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Can Labor Market Imperfections Explain Changes in the Inverse Farm Size–Productivity Relationship?

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Abstract

A large national farm panel from India covering a quarter century (1982, 1999, and 2008) is used to show that the inverse farm size-yield relationship weakened significantly over time, despite an increase in the dispersion of farm sizes. Key reasons are substitution of capital for labor in response to nonagricultural labor demand. Family labor was more efficient than hired labor in 1982–99, but not in 1999–2008. In line with labor market imperfections as a key factor, separability of labor supply and demand decisions cannot be rejected in the second period, except in villages with very low nonagricultural labor demand.

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Can Labor Market Imperfections Explain Changes in the Inverse Farm Size–Productivity Relationship?
Longitudinal Evidence from Rural India

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1. Introduction

The existence of an inverse relationship between farm size and output per unit of land is among the most well established empirical regularities in agricultural production economics. First identified descriptively in Russia (Chayanov 1926), then confirmed by Indian farm management studies (Bardhan 1973; Srinivasan 1972), and subsequently supported by a much larger set of countries (Berry and Cline 1979), this relationship proved robust to numerous challenges. These include but are not limited to better accounting for land quality and soil characteristics, use of profits instead of output, increased precision of plot area estimates, and more sophisticated panel estimation techniques.

Barring measurement error, the most likely explanation for this phenomenon is market imperfections. However, which markets are affected, how amenable they are to change, and what this implies for policy have not been established clearly in the existing literature. To take one example, the inverse relationship has been used to recommend redistributive land reform as a way to improve efficiency and equity (Lipton 2009), but how to implement this in a way that avoids negative side effects is far from clear. Recognition that rising non-agricultural wages and new technology will affect factor price ratios, supervision requirements, and the presence and extent of market imperfections that might have caused the relationship in the first place has led many to reassess such policy advice to also account for political viability.

Also, changes in factor price ratios, which increase the attractiveness of substituting capital for labor in agricultural production, together with continued subdivision in the context of generational change, may limit the scope for mechanization and thus contribute to a reversal of the inverse relationship, making some farms (or, more accurately, plots) too small for efficient cultivation (Foster and Rosenzweig 2010). Recent innovations in crop breeding, tillage, and information technology also may make agricultural production more knowledge-intensive, thus eliminating part of small farmers' traditional advantage. Most studies use cross-sectional evidence or short panels; longitudinal data at the household level could improve understanding of the relationship between farm size and productivity by allowing researchers to test how that relationship evolves in response to changes in technology and market functioning.

To provide such insights, this paper draws on panel data from close to 5,000 farm households in 17 Indian states over a 25-year period (1982–2008). We use panel regressions for the early (1982–1999) and late

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(1999–2008) periods—in both cases, instrumenting size of cultivated area in order to explore changes in the inverse relationship over time and to test different explanations of this relationship with particular focus on the role of labor market imperfections. This study thus complements existing studies that explored the relationship from a cross-sectional perspective or with much shorter panels. Over the period covered by the data, India's economy experienced dramatic changes, the magnitude of which differed considerably across states; this finding provides a source of identification not available to cross-sections.

The results from nonparametric and parametric regressions point towards an attenuation of the inverse farm size-productivity relationship over the period studied, consistent with the findings in other countries. To appreciate mechanisms that may underpin such changes, we note that higher wage levels in the 2000s were likely to have increased demand for fertilizer and machinery as a substitute for labor. It is thus expected that production would have become more capital intensive, possibly weakening the supervision cost advantage of family labor. Indeed, panel labor demand functions for the first and second periods support this, suggesting that the family–to–hired labor ratio is significant and positive in the first period, but insignificant in the second, which is consistent with the notion that equality of productivity between family and hired labor can no longer be rejected in the second period.

If supervision matters less with higher capital intensity of production, we would expect the coefficient on the family–to–total labor ratio in a labor demand function to become insignificant, something indeed observed in the data. Alternatively if, with improved functioning of agricultural and non-agricultural labor markets, own and hired labor are fully substitutable, then households' labor endowment should not affect labor use in agriculture. Although separability between households' endowment and labor demand is indeed rejected for the first period, pointing toward labor market imperfections, it is not rejected for the second period, suggesting that the functioning of labor markets improved over time. Consistent with this notion, we find that, while separability in the second period is not rejected for the entire sample, it is rejected in villages with inactive non-agricultural labor markets. This could be an explanation for the persistence of the inverse relationship between farm size and productivity, albeit less pronounced, in period two.

The finding of improved labor market functioning provides a coherent explanation of the observed weakening of the negative farm size and productivity relationship over time in India; it suggests that efforts to improve rural factor market functioning as a policy response will be appropriate in virtually all conditions. By contrast, policies of redistributive land reforms will be a desirable response only if implementation is possible as, without this, they may create further impediments to the functioning of rural factor markets.

The paper is structured as follows. Section two reviews literature and presents data and descriptive statistics, illustrating the far-reaching changes experienced by India over the period considered. Section three presents results regarding the inverse size-productivity relationship and factor demand equations. Section four

focuses on explanations in terms of labor market imperfections by testing the relevance of family labor and separability of labor supply and demand. Section five concludes with implications for policy and research.

2. Motivation and relevance

This section draws on conceptual and empirical literature to document strong empirical evidence in support of a negative farm size–productivity relationship and to discuss potential channels underpinning it. If an inverse relationship between productivity and farm size can be attributed to labor market imperfections, one would expect the relationship to weaken as better labor market functioning and increased demand for non-agricultural labor lead to choice of more capital-intensive modalities of agricultural production. Longitudinal data spanning a long enough period should uncover such patterns.

2.1 The farm size-productivity relationship

A negative relation between farm size and output per unit of land in un-mechanized agriculture has been confirmed many times in the literature (Eastwood, Lipton, and Newell2010). It was first noted in Russia (Chayanov 1926), in Indian farm management studies (Bardhan 1973; Srinivasan 1972), and in a host of other countries using simple ordinary least square (OLS) regressions (Berry and Cline 1979) that likely failed to fully control for land quality (Bhalla and Roy 1988). Yet, even when these shortcomings are addressed, the inverse relationship has remained surprisingly robust. Correcting for measurement errors may weaken the relationship, but it does not invalidate it (Lamb 2003). In fact, use of panel estimation techniques (Assuncao and Braido 2007) or careful controls for soil quality affected the magnitude but not the existence of the relationship (Barrett, Bellemare, and Hou 2010). Although biased estimates of area cultivated by farmers may reinforce the inverse relationship, it persists even when objective, GPS-based measures of area cultivated are used (Calogero, Savastano, and Zezza 2011). Use of profits rather than output per hectare in India has the same result—namely, the relation weakens but does not disappear (Rosenzweig and Binswanger 1993). The relationship is also not limited to South Asia; a recent study finds that in Mexico, large farms have a lower value of output per hectare than small farms and produce further from the efficiency frontier (Kagin, Taylor, and Yunez-Naude 2016).

Conceptually, imperfections in land, labor, and credit or insurance markets affect the relationship between farm size and productivity. If, as suggested by most empirical studies, the agricultural production function is characterized by constant returns to scale (Deininger, Nizalov, and Singh 2013), imperfections in two markets—most often those for land and labor—will generate a systematic relationship between size of cultivated area and output (Feder 1985). Family labor has long been shown to resolve informational asymmetries in rural labor markets (Bharadwaj 2015). If hired labor is used, owner-operators, who are residual claimants to profit, will be more likely to exert effort than will wage workers, who, in light of the

dispersion of agricultural production processes over space and time, require supervision (Frisvold 1994). This generates a negative relationship between productivity and size as the ratio of hired to family labor increases with farm size. Owner operators' advantage could be enhanced by their knowledge of local soil or climatic conditions which may have been accumulated over generations (Rosenzweig and Wolpin 1985).

Small farmers' advantages in labor supervision, knowledge, and the ability to adjust to labor demand can be offset by their difficulty in accessing capital and insurance, which arise from the high transaction cost of providing formal credit in rural markets, possibly exacerbated by the difficulty of small farms to provide collateral. Frictions created by fixed costs of participating in labor or land markets could motivate small farmers who are unable to rent to rationally apply family labor to cultivate their fixed land endowment more intensively than would be the case in a situation where such markets function perfectly. An inverse relationship can also emerge if labor and credit market imperfections are combined with a fixed cost element for production (Eswaran and Kotwal 1986) or if there is heterogeneity in farmers' skills in the presence of credit market imperfections (Assuncao and Ghatak 2003). Land and insurance market imperfections can prompt small farmers who are net buyers of food to use family labor more intensively in an attempt to reduce potentially adverse effects of price fluctuations (Barrett 1996). The lumpiness of certain inputs (for example, machinery, draft animals and management skills) together with the advantages in getting access to working capital or their capacity to diffuse risk may, in practice, lead the relationship between farm size and productivity to be U-shaped (Heltberg 1998). As a result, with few exceptions,[2] agricultural production continues to rely on owner-operated firms (Allen and Lueck 1998, Deininger and Feder 2001).[3]

Although supervision constraints can account for the inverse relationship if hired labor is used, off-farm labor market rationing that constrains opportunities for smallholders participation can provide an alternative explanation (Benjamin 1992). If this is a binding constraint, farmers with more abundant labor endowment relative to cultivated land would apply more labor per unit of land, leading to a negative relationship between farm size and productivity. Such imperfections, while often taken as exogenously given in cross-sectional studies, will vary with time and the way in which family labor is valued.[4]

Although labor market imperfections will strengthen the negative relationship, capital market imperfections are likely to result in the opposite outcome. Lumpiness of certain technologies such as bullocks or tractors

[2] A well-known exception to the advantages of owner-operated units of production over those relying on wage labor is in perishable plantation crops. In such crops, economies of scale in processing may be transmitted to the production stage (Binswanger and Rosenzweig 1986), and employment is often year-round, so that the optimum size of a unit is determined by the factory's processing capacity.

[3] As of the end of2009, only seven publicly listed farming companies existed worldwide—three in South America and four in Ukraine and the Russian Federation (Deininger *et al.* 2011). This contrasts with processing, input industries, and sometimes output markets, all of which are characterized by large fixed costs (such as for research and development or processing) that give rise to economies of scale and often a highly concentrated industry structure (Deininger and Byerlee 2012a).

[4] In Rwanda, a very strong negative relationship between farm size and productivity weakens or is completely eliminated if family labor is valued at shadow or market wages, respectively (Ali and Deininger 2015).

that cannot easily be addressed via factor markets may attenuate small farmers' advantage. Rising wages may make substitution of capital for labor desirable even without changes in technology (Yamauchi 2016). Technical change, via innovations in crop breeding, tillage, and information technology (Deininger, Nizalov, and Singh 2013) may reduce supervision requirements further.[5] Information technology may improve smallholders' competitive position at the marketing stage (Aker and Fafchamps 2014) more intensive use of technology at the production stage will require capital. If, for example, the transaction cost of providing formal credit in rural markets—which could possibly be exacerbated by constraints on those markets' ability to post collateral due to lack of formal documentation—may cause small farmers to have difficulty in accessing capital; thus, the inverse relation may actually be reversed. In fact, it has been shown that in India, large benefits may be forgone due to land market imperfections that make it impossible to assemble parcel sizes large enough to allow mechanization (Foster and Rosenzweig 2010).

The above suggests that magnitude and possibly direction of the relation between farm size and productivity will vary in response to exogenous factors including non-farm labor demand, availability of technology, and the functioning of capital markets. These factors, together with policy, will affect factor prices, the optimum ratio of capital to labor used in production, supervision needs, and the extent of market imperfections that have traditionally given rise to the inverse farm size-productivity relationship (Otsuka, Liu and Yamauchi 2013). This finding is illustrated by the differential effect of the increased importance of knowledge and capital in the wake of the green revolution (Deolalikar 1981), which weakened the farm size-productivity relationship in districts suited to new technology but left it intact, with small producers most efficient, where traditional methods prevailed.

2.2 Data and descriptive statistics

To empirically address this issue, we use data from the 1982, 1999 and 2008 rounds of the Additional Rural Incomes Survey & Rural Economic & demographic survey (ARIS/REDS), a panel survey covering 242 villages in 17 Indian states conducted by India's National Council for Applied Economic Research (NCAER).[6] Each round's household schedule provides data on demographics, assets, income, consumption, and economic activities at the household level, as well as detailed information on labor and non-labor inputs used in agriculture and outputs from agricultural production. In case of splits,[7] all of a household's

[5] Pest-resistant and herbicide-tolerant varieties facilitated broad adoption of zero tillage and, by reducing the number of steps in the production process as well as the labor intensity of cultivation, allowed management of larger areas. The ability to have machinery operations guided by GPS technology reduces skill requirements and allows remote supervision. The scope for substituting crop and pest models and remotely sensed information on field conditions for personal observation also reduces the advantage of local knowledge (Deininger and Byerlee 2012b).

[6] The sample included three strata: (i) districts that were part of the Intensive Agricultural District Program (IADP), an extension and input provision program placed in areas thought to have high potential for crop productivity growth; (ii) districts covered by the Intensive Agricultural Area Program, and (iii) all other districts.

[7] The 1999 round covered a total of 7,474 households, including all original households and their split-offs. The 2008 round (for the 2007–08 season) has a sample size of 8,659 households from 242 villages and includes all households surveyed in 1999, the split-off households residing within these villages, and a set of randomly selected new households.

successors in the same village were interviewed. We limited our analysis sample to dynasties that were interviewed in all three rounds.

Over the 25-year period covered by the data, India's economy experienced far-reaching changes. Although growth rates remained well below 5% in the 1980s, liberalization in the early 1990s resulted in much stronger economic growth, largely driven by the non-farm economy. As a result, from 1980 to 2010, real per capita gross domestic product (in 2000 US dollars) grew from $225 to $786 and the share of the population in agriculture decreased from 63% to 48%, reaching a point at which the non-farm sector employs close to 40% of the country's rural workforce (Himanshu *et al.* 2013). Although land market imperfections impact non-agricultural productivity (Duranton *et al.* 2015a), functioning of credit markets improved markedly since the 1990s (Duranton *et al.* 2015b). From 2006 to2008, the national rural employment guarantee scheme contributed to a significant increase in rural labor demand, which increased reservation wages (Imbert and Papp 2015a), reduced seasonal migration (Imbert and Papp 2015b), and resulted in some crop diversification. Inheritance-induced subdivision led to a significant decrease of plot size, below the point at which mechanization is possible, giving rise to inefficient patterns of cultivation (Foster and Rosenzweig 2010).

Table 1 includes sample means for key variables in each survey round. Although mean household size dropped from 7.4 in 1982 to 5.7 in 2008, the head's level of education more than doubled from 2.2 to 5.0 years. As a result of population growth, mean land endowments declined markedly, from 8.6 to 4.5 acres. At the same time, real wage rates more than tripled, from 15 (16) rupees per day in 1982 to 48 (68) rupees per day in 2008 for agricultural (non-agricultural) work. Intensity of agricultural labor use declined from 81 days per acre in 1982 to 51 days per acre in 2008, accompanied by an almost ten-fold increase in per acre expenses on machinery hire from 115 rupees to 989 rupees. Between 1982 and 2008, the share of households that hired machine services or owned a machine themselves increased from 51% (7%) to 74% (32%), whereas the share of households that neither owned nor hired tractors decreased from 49% to 26%. The data also point towards a significant increase in off-farm labor opportunities, with the share of farmers who worked in non-farm wage or self-employment increasing from low levels (6% each, respectively) to 20% and 18% and the share of income from agriculture decreasing from 84% in 192 to 61% in 2008.

Table 2 presents the same variables by size of cultivated land area for small (less than 2.0 acres), medium (2.0–6.0 acres), and large (more than 6.0 acres) farms. Although it is not surprising to find that large farmers' income was about triple that of small farmers throughout, large farms also obtained a much higher share of income (94% in 1982, 77% in 1999, and 81% in 2008) from agriculture in all three survey years. Household size dropped over time for all farm size classes, with declines from 8.4 in 1982 to 6.6 in 2008 for large, 6.8 to 5.7 for medium, and 5.9 to 5.3 for small farms. The share of agricultural income in total income declined

across all farm sizes, with 51% in 1982 to 34% in 2008 for small, and 72% to 58% for medium farms, the decline was relatively bigger for the small and medium farms but changes only little (from 94% to 81%) for the large ones.

Possibly in response to higher wages, farmers of all sizes shifted from labor-intensive to more capital-intensive modes of production. Although this change affected all farm sizes, it was most pronounced for large farms: from 1982 to 2008, small, medium, and large farms decreased labor days per acre by 48%, 54%, and 64%, respectively. The share of farms owning machinery, however, increased drastically over the survey periods with rates of increase more pronounced for larger farm sizes. Compared to few small and medium farms (1% and 3%, respectively) owning machinery in 1982, 18% and 38%, respectively, did so in 2008; machinery ownership for large farms increased from 13% in 1982 to 33% in 1999 and then to 52% in 2008. A similar increase is observed in the rate of renting machinery, although the rate of increase is more similar across farm size groups.

3. Changes in the farm size-productivity relationship and factor demand over time

Non-parametric and parametric regressions suggest that the inverse relationship between farm size and productivity was attenuated over time to the point at which it had almost disappeared at the end of the 1992–2008 period. Estimates of factor demand suggest that higher wages precipitated a substitution of machinery for labor via more capital intensive production technology, which is consistent with the finding that family labor was more efficient than hired labor in the 1982-1999 but not in the 1999-2008 period. Tests for separability between labor supply and labor demand further support the attenuation of labor market imperfection over time as one of the factors contributing to a weakening of the inverse farm size-productivity relationship.

3.1 Evidence on evolution of the relationship

To illustrate the issue at hand, Figure 1 plots results from locally weighted polynomial regressions of total monetary yield against farm size for 1982, 1999, and 2008. The 95% confidence intervals are also displayed. Between 1982 and 1999, farm size decreased significantly as a large number of very small farms emerged, presumably due to inheritance. This change led to a clear widening of the farm size distribution and a marked decline of mean holding size. Between 1982 and 1999, productivity increased markedly, though with little impact on the relation between farm size and productivity (the figure's slope). In the second period (1999–2008), average productivity changed much less, and the relationship flatted out, largely due to differential performance by farms smaller than and larger than 1 hectare with productivity declining (though insignificant at 5%) for the former but increasing (and significant at 5%) for the latter.

To appreciate the underlying factors, we denote households by *i*, communities by *j*, and time by *t*, to estimate a standard production function of the form

$$\ln y_{ijt} = \alpha_{\mathrm{i}} + \alpha_2 \ln A_{ijt} + \alpha_3 H_{ijt} + \alpha_4 D_t + \varepsilon_{\mathrm{ijt}} , \qquad (1)$$

where y_{ijt} is value of crop output,[8] A_{ijt} is total cropped area including land quality indicators; H_{ijt} is a vector of household characteristics, D_t is a year year dummy; α_{i} is household fixed effects and $\varepsilon_{\mathrm{ijt}}$ is a random error term. We estimate the model for the early (1982–1999) and the later period (1999–2008) separately and are interested in α_2, the estimated elasticity of yield with respect to cultivated area so that values of less (more) than 1 imply a negative (positive) farm size-productivity relationship. If, for example, due to improved functioning of other markets, the magnitude of this relationship decreased over time, the estimate of α_2 in the second period would be larger than in the early one.

For Example, due to unobserved production shocks that affect cultivated area, there is a possibility that $E(\ln A_{ijt}\varepsilon_{ijt}) \neq 0$, OLS could yield biased and inconsistent estimates of α_2. To address potential endogeneity of cultivated area, we use an instrumental variable (IV) strategy in which the sizes of inherited land and assets serve as instruments, relying on the assumption of production shocks being independently distributed over time (Foster and Rosenzweig 2002). For 1999–2008, we also use lagged household characteristics (which are not available in 1982), as instruments. Over-identification and weak instrument tests are used to check whether this is met in practice, and standard errors are adjusted for clustering at the village level throughout.

Table 3 reports results of estimating equation (2) for the 1982–1999 (cols. 1&2) and 1999–2008 periods (columns 3 and 4), in both cases without (columns. 1 and 3) and with (columns. 2 and 4) instrumentation. Because instruments easily pass relevant tests in both cases,[9] we focus discussion on instrumented estimates in columns 2 and 4. Three results stand out. Most important, although a negative farm size–productivity prevailed in both periods, estimated coefficients of 0.734 in the initial and 0.946 in the last period suggest that the magnitude of the relationship decreased markedly over time. In addition, the coefficient on the share of irrigated land is positive and significant in both periods although its magnitude is slightly smaller in the second period. Finally, although some household characteristics, such as asset endowments, education, and composition, are significant determinants of production in the first period, their significance disappears in the second period, possibly due to improved functioning of rural factor markets, an issue to which we will return below.

[8] The total output value is deflated by output price index using 1999 as a base year. The output price index is generated as a weighted average of prices of all types of outputs in the sample using the quantities as weights.

[9] The Hansen J over-identification test (p-value=0.18 for the early panel period and 0.28 for the later period) and weak-IV tests (p-value = 0 from both the Anderson-Rubin Wald test and the Stock-Wright Lagrange Multiplier test in both panel periods).

3.2 Determinants of factor demand

To explore whether rising labor and other input costs may have led to substitution between different types of inputs, we estimate input demand equations for key inputs, including machinery, fertilizer, and hired or family labor. Denoting households by i, villages by j, and time by t, input demand equations take the form

$$lnX_{ijt} = \alpha_i + \beta_1 lnA_{ijt} + \beta_2 lnw_{jt} + \beta_3 lnP_{jt} + \beta_4 H_{ijt} + \beta_5 D_t + \varepsilon_{ijt} \qquad (2)$$

where X_{ijt} is the value (or, for labor, quantity) of input use; A_{ijt} is the total crop area; w_{jt} is the agricultural and non-agricultural wage rate; P_{jt} is a vector of prices for output, machine rental, or fertilizer; H_{ijt} is a vector of household-level controls, such as the head's gender, age, and education, and number of members aged 14 to 64 years; D_t is a vector of year dummies; and α_i is a household fixed effect. We are particularly interested in β_2, the estimated elasticity of demand for different inputs with respect to wages.

Relevant elasticities in Table 4 suggest that increases in agricultural wages prompt a substitution of capital in the form of machinery (with an estimated elasticity of 0.29), fertilizer (0.56), and, to some extent, family labor (0.13) for hired labor. Overall labor intensity declines (with an elasticity of −0.12), as increased supply of family labor is not large enough to compensate for the decline in hired labor (−0.24). By contrast, higher non-farm wages are estimated to have a similarly sized impact on machinery use (elasticity of 0.31) but a negative impact on the labor intensity of agricultural production (−0.09), as a reduction in family labor supply (−0.08) is only partly compensated for by increased use of hired labor (0.35).[10] Increases in the price of bullocks are estimated to result in lower demand for hired labor (elasticity of −0.14, possibly due to reduced demand for drivers that come with bullocks; this is compensated for by higher use of family labor (0.12), fertilizer (0.09), and, to some extent, machinery (0.05). Higher prices of fertilizer are estimated to lead to significant reductions in fertilizer use (−0.39) and increases in total labor (0.07) and machinery (0.05), with no appreciable impact on hired labor use. Higher output prices are estimated to increase demand for all inputs except family labor, due in part to the inelastic supply of the latter. Coefficients for endowments are largely consistent with expectations—that is, higher family labor endowments, as well as female headship, result in increased use of family labor. Although we do not find significant impacts of the head's age, having a more educated head is estimated to be associated with significantly higher levels of machine use.

4. Exploring changes in and determinants of labor market functioning

To test more directly whether labor market imperfections may have been a reason for the observed relationship, we explore efficiency differences between family and hired labor, which may be indicative of

[10] In other words, regressions suggest that a 10.0% increase of the agricultural (non-agricultural) wage rate would result in an increase of rented machine use of 2.0% (2.7%), an increase in the use of fertilizer by 3.4%, and a 3.2% (1.3%) reduction in hired or total labor use, respectively.

supervision constraints. The results imply that such differences seem to have been present in the first but not the second period; therefore, we test whether, in the second period, labor demand and supply were separable not only overall but also for villages with different levels of non-agricultural labor demand. Non-separability is rejected in the aggregate but not for villages with very little non-agricultural labor demand, supporting the importance of non-agricultural labor markets in attenuating the negative farm size–productivity relationship over time.

4.1 Differences in efficiency between own and hired labor

The fact that wage workers require supervision by family members has long been viewed as a root cause of the inverse relationship (Frisvold 1994). We can test this by including the ratio of hired to total labor in a modified labor demand function (Benjamin 1995). To do so, we estimate

$$lnL_{ijt} = \alpha_i + \gamma_1 lnA_{ijt} + \gamma_2 lnw_{ij} + \gamma_3 lnP_{jt} + \gamma_4 H_{ijt} + \gamma_5 D_t + \gamma_6 lnM_{ijt} + \varepsilon_{ijt} \qquad (3)$$

where lnL_{ijt} is the logarithm of labor used in crop production and most other variables are as defined in equation (2); the only addition is that $M_{ijt} = (F_{ijt}+1)/L_{ijt}$, which is the ratio of family (F) to total labor. The sign and significance of γ_6 provide a direct test for differences in efficiency between family and hired labor: a significantly negative (positive) value would imply that family labor is more (less) efficient than hired labor. We again estimate (3) for the early (1982–1999) and late (1999–2008) periods separately. Data for the latter period also allow us to differentiate pre-harvest from harvest labor, and as a robustness we include the pre-harvest family labor ratio and the post-harvest family ratio as explanatory variables. In these regressions, household fixed effects allow us to control for unobserved household-level attributes, including ability. As the family–to–total labor ratio, M_{ijt}, is likely to be affected by endogenous household division and simultaneity bias so that instrumentation is needed (Benjamin 1995). We use demographic variables as instruments and also use the amount of land inherited by a household as an instrument for A_{ijt}.

Table 5 reports the results with and without IVs for the early (columns 1 and 2) and the later (columns 3-6) periods, respectively. In the regressions without IVs, we control for more household demographics, including household size, share of female and male primary adults, and share of elderly females and males. Our interpretation is based on the IV results. The significant negative coefficient (–0.26) on M_{ijt} in the first period suggests that, an additional unit of hired labor added less to output than an additional unit of family labor, plausibly due to supervision constraints. In 1999–2008, by contrast, the coefficient (in column 4) is no longer significant, pointing toward disappearance of the productivity difference between hired and family labor, possibly due to greater use of machinery or capital-intensive methods of production that require less supervision and that allows for easier monitoring. Test statistics in the bottom of Table 5 suggest that, for the early period, IVs fail to pass the over-identification and weak IV tests, reducing our confidence

in the IV results for this period. However, results comfortably pass both tests for the later period where IV results are consistent with those from the fixed effects (FE) regression,[11] suggesting that the contribution to output by own labor and hired labor is not significantly different (column 4). This result does not change even if relevant ratios are entered separately for pre-harvest and harvest seasons (column 6).[12] A plausible explanation for this finding is that, with wage growth, the mix of inputs for agricultural production shifted from labor towards machinery. Although Indian agriculture is a far cry from the highly skill- and data-intensive methods of modern precision farming, use of machinery reduces scope and need for supervision while also requiring hired workers to have higher levels of skill.

4.2 Testing for separability between labor supply and demand decisions

In the presence of labor market frictions, households' labor endowment will affect total labor use, resulting in non-separability between farm households' consumption (labor supply) and production (labor demand) decisions. An alternative to substitution of capital for labor is that farmers who may have been rationed out of non-agricultural labor markets in period one may have been able to overcome barriers to participation in the second period, thus leading to the disappearance of the relationship. Following Benjamin (1992), we can explore this by testing whether, if expected with well-functioning labor markets, the amount of labor used in agricultural production is independent of households' labor endowment. We estimate

$$lnL_{ijt} = \alpha_i + \delta_1 lnA_{ijt} + \delta_2 lnw_{ij} + \delta_3 lnP_{jt} + \delta_4 H_{ijt} + \delta_5 D_t + \delta_6 n_{ijt} + \varepsilon_{ijt} \qquad (4),$$

where lnL_{ijt} is the logarithm of labor used in crop production; A_{ijt}, H_{ijt}, and D_t are as above; w_{ij} is a vector of prices, including farm and non-farm wages, fertilizer price, and price of bullocks; n_{ijt} is a vector of household demographic variables, including household size, share of prime male adult in household, share of prime female adult in household, share of elder male adult in household and share of elder female adult in household; and α_i is household fixed effect. The parameters of main interest are the vector of δ_6, and the absence of labor market imperfection would imply $\delta_6 = 0$ jointly for all demographic variables. Availability of panel data allows us to include household fixed effects. To assess changes in labor market performance over time, we estimate separate regressions for early (1982–1999) and late (1999–2008) periods, distinguishing pre-harvest labor demand from total labor demand.

Results in table 6 suggest that, in 1982–1999, household size had a significant impact on total labor demand. Although few other coefficients are individually significant at conventional levels, the joint test at the bottom of table 6 points toward rejection of the hypothesis of household characteristics, including size,

[11] The fact that the exclusion restriction test failed for the early panel (1982–1999) but not for the late panel (1999–2008) is consistent with the result of the separability test (equation (5)): separability holds for the late panel period but not for the early panel period.

[12] F tests fail to reject the null hypothesis that the ratios of family labor to total labor for pre-harvest and harvest seasons are jointly insignificant for the regressions without IVs (column 5) and with IVs (column 6) at any conventional levels.

female headship, and age composition, not having affected use of labor for crop production in the pre-harvest stage and overall (columns 1 and 2). The relationship disappears for the second period (columns 3 and 4), pointing toward better functioning of agricultural labor markets. To formally check whether labor market imperfections were present in the first but not the second period, we test for coefficients on all demographic variables jointly being equal to zero. Results, reported in the bottom of Table 6, suggest that this null hypothesis can be rejected at 5% in the first but not the second period, supporting the notion that labor market functioning did improve over time.

Estimated coefficients on other variables are mostly consistent with expectations. Negative and significant coefficients on agricultural wage in all equations are consistent with downward sloping demand curves. The same is true for negative and significant coefficients (except in one case) on non-agricultural wages. Where they are significant, coefficients on fertilizer and bullock prices have the expected positive signs. Demand for pre-harvest farm labor increases in cropped area as well and the household head's gender was a key determinant of labor demand in the early but not the later panel period.

Although results from estimating equation (5) allowed us to reject the hypothesis of labor market imperfections in 2008 for the total sample, we noted earlier that the negative relationship between farm size and productivity did not disappear entirely in the second period (see Figure 1 & Table 3). If this is a result of labor market imperfections, we should be able to document these empirically. Variation in non-agricultural labor demand across villages can provide a plausible cause of exogenous variation and to explore whether it may play a role, we split sample villages into three equally sized groups with high, medium, and low levels of nonfarm labor demand and we interacted household characteristics with dummies for each group to test for heterogeneity along this dimension in the 1999–2008 panel.[13]

Results shown in Table 7 are for specifications in which we interacted total household size (columns 1 and 2) and other demographic variables (columns 3 and 4) with indicator variables for high, medium and low levels of village level activity in non-agricultural labor markets. Estimated coefficients provide support for the hypothesis that low levels of non-agricultural labor demand are a key driver of observed labor market imperfections and most plausibly the inverse relationship between farm size and productivity. Although the interactions indicate that household size is insignificant in villages with more active non-agricultural labor demand, the coefficient on household size interacted with low level of non-farm labor market activity is highly significant for pre-harvest and total labor (columns 1 and 2). The result is robust to the addition of more indicators of household composition. In addition to one of these indicators (share of elderly male adults) being significant for the group with low level of nonagricultural labor market activity, results from

[13]The level of nonfarm labor demand is measured by the share of nonfarm work to total work in the village. Such information is not available in the 1982 survey so we only use the later panel for this analysis.

the test for joint significance on all household composition variables for the respective category (see the bottom rows of columns 3 and 4) reject the null hypothesis of coefficients on demographic variables jointly being equal to zero for this group. In contrast, these demographic variables are jointly insignificant for the groups with high and medium levels of nonagricultural labor market activity.

5. Conclusion and policy implications

Although cross-country differences in non-agricultural productivity have long attracted interest in the literature (Hsieh and Klenow 2009), differences of equal magnitude exist in agricultural production (Adamopoulos and Restuccia 2014a). In fact, a number of studies have started to explore such differences, pointing toward imperfections in factor markets, especially those for land (Restuccia and Santaeulalia-Llopis 2015) and labor, in addition to policies (Adamopoulos and Restuccia 2014b) as key factors.

We complement this literature by using longitudinal data from India spanning the 1982–2008 period to examine how the inverse relationship between farm size and land productivity evolved over time. The main finding that the relationship weakened significantly but did not disappear even in the last period (2008) can be attributed to two factors. Expanded technological options and real wage growth imply that capital was increasingly substituted for hired labor via mechanization of the production process, reducing the need for family labor-based supervision. However, improved functioning of rural labor markets via greater non-agricultural demand seems to have helped establish the preconditions for separability in the second period, possibly by allowing farmers who had earlier been rationed out to participate in such markets.

Our main finding is that, over the 25 years considered, the inverse farm size–productivity relationship in India weakened in response to better labor market functioning. This not only provides an economically meaningful explanation of an empirically robust phenomenon, but also has implications for policies recommended to address this issue. Where it could be implemented, redistributive land reform has helped increase small farmers' productivity and investment.[14] But if market imperfections, the nature and severity of which are likely to change over time, have a significant impact, a response that carefully considers historical context and weighs advantages and disadvantages of different options will be needed.[15]

Although better labor market functioning can explain the gradual disappearance of the inverse relationship overall, labor productivity still shows considerable divergence across producers. Exploring the extent to

[14] In India, land reform had positive impacts (Besley and Burgess 2000), though it often did not target the poorest (Besley *et al.* 2016) where it could be implemented (Deininger, Jin and Nagarajan 2009). However, it triggered tenant evictions (Appu 1997), which is consistent with implementation challenges in other contexts (Benjamin *et al.* 2012). In India, with implementation having come to a virtual standstill, mechanisms used (such as prohibition of rentals) remain in place and undermine investment incentives (Deininger, Jin and Yadav *et al.* 2013) and land market functioning (Deininger, Jin and Nagarajan 2008).

[15] See Keswell and Carter (2014) for a recent study suggesting that, in the context of South Africa, benefits from such reform can still outweigh the considerable costs incurred by such a policy.

which such differences may be attributed to restrictions on the operation of land lease markets that were in place throughout the period could provide evidence to inform ongoing initiatives aiming to make it easier for states to liberalize such markets as a way to improve India's agricultural productivity.[16]

[16] Efforts in this direction have recently led to circulation of a model law to allow land leasing that can be adopted by states (Niti Aayog 2016).

Table 1: Household characteristics, 1982, 1999, and 2008

	1982	1999	2008
Household characteristic			
Head's age (years)	50.91	50.07	51.64
Head's education (years)	2.20	4.29	4.97
Household size	7.35	6.53	5.68
# members >14 years old	4.65	4.38	4.01
Share of prime males	48.59	46.82	44.08
Share of prime female	43.37	43.10	42.76
Share of elderly males	4.71	5.33	6.69
Share of elderly females	3.33	4.75	6.48
Income and endowments			
Area owned (acres)	8.64	5.60	4.46
Area cultivated (acres)	9.01	7.18	6.83
Owns machinery	0.07	0.20	0.32
… if yes, value ('000 Rs./acre)	6.44	55.40	64.26
Total income ('000 Rs)	30.71	65.65	79.96
Share from agriculture (%)	84.48	64.64	61.26
Labor use			
Using any labor for agric. production (%)	98.70	97.17	97.16
Using family labor for agric. production (%)	97.58	97.01	95.37
Using hired labor for agric. production (%)	70.01	72.13	77.86
Worked as agric. labor (%)	25.06	20.94	20.29
Worked as nonagricultural labor (%)	5.76	13.70	19.86
Engaged in nonagricultural self-employment (%)	6.41	7.86	17.55
Cultivation details			
Agricultural labor (days/acre)	80.81	52.58	51.39
Own labor (days/acre)	42.67	44.04	37.85
Male labor		28.25	26.49
Female labor		11.54	10.80
Hired labor (days/acre)	38.56	8.54	13.55
Male labor		5.33	8.72
Female labor		2.82	4.60
Rents or owns machinery	50.71	62.04	74.11
… if yes, value (Rs./acre)	115.01	509.20	989.38
Share of who neither own nor hire tractors	0.49	0.38	0.26
Village level variables			
Agricultural wage (Rs/day)	15.00	45.35	47.59
Nonagricultural wage (Rs/day)	16.01	57.48	67.74
Number of observations	3,746	4,518	4,924

Source: Own computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Table 2: Key characteristics by farm size group, 1982, 1999, and 2008

	1982			1999			2008		
Variable	**Small**	**Med.**	**Large**	**Small**	**Med.**	**Large**	**Small**	**Med.**	**Large**
Household characteristics									
Head's age (years)	48.64	49.91	52.75	47.77	49.78	54.08	49.75	52.19	54.60
Head's education (years)	2.03	2.18	2.28	3.98	4.25	4.84	4.76	4.75	5.86
Household size	5.90	6.80	8.45	5.89	6.25	7.98	5.26	5.65	6.62
# members >14 years old	3.80	4.33	5.29	3.81	4.22	5.50	3.59	4.05	4.85
Share of prime males	48.07	48.83	48.59	47.37	46.86	46.16	43.92	44.41	43.80
Share of prime female	44.51	43.82	42.69	43.55	43.18	42.53	43.89	42.43	41.50
Share of elderly male	4.10	4.46	5.08	4.56	5.42	6.06	6.07	6.83	7.43
Share of elderly females	3.32	2.89	3.65	4.52	4.54	5.25	6.12	6.33	7.28
Income and endowments									
Area owned (acres)	1.07	3.73	16.13	1.21	4.04	14.80	1.08	3.71	13.09
Area cultivated (acres)	1.75	4.54	15.98	2.25	5.70	17.11	2.09	5.88	18.72
Owns machinery	0.01	0.03	0.13	0.14	0.17	0.33	0.18	0.38	0.52
…if yes, value ('000 Rs./acre)	1.34	1.53	7.62	8.50	31.37	106.35	22.944	43.88	122.87
Total income ('000 Rs)	14.50	21.85	45.30	38.80	55.39	123.11	46.539	74.49	161.55
Share from agriculture	50.67	71.87	94.37	40.34	63.55	77.23	34.48	58.01	80.52
Labor use									
Using any labor for agric. production (%)	96.42	99.13	99.32	92.87	99.88	99.55	93.46	99.89	99.80
Using family labor for ag. (%)	96.11	97.86	97.97	92.76	99.71	99.36	92.40	98.29	96.11
Using hired labor for ag. (%)	52.96	67.30	79.72	58.36	78.18	83.92	65.88	84.16	91.30
Worked as agric. labor (%)	47.66	29.77	11.22	34.28	17.28	6.09	29.50	17.44	6.14
Worked as nonagric. labor (%)	7.17	6.65	4.39	20.55	11.98	5.81	29.21	15.20	8.90
Worked as nonagric. self-employment (%)	7.63	5.70	6.49	9.13	6.52	7.99	18.33	17.71	15.56
Cultivation details									
Agric.labor (days/acre)	158.98	79.72	48.80	91.55	37.75	19.86	82.50	36.56	17.71
Own labor (days/acre)	107.37	39.68	17.96	79.66	30.37	14.31	63.98	24.91	10.47
Male labor				51.41	19.24	9.12	44.64	17.39	7.69
Female labor				20.98	8.05	3.47	18.30	7.21	2.70
Hired labor (days/acre)	52.11	40.46	31.22	11.89	7.38	5.55	18.52	11.65	7.24
Male labor				7.64	4.32	3.59	11.83	7.55	4.73
Female labor				3.72	2.69	1.71	6.45	3.86	2.34
Rents or owns machinery	44.08	49.64	54.50	55.30	66.78	65.03	69.63	76.30	79.43
… if yes, value (Rs./acre)	192.46	109.04	92.47	644.70	485.76	369.08	1413.15	795.49	555.44
Village level variables									
Agricultural wage (Rs/day)	15.66	14.51	15.14	45.19	45.22	45.81	46.83	47.93	48.57
Nonagric. wage (Rs/day)	15.73	15.25	16.78	58.16	56.85	57.42	66.39	67.81	70.50
Number of observations	825	1,363	1,558	1,698	1,719	1,101	2,078	1,869	977

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Note: Large, medium, and small farms are defined as having more than 5, between 2.5 and 5, and less than 2.5 acres of land, respectively.

Table 3: Estimates of the farm size-productivity relationship, 1982/99 and 1999/2008

	1982-99		1999-2008	
	FE	**FE-IV[a]**	**FE**	**FE-IV[a]**
	(1)	**(2)**	**(3)**	**(4)**
Log of land	0.781***	0.734***	0.816***	0.946***
	(15.329)	(9.380)	(21.452)	(21.538)
Share of irrigated land	0.248***	0.254***	0.191***	0.166**
	(4.093)	(4.192)	(2.912)	(2.575)
Share of leased land	0.103	0.121	0.016	-0.016
	(1.217)	(1.387)	(0.240)	(-0.246)
Log of assets	0.048***	0.052***	0.022***	0.013
	(3.882)	(4.045)	(2.687)	(1.496)
Head is female	-0.027	-0.025	-0.036	-0.042
	(-0.398)	(-0.359)	(-0.728)	(-0.850)
Head's age	-0.197	-0.399	-0.346	0.177
	(-0.202)	(-0.383)	(-0.298)	(0.155)
Head's age squared	0.022	0.051	0.061	-0.014
	(0.169)	(0.367)	(0.408)	(-0.093)
Head's education	-0.077***	-0.074***	0.027	0.029
	(-2.642)	(-2.599)	(1.373)	(1.464)
No. of males 14-64 years old	0.004	0.014	0.063**	0.039
	(0.109)	(0.348)	(2.135)	(1.374)
No. of females 14-64 years old	0.017	0.025	0.033	0.016
	(0.436)	(0.701)	(1.364)	(0.636)
No. of males > 64 years old	0.117	0.115	0.073	0.020
	(0.356)	(0.341)	(0.477)	(0.118)
No. of females > 64 years old	0.496**	0.520**	0.136	0.106
	(2.017)	(2.097)	(0.766)	(0.599)
R^2	0.546	0.545	0.580	0.568
No. observation	7,146	7,146	6,389	6,389
No. farmers	4,145	3,001	3,805	2,584
Hansen J statistic (overid test)		8.849		9.783
Chi-sq P-value		0.1822		0.2806
Anderson-Rubin Wald test		21.68		25.69
Stock-Wright LM S statistic		82.57		93.87

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Note: The dependent variable is the output value per acre in logs. All specifications are for the first difference. Standard errors adjusted for clustering effect at the village level, and asymptotic t-ratios are given in parentheses (*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$). [a] As discussed in the text, we instrument cultivated area using as instruments inherited land and assets, the number of male and female claimants, and the variance of their schooling, as well as the number of co-resident brothers of the head. FE=fixed effects; IV=instrumental variable.

Table 4: Estimates of factor demand equations

	Hired labor (1)	Family labor (2)	Total labor (3)	Machine use (4)	Fertilizer (5)
Output price	0.067***	-0.033***	0.056***	0.102***	0.127***
	(6.474)	(-5.439)	(15.056)	(12.648)	(11.507)
Factor prices					
Agricultural wage	-0.244***	0.133***	-0.116***	0.293***	0.564***
	(-4.111)	(3.831)	(-4.032)	(6.307)	(8.911)
Nonagricultural wage	0.348***	-0.077**	-0.089***	0.310***	0.104
	(5.792)	(-2.203)	(-3.069)	(6.595)	(1.627)
Price of bullocks	-0.142***	0.122***	0.035**	0.050*	0.086**
	(-4.192)	(6.185)	(2.160)	(1.902)	(2.385)
Price of fertilizer	-0.004	0.092***	0.070***	0.048**	-0.388***
	(-0.145)	(5.314)	(4.893)	(2.075)	(-12.299)
Endowments					
Log of land	0.672***	0.321***	0.413***	0.201***	0.605***
	(20.868)	(17.045)	(26.458)	(7.995)	(17.626)
# members14-64 years old	-0.479***	0.145***	0.027	-0.024	0.105**
	(-10.588)	(5.502)	(1.232)	(-0.665)	(2.178)
Head is female	-0.163	0.188***	0.104*	-0.216**	0.001
	(-1.470)	(2.895)	(1.932)	(-2.479)	(0.007)
Head's age	1.156	0.582	0.701	-0.924	0.023
	(0.744)	(0.642)	(0.932)	(-0.761)	(0.014)
Head's age squared	-0.153	-0.050	-0.074	0.144	-0.016
	(-0.756)	(-0.423)	(-0.752)	(0.911)	(-0.074)
Head's education	0.013	-0.005	-0.011	0.101***	-0.037
	(0.414)	(-0.262)	(-0.701)	(4.034)	(-1.079)
Constant	0.791	1.396	3.018**	0.212	2.791
	(0.266)	(0.806)	(2.100)	(0.091)	(0.883)
No. of observations	10,204	10,204	10,204	10,204	10,204
R-squared	0.143	0.115	0.384	0.247	0.189
Number of farmers	4,254	4,254	4,254	4,254	4,254

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Note: Dependent variable is the log of labor quantity (days) or monetary value of machinery hire and fertilizer. All regressions include time dummies and farmer fixed effects. Standard error is adjusted for clustering at the village level throughout, and asymptotic t-ratios are given in parentheses (*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$).

Table 5: Test for efficiency differences between own labor and hired labor

	1982-99		1999-08			
	FE	FE (IV)[a]	FE	FE (IV)[a]	FE	FE (IV)[a]
	(1)	(2)	(3)	(4)	(5)	(6)
Log of land	0.348***	0.420***	0.364***	0.350***	0.341***	0.357***
	(16.181)	(14.645)	(16.681)	(16.084)	(17.047)	(11.480)
Agriculture wage	-0.475***	-0.260**	-0.171***	-0.211***	-0.195***	-0.197***
	(-11.569)	(-2.543)	(-2.941)	(-3.708)	(-3.665)	(-3.223)
Non agriculture wage	-0.057	-0.205***	-0.092*	-0.191***	-0.195***	-0.155*
	(-1.480)	(-4.572)	(-1.839)	(-4.050)	(-4.241)	(-1.819)
Price of fertilizer	-0.004	-0.019	0.151***	0.152***	0.149***	0.149***
	(-0.190)	(-0.755)	(7.295)	(7.929)	(7.898)	(7.467)
Price of bullocks	0.200***	0.092***	0.006	-0.021	-0.036	-0.050
	(10.378)	(3.109)	(0.142)	(-0.511)	(-0.994)	(-0.946)
Head's age	0.936	1.005	-2.094	-2.416**	-2.404**	-2.684**
	(1.030)	(1.026)	(-1.635)	(-2.204)	(-2.054)	(-2.144)
Head's age squared	-0.109	-0.126	0.274	0.316**	0.314**	0.350**
	(-0.903)	(-0.984)	(1.629)	(2.201)	(2.044)	(2.135)
Head's education	-0.023	-0.045	0.001	-0.007	-0.011	-0.009
	(-1.035)	(-1.493)	(0.066)	(-0.380)	(-0.549)	(-0.471)
Head is female	0.141*	0.253**	0.095	0.067	0.085	0.119
	(1.757)	(2.494)	(1.262)	(0.975)	(1.242)	(1.233)
Log of household size	0.149***		0.031		0.015	
	(4.114)		(0.753)		(0.384)	
Share males 14-64 years old	-0.033		-0.090		-0.106	
	(-0.359)		(-0.865)		(-1.115)	
Share females 14-64 years old	0.108		0.035		0.034	
	(0.934)		(0.284)		(0.299)	
Share males >64 years old	0.024		0.018		0.003	
	(0.114)		(0.088)		(0.013)	
Share females >64 years old	0.242		-0.202		-0.232	
	(1.236)		(-1.092)		(-1.370)	
Family to total labor ratio	-0.264***	0.283	0.019	0.118		
	(-15.180)	(1.308)	(0.662)	(0.655)		
Family to total labor ratio (pre-harvest)					-0.045	-0.314
					(-1.273)	(-0.561)
Family to total labor ratio (harvest season)					0.006	0.359
					(0.178)	(0.625)
Constant	3.133*	3.795*	8.115***	10.038***	9.992***	10.463***
	(1.827)	(1.884)	(3.326)	(4.773)	(4.483)	(4.477)
Observations	8,314	8,314	7,400	7,400	7,400	7,400
R-squared	0.341	0.227	0.236	0.255	0.259	0.229
Number of farmers	4,925	4,925	4,439	4,439	4,439	4,439
Hansen J stat. (overid test)		11.603		4.22		3.396
Chi-sq P-val		0.0206		0.3771		0.2725
Kleibergen-Paap rk LM stat.		28.075		35.3		0.3345

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Notes: Dependent variable is the log of total labor days used for agricultural production. All specifications include farmer fixed effects and year fixed effects. Absolute value of asymptotic t-ratios are in parentheses. FE = fixed effects; IV = instrumental variable.a Instruments include household size, share of prime male workforce, share of prime female workforce, share of elderly male workforce, and share of elderly female workforce. All prices are in logarithms.

Table 6: Test for labor market efficiency for preharvest and total labor

	1982-99		1999-08	
	Pre-harvest	**Total**	**Pre-harvest**	**Total**
	(1)	**(2)**	**(3)**	**(4)**
Log of land	0.377***	0.381***	0.365***	0.344***
	(17.062)	(17.182)	(16.823)	(17.335)
Agricultural wage	-0.588***	-0.138***	-0.173***	-0.200***
	(-14.100)	(-3.285)	(-2.980)	(-3.766)
Non-agricultural wage	-0.062	-0.199***	-0.091*	-0.195***
	(-1.566)	(-5.010)	(-1.815)	(-4.261)
Price of fertilizer	-0.012	-0.009	0.151***	0.150***
	(-0.574)	(-0.422)	(7.316)	(7.951)
Price of bullocks	0.177***	0.119***	0.008	-0.031
	(8.916)	(5.985)	(0.192)	(-0.861)
Log of household size (α_1)	0.108***	0.098***	0.029	0.010
	(2.893)	(2.611)	(0.693)	(0.260)
Share males 14-64 years old (α_2)	-0.091	-0.161*	-0.094	-0.113
	(-0.962)	(-1.696)	(-0.903)	(-1.196)
Share females 14-64 years old (α_3)	0.047	-0.007	0.035	0.036
	(0.396)	(-0.061)	(0.280)	(0.317)
Share males >64 years old (α_4)	0.021	0.019	0.019	0.006
	(0.099)	(0.089)	(0.091)	(0.030)
Share females >64 years old (α_5)	0.084	-0.086	-0.198	-0.223
	(0.413)	(-0.425)	(-1.071)	(-1.321)
Head's age	0.886	1.205	-2.079	-2.364**
	(0.943)	(1.278)	(-1.624)	(-2.021)
Head's age squared	-0.107	-0.148	0.273	0.310**
	(-0.854)	(-1.182)	(1.620)	(2.012)
Head's education	-0.043*	-0.022	0.002	-0.010
	(-1.898)	(-0.967)	(0.081)	(-0.521)
Head is female	0.196**	0.176**	0.093	0.080
	(2.363)	(2.114)	(1.241)	(1.166)
Constant	3.940**	2.662	8.089***	9.928***
	(2.224)	(1.497)	(3.316)	(4.455)
No. of observations	8,314	8,314	7,400	7,400
R-squared	0.291	0.375	0.236	0.257
Number of farmers	4,925	4,925	4,439	4,439
Joint F-test (α_1+ α_2+ α_3+ α_4+ α_5=0)	2.899	3.437	0.819	0.942
P-value	0.0129	0.00423	0.536	0.452

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Note: Dependent variable is the log of preharvest and total labor days used for agricultural production. Absolute values of asymptotic t-ratios are in parentheses. All specifications include farmer fixed effect and year fixed effect. Standard errors adjusted for clustering effect at the village level. All prices are in logarithms.

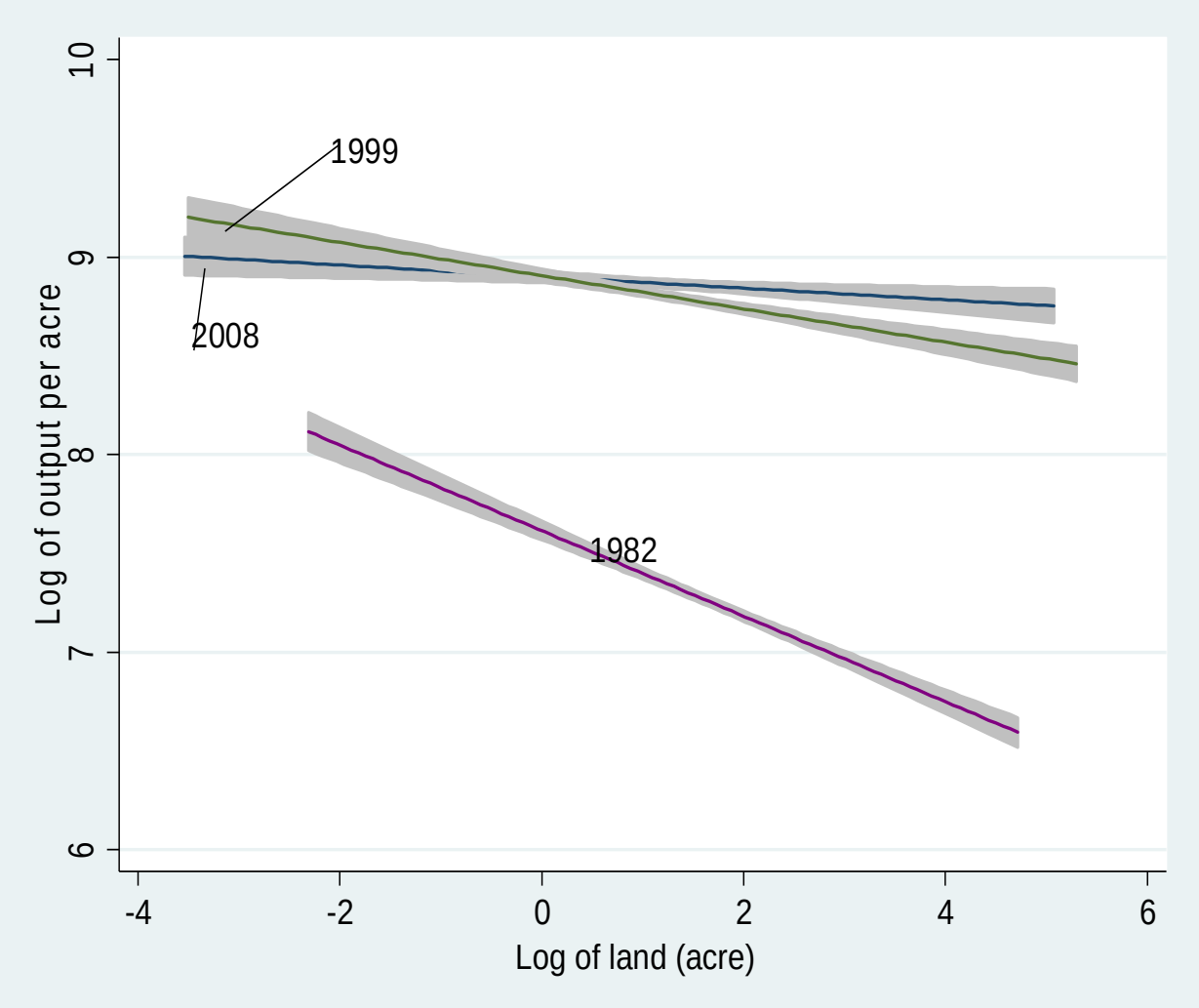
Table 7: Exploring heterogeneity in labor markets

	Pre-harvest labor (1)	Total labor (2)	Pre-harvest labor (3)	Total labor (4)
Log of land	0.359***	0.338***	0.360***	0.340***
	(16.998)	(17.329)	(16.863)	(17.246)
Log of agricultural wage	-0.228***	-0.238***	-0.232***	-0.242***
	(-4.005)	(-4.514)	(-4.060)	(-4.582)
Log of nonagricultural wage	-0.088*	-0.197***	-0.082*	-0.192***
	(-1.786)	(-4.348)	(-1.666)	(-4.230)
Price of fertilizer	0.155***	0.152***	0.153***	0.150***
	(7.656)	(8.138)	(7.573)	(8.056)
Village price of hiring bullock	-0.083**	-0.085**	-0.086**	-0.085**
	(-2.124)	(-2.333)	(-2.186)	(-2.336)
Log of household size * Low (α_1)	0.135***	0.120**	0.143**	0.121**
	(2.653)	(2.545)	(2.518)	(2.291)
Log of household size * Medium (β_1)	0.065	0.035	0.037	0.009
	(1.369)	(0.809)	(0.671)	(0.166)
Log of household size * High (γ_1)	-0.039	-0.044	-0.053	-0.076
	(-0.735)	(-0.916)	(-0.883)	(-1.360)
Share of prime male adult * Low (α_2)			-0.071	-0.138
			(-0.458)	(-0.963)
Share of prime male adult * Medium (β_2)			-0.136	-0.126
			(-0.843)	(-0.842)
Share of prime male adult * High (γ_2)			0.004	-0.040
			(0.028)	(-0.283)
Share of prime female adult * Low (α_3)			-0.033	0.067
			(-0.166)	(0.367)
Share of prime female adult * Medium (β_3)			-0.023	-0.036
			(-0.123)	(-0.208)
Share of prime female adult * High (γ_3)			0.087	0.028
			(0.465)	(0.162)
Share of elderly male adult * Low (α_4)			0.632**	0.516*
			(2.074)	(1.834)
Sh of elderly male adult * Medium (β_4)			-0.134	-0.029
			(-0.471)	(-0.110)
Sh of elderly male adult * High (γ_4)			-0.213	-0.322
			(-0.738)	(-1.211)
Sh of elderly female adult * Low (α_5)			-0.118	-0.080
			(-0.362)	(-0.266)
Sh of elderly female adult * Medium (β_5)			-0.295	-0.358
			(-1.095)	(-1.438)
Sh of elderly female adult * High(γ_5)			0.033	-0.030
			(0.116)	(-0.115)
No. of observations	7,400	7,400	7,400	7,400
R-squared	0.217	0.251	0.218	0.251
Joint F-test for Low level			2.493	2.443
P-value1			0.0292	0.0322
Joint F-test for Medium level			0.814	0.640
P-value2			0.540	0.669
Joint F-test for high level			0.394	0.676
P-value3			0.853	0.641

Source: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Note: Constant, time dummy, gender, education, and age of head and its square are included but not reported. Low, medium, and high refer to terciles of village-level demand for nonagricultural labor. t-statistics are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure 1: The changes in the farm size-productivity relationship 1982, 1999, and 2008



Source: Based on results from locally weighted polynomial regressions of total money yield against farm size for 1982, 1999, and 2008.
Note: Confidence intervals are at 95 percent level.

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