficient on $\ln L_i^A$ was -0.287 (t-ratio of 8.05), while the partial adjustment coefficient fell to 0.217 (7.09).

Table 2 gives the estimated partial adjustment coefficients when various controls are added step-by-step (cumulatively). We give national results and a breakdown by region. Let us focus first on the national results. Consistently with Figure 1, all of our tests indicate a highly significant positive coefficient on the initial efficiency loss, implying that the land re-allocation process was in the direction of a more efficient allocation. However, as can be seen from Table 2, the partial adjustment coefficient falls to less than half the value implied by Figure 1 when all controls are added. This is the combined effect of both relaxing homogeneity and adding the controls including commune fixed effects. Of all these changes, relaxing homogeneity and adding commune effects does most of the work; with just these two changes, the partial adjustment coefficient falls to 0.155 (t=5.18), while adding the rest of the control variables only brings it down slightly more, to 0.131 (Table 2).

There are regional differences in the estimated adjustment coefficients, though the pattern of declining coefficients as controls are added is similar across regions. There is little sign of a difference between the north and the south; while the highest coefficient without controls is for the Northern Uplands, the south's Mekong Delta is the second highest.

Our results suggest that any non- competitive forces being picked up by our controls tended to be cooperant with competitive market forces, as captured by the adjustment coefficient to initial losses from the administrative allocation. This is evident from the fact that, on balance, controls that raise (lower) land allocation tend to be positively (negatively) correlated with the efficiency loss due to the initial allocation. The only exception is for the controls for demographic shocks, which tended to work in the opposite direction (as is evident in Table 2), though the effect on the partial adjustment coefficient is small.

In Table 3 we give the complete results for the most comprehensive model we estimated. For this we also added interaction effects between the initial loss variable and both initial land allocation and head's education, to allow the adjustment coefficient to vary within regions. The interaction effect with education was insignificant nationally and in most regions. However, we find a significant interaction effect between the initial loss relative to the efficient allocation and the initial land allocation. The speed of adjustment toward the efficient allocation was higher for those who started off with less land.

We find a number of other factors that influence land re-allocation. There is a highly significant effect of an increase over the time period in the number of persons of working age and new people joining the household. (We also tried dropping the latter variable given possible endogeneity concerns, but other results were affected little in the national model.) Households with male heads were also favored in the land re-allocation process. Having higher amounts of other types of land resulted in significantly higher access to allocated land.

There are some regional differences in the model with controls. The significant negative interaction effect (such that there is a higher adjustment coefficient for households with less land) is only found in the Mekong. Whether this is a market response is unclear; it could also reflect the efforts of local officials in the Mekong to avoid rising landlessness (Section 2).

The impacts of demographic and labor force changes appear to be generally stronger in the northern provinces. This is also where local authorities are more likely to enforce periodic land re-allocations. Being from an ethnic minority household helped increase annual land holdings in the north, and (especially) the Central Coastal region, while it tended to reduce holdings in the Mekong Delta; note, however, that the ethnic groups are not the same in these two regions. Ethnic effects also become significant and positive in the Northern Uplands and North Coast regions when we omit the number of new household members in 1998. Having a member who works for an SOE has a pronounced negative impact on annual land changes in the Northern Uplands and the Central Coast, though it has no impact elsewhere. In both the Northern Uplands and Central Coast regions a higher share of good quality irrigated land reduced the land re-allocation over time.[15] The tendency to favor male heads of household is strongest in the north.

We also tested for effects of the initial efficiency of land allocation on the probability of becoming landless (in terms of allocated annual land). Table 4 gives the proportion of the 1997/98 sample that had no allocated land classified by the estimated initial loss relative to the efficient allocation in 1992/93. The higher the loss relative to the efficient allocation the higher the probability of having no allocated land in 1997/98.

We also estimated probits for landlessness using the same regressors as in Table 3. We did this for both disposal of allocated annual land and disposal of all cultivated land. Virtually the only significant predictors in any of these regressions was the proportionate efficiency loss, which had a significant positive coefficient in most cases, and geographic dummy variables. Becoming landless was more likely for households who had too little land relative to the efficient allocation, and it was more likely in the south than in the north.

Our results are suggestive of a "land polarization" process among those who started off with too little land relative to the efficient allocation. The bulk of these households "traded up,"

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[15] We tested a dummy for being a social fund transfer recipient, one of the few ways to identify households that may be treated preferentially by local authorities. This was insignificant in the national model and all regions except the North Coast where it had a positive effect.

acquiring more land in the more market-oriented economy. However, a minority simply disposed of their allocated land. The results in Table 4 are suggestive of an interpretation in which a subset of those households who started out with too little land (relative to the efficient allocation) simply "cashed in," possibly to take up other non-farm activities or pay off debts.

The difference in behavior of those households who disposed of their allocated land raises a concern about the possibility of sample selection bias in our main regressions for land re-allocation.[16] In fact there are two possible sources of such bias. The first stems from the fact that our preferred specification for the functional form entailed that some observations had to be dropped; the second is panel attrition, in that some of the original random sample could not be interviewed in the second survey for various reasons (they had left their original address or they chose not to participate again). Motivated by the approach to testing for panel attrition bias in Fitzgerald, Gottschalk and Moffitt (1998), we tested for both sources of bias using initial land allocation as the auxiliary endogenous variable in a probit for whether a household dropped out of the sample (for either reason), with controls for all other observable exogenous characteristics in the baseline survey. (We used the same set of controls as in our model of land re-allocation.) The initial land allocation variable was statistically insignificant (at the 10% level) nationally and for all regions, suggesting that there is little or no bias due to sample selection in our regressions for land re-allocation.

[16] It might be conjectured that this explains why we get a better fit using the log difference specification; since the observations that disposed of their allocated land behaved very differently to differences in the initial inefficiency of their allocation, dropping these (because one cannot take the log of zero) improved the fit. However, we got a better fit with the log specification across the same (truncated) sample when compared to other values of η (tested at 0.1 intervals over the [0,1] interval).

6. Conclusions

The standard policy prescription for transforming a socialist command economy into a market economy is to privatize productive assets and then change the law to permit free transactions in those assets. We have put this model to the test in the context of Vietnam's agrarian transition.

We find some support for the standard model during a period that included major liberalizing reforms to land laws. There are signs that land allocation responded to the inefficiencies of the initial administrative assignment at de-collectivization. Households who started with an inefficiently low (high) amount of crop land under the administrative assignment tended to increase (decrease) their holdings over time, through the process of re-allocation allowed under the new land laws. The partial adjustment coefficient was about 1/3 in the aggregate, meaning that one third of the initial proportionate gap between the actual allocation and the efficient allocation was eliminated within five years.

We find an appreciably lower adjustment coefficient when we relax the standard homogeneity assumption in partial adjustment models (whereby the initial allocation does not influence the change in land allocation independently of the initial loss relative to the market allocation). At a given land deficit or surplus relative to the efficient allocation, households who started with the least crop land under the administrative assignment tended to see the largest increase in holdings during the transition. The speed of adjustment to inefficiencies in the administrative allocation also tended to be higher for those who started with less land. In other words, the transition process favored the "land-poor."

The adjustment coefficient falls when we add controls for commune effects, demographic shocks and possible non-market factors influencing land allocation. The process favored households with long-term roots in the community, with male heads, better education and with more non-allocated land. We find that these controls tend to be cooperant with market forces, in that they are jointly positively correlated with land re-allocation and the efficiency losses from the initial administrative allocation.

This is not what one would expect if the controls reflected strong non- competitive forces working against efficient land reallocation. The seemingly slow response to the initial inefficiencies of the administrative allocation does not appear to stem from countervailing non-competitive forces, but appears to be inherent to the workings of the market process in this setting.

Table 1: Variable definitions and summary statistics

	Mean	st.dev.
Log difference in allocated irrigated land equivalent (m^2)	0.142	0.66
Proportional efficiency loss (log efficient allocation minus log actual in 1993)	-0.016	0.78
Religion: 1 if h'hold head is Buddhist or Christian (0 if other, animist or none)	0.307	0.46
Ethnic: 1 if h'hold head is of ethnicity other than majority Kinh or Chinese	0.121	0.33
Local born: 1 if head is born locally	0.861	0.35
Gender of household head (male=1)	0.791	0.41
Labor age adult member is handicapped	0.007	0.09
SOE: member has primary or secondary occupation in State owned enterprise	0.018	0.14
Gov't job: member works for gov't in primary/ secondary occupation or retired from gov't (professional codes 20 and 21)	0.059	0.25
Social subsidy: dummy var. for receipt of gov't transfers to war heroes, martyrs, disabled etc	0.103	0.30
Household head's years of education	6.107	3.83
Other h'hold adults' years of education	10.648	9.22
H'hold's private irrigated land (m^2)	158.853	658.68
H'hold's private non-irrigated land (m^2)	228.824	955.31
H'hold's private perennial land (m^2)	349.057	1492.13
H'hold's private water surface land (m^2)	55.913	478.74
H'hold cultivates swidden land=1	0.108	0.31
Share of good irrigated land	0.304	0.39
Share of good non-irrigated land	0.374	0.46
No. >=16 in 1993 who died by 1998	0.109	0.33
No. >=50 in 1993 who died by 1998	0.089	0.30
Change in number of disabled adults 1993-98	-0.004	0.15
Change in no. of able bodied working age members 1993-98	-0.138	1.19
H'hold has new individual aged 8-99 in 1998	0216	0.60

Source: 1992/93 and 1997/98 Viet Nam Living Standards Surveys. 2559 observations except for the change in log allocated land for which n=2361.

Table 2: Effects of adding controls on the partial adjustment coefficients by region

	Northern Uplands	Red River	North Coast	Central Coast	Mekong Delta	Full Sample
No controls	0.476	0.294	0.306	0.172	0.350	0.328
	(5.97)	(6.81)	(3.35)	(2.17)	(4.51)	(9.82)
Adding initial land allocation	0.170	0.094	0.129	0.025	0.221	0.218
	(1.61)	(2.67)	(1.24)	(0.37)	(3.06)	(7.09)
Adding commune effects	0.205	0.123	0.132	0.079	0.171	0.155
	(3.96)	(2.98)	(1.52)	(1.32)	(1.62)	(5.18)
Adding controls for demographic shocks	0.255	0.150	0.175	0.074	0.215	0.182
	(4.89)	(4.02)	(2.24)	(1.15)	(2.20)	(6.46)
Adding controls for connections and assets	0.268	0.071	0.173	0.069	0.074	0.131
	(4.54)	(1.39)	(1.68)	(1.16)	(0.73)	(4.09)
No. observations	432	790	459	269	308	2361

Note: The table gives regression coefficients of the change in log annual land allocation on the estimated proportionate loss from the initial administrative allocation relative to the counter-factual market allocation. The regressions are cumulative in that as controls are added the previous controls are kept in.

Table 3: Determinants of changes in allocated annual agricultural land

	Northern Uplands	**Red River**	**North Coast**	**Central Coast**	**Mekong Delta**	**Full sample**
Proportional loss from	0.433	0.197	0.501	0.230	1.494	0.700
admin. allocation	(2.65)	(0.52)	(1.09)	(0.67)	(2.90)	(4.51)
Log initial land	-0.481	-0.434	-0.298	-0.495	-0.394	-0.405
allocation	(7.20)	(6.32)	(3.47)	(10.04)	(4.01)	(11.78)
Interaction of loss with	-0.024	-0.017	-0.047	-0.022	-0.168	-0.077
initial land	(1.06)	(0.34)	(0.84)	(0.52)	(3.02)	(3.87)
Adult member died	0.096	0.110	0.043	-0.059	0.170	0.043
1993-98	(0.52)	(1.22)	(0.18)	(0.53)	(1.07)	(0.53)
Elderly member died	-0.150	-0.118	-0.034	-0.143	-0.162	-0.080
1993-98	(0.67)	(1.18)	(0.14)	(0.96)	(0.99)	(0.88)
Change in no. disabled	0.204	0.240	0.122	0.043	-0.008	0.119
1993-98	(2.15)	(1.66)	(1.77)	(0.43)	(0.04)	(2.03)
Change in no. of able	0.119	0.150	0.119	0.052	0.05	0.100
bodied members	(5.08)	(8.70)	(5.56)	(1.44)	(1.72)	(8.92)
New member 8-99	0.113	0.189	0.111	0.050	0.205	0.124
1993-98	(2.20)	(4.59)	(1.73)	(0.94)	(3.74)	(5.00)
Religion	0.151	-0.049	0.020	-0.054	0.126	0.005
	(2.13)	(1.12)	(0.20)	(0.45)	(2.61)	(0.16)
Ethnicity	0.254	-0.128	0.089	1.014	-0.288	0.096
	(2.06)	(3.40)	(0.75)	(14.57)	(1.44)	(0.93)
Born locally	0.159	0.018	0.160	0.178	-0.026	0.093
	(1.71)	(0.25)	(1.36)	(2.15)	(0.22)	(2.13)
Gender of head	0.121	0.121	0.097	0.091	0.068	0.123
(male=1)	(3.93)	(2.73)	(1.61)	(1.27)	(0.64)	(4.35)
Government job	-0.142	-0.060	-0.142	-0.171	0.124	-0.090
	(1.01)	(0.75)	(1.58)	(0.86)	(0.94)	(1.56)
SOE job	-0.462	0.104	-0.087	-0.216	0.174	0.036
	(4.19)	(0.56)	(0.37)	(2.06)	(1.05)	(0.28)
Education of head	-0.006	0.011	-0.000	-0.001	0.028	0.006
	(0.78)	(2.48)	(0.05)	(0.18)	(1.40)	(1.58)
Education of other	0.004	0.004	-0.001	0.007	0.009	0.004
adults	(1.52)	(1.60)	(0.20)	(2.79)	(2.09)	(2.18)
Share of good quality	-0.032	-0.047	0.032	-0.058	0.005	-0.009
non-irrigated land	(0.38)	(0.81)	(0.50)	(0.63)	(0.06)	(0.27)
Share of good quality	-0.256	-0.001	-0.088	0.118	0.271	-0.063
irrigated land	(2.21)	(0.01)	(0.84)	(1.59)	(1.94)	(1.23)
Private irrigated x 10^3	0.051	0.249	0.275	-0.020	0.051	0.058
	(0.61)	(1.57)	(1.92)	(0.18)	(2.56)	(2.44)
Private non-irrigated	0.077	0.111	0.195	0.056	0.080	0.042
x 10^3	(0.78)	(4.04)	(2.06)	(0.92)	(7.34)	(1.88)
Private perennial x 10^3	-0.031	0.015	-0.139	0.092	0.044	0.024
	(0.063)	(0.016)	(1.29)	(1.11)	(2.00)	(2.04)
Private water surface x	0.334	0.027	-0.043	--	0.041	0.059
10^3	(2.72)	(0.52)	(0.31)		(5.45)	(3.86)
Swidden land dummy	-0.149	0.266	0.242	0.122	0.171	0.064
variable	(2.37)	(6.75)	(1.85)	(0.88)	(3.09)	(0.94)

Commune dummy variables	Yes	Yes	Yes	Yes	Yes	Yes
Constant	2.938	2.793	2.067	4.235	2.165	2.615
	(6.97)	(5.57)	(3.68)	(8.68)	(2.56)	(7.82)
R^2	0.631	0.461	0.435	0.548	0.438	0.490
RMSE	0.472	0.390	0.454	0.420	0.610	0.483
No. observations	432	790	459	269	308	2361

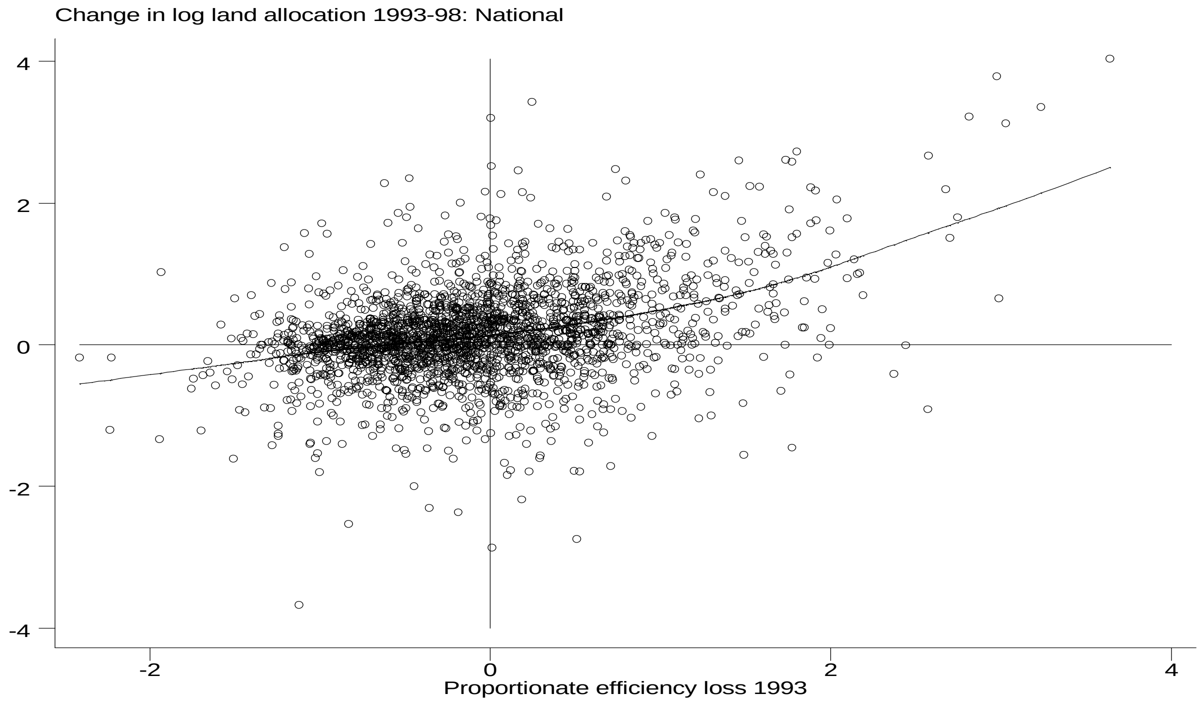
Note: The dependent variable is the change in log annual agricultural allocated land between 1993 and 1998. Absolute t-ratios in parentheses are based on standard errors corrected for heteroskedasticity and clustering. Unless otherwise noted, all variables are initial 1993 values.

Table 4: Disposal of allocated land

% landless in 1998/99		% landless
Quintiles of households ranked by the efficient allocation minus the from administrative allocation, 1992/93	1 (Gained relative to the efficient allocation)	4.6
	2	2.6
	3	5.9
	4	10.7
	5 (Lost relative to the efficient allocation)	16.4
		7.7

Note: % of households having no allocated annual agricultural land in 1997/98; total number of sampled households is 2559.

Figure 1: Proportionate land re-allocation 1993-98 against the proportionate loss from the administrative allocation in 1993



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