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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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Contents

Abstract	V
Acknowledgments	vi
1. Introduction	1
2. Motivation and Relevance	3
3. Changes in the Farm Size–Productivity Relationship and Factor Demand Over Time	8
4. Exploring Changes in and Determinants of Labor Market Functioning	12
5. Conclusion and Policy Implications	18
References	19

Tables

2.1 Household characteristics, 1982, 1999, and 2008	6
2.2 Key characteristics by farm size group, 1982, 1999, and 2008	7
3.1. Estimates of the farm size-productivity relationship, 1982-1999 and 1999-2008	10
3.2 Estimates of factor demand equations	11
4.1 Test for efficiency differences between own labor and hired labor	13
4.2 Test for labor market efficiency for preharvest and total labor	15
4.3 Exploring heterogeneity in labor markets	16
Figures	
3.1 The changes in the farm size–productivity relationship, 1982, 1999, and 2008	8

ABSTRACT

The existence of an inverse relationship between farm size and output per unit of land is well documented. However, little research focuses on if and how an inverse relationship between farm size and productivity changes when labor market performance improves. To fill this gap, we use a large national farm panel from India covering a quarter-century (1982, 1999, and 2008) to show that the inverse relationship weakened significantly over time, despite an increase in the dispersion of farm sizes. A key reason was the substitution of capital for labor in response to nonagricultural labor demand. In addition, family labor was more efficient than hired labor in the 1982–1999 period, but not during the 1999–2008 period. 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.

Keywords: inverse farm size-productivity relationship; labor market imperfections; farm size in 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—including, but 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, it is not clear which markets are affected, how amenable they are to change, and what this implies for policy. In particular, although the inverse relationship has been used to recommend redistributive land reform as a way to improve efficiency and equity (Lipton 2009), political resistance has often made such reforms difficult to implement. Recognition that rising nonagricultural 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 reconsider such policy advice. 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 relationships evolves in response to changes in technology and market functioning.1

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). The paper uses panel regressions for the early (1982–1999) and late (1999–2008) periods—in both cases, instrumenting size of cultivated area in order to explore the changes in the inverse relationship over time and to test different explanations of this relationship in terms of 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 across states; this finding provides a source of identification not available to cross-sections.

The results from nonparametric and parametric regressions suggest that the inverse farm size—productivity relationship attenuated significantly over the 1982–2008 period, consistent with the findings in Vietnam from 1992-2008 (Liu, Violette, and Barrett 2016). To understand mechanisms underlying such changes, we note that higher wage levels in the 2000s were likely to have translated into 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 using the same approach suggest 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 the labor demand function to become insignificant—and this is indeed observed in the data. Alternatively if, with improved functioning of agricultural and

¹ Liu, Violette and Barrett (2016) is probably the only study that used panel data to explore the evolution of the relationship between farm size and productivity over a long period.

nonagricultural 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 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, although separability is not rejected overall in the second period, it is rejected in villages with inactive nonagricultural labor markets, possibly explaining the persistence of the inverse relationship between farm size and productivity, though less pronounced, in the second period. The finding of better functioning of labor markets provides a coherent explanation for the observed phenomenon that the negative farm size—productivity relationship weakened significantly over time in India; it suggests that efforts to improve functioning of rural factor markets are a more appropriate response to this phenomenon than redistributive land reforms, which have often been difficult to implement and, as a result, created further impediments to the functioning of rural factor markets.

The paper is structured as follows: Section 2 discusses the literature, data, and descriptive statistics, indicating the far-reaching changes experienced by India during the period considered. Section 3 presents results on the inverse farm size–productivity relationship and the factor demand equations. Section 4 focuses on explanations in terms of labor market imperfections by testing the relevance of family labor and separability of labor supply and demand. Section 5 concludes with implications for policy and future research.

2. MOTIVATION AND RELEVANCE

This section draws on conceptual and empirical literature to highlight the strong evidence in support of a negative farm size—productivity relationship and discusses potential channels underpinning this phenomenon. If an inverse relationship between productivity and farm size can be attributed to labor market imperfections, whereas capital market imperfections have the opposite effect, one would expect the relationship to weaken as agricultural production becomes more capital intensive and as nonagricultural labor demand increases. Longitudinal data spanning a long enough period should uncover such patterns.

The Farm Size-Productivity Relationship

A negative relation between farm size and output per unit of land in unmechanized agriculture has been confirmed many times in the literature (Eastwood, Lipton, and Newell 2010). 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 is 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 bias in farmers' estimates of area cultivated 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). 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 Yúnez-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—for example, 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 by family labor (Frisvold 1994). This dispersion generates a negative relationship between productivity and size as the ratio of hired to family labor increases with farm size.²

Small farmers' advantages in labor supervision, knowledge, and organizational advantages can be offset by their difficulty in accessing capital and insurance. This difficulty arises from the high transaction cost of providing formal credit in rural markets, possibly exacerbated by the difficulty of small farms to provide collateral. Frictions in labor market participation and land markets—due, for example, to transaction costs—could motivate small farmers who are unable to rent to rationally apply family labor to cultivate their fixed land endowment more intensively than with perfect markets. 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) plus the advantages in getting access to

² Owner-operators' advantage could be further enhanced by their knowledge of local soil and climatic conditions, often accumulated over generations (Rosenzweig and Wolpin 1985).

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). Thus, with few exceptions,³ agricultural production in practice relies on owner-operated firms (Allen and Lueck 1998; Deininger and Feder 2001).⁴

Although supervision constraints can account for the inverse relationship if hired labor is used, off-farm labor market rationing that constrains opportunities for smallholder 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 the negative relationship between farm size and productivity. These imperfections, taken as exogenously given in cross-sectional studies, will vary not only with time but also with how family labor is valued.⁵

Although labor market imperfections will strengthen the negative relationship, capital market imperfections are likely to go the opposite way. Lumpiness of certain technologies (such as bullock and tractors) that cannot be easily overcome without functioning rental markets may attenuate small farmers' advantage. Innovations in crop breeding, tillage, and information technology (Deininger, Nizalov, and Singh 2013) tend to reduce supervision requirements.6 Although technology can improve smallholders' competitive position at the marketing stage (Aker and Fafchamps 2014), use of sophisticated techniques at the production stage will require capital, and rising wages may allow substitution of capital for labor (Yamauchi 2016). 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).

As a consequence, magnitude and possibly even direction of the relation between farm size and productivity will vary in response to exogenous factors, including nonfarm labor demand, availability of technology, and the functioning of capital markets. These factors, as well as policy, will affect factor prices, the optimum ratio of capital to labor used in production, supervision needs, and the extent of market imperfections that traditionally gave 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.

Data and Descriptive Statistics

This study's data came 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.⁷ Each round's

³ 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.

⁴ As of the end of 2009, only seven publicly listed farming companies existed worldwide—three in South America and four in Ukraine and Russia (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).

⁵ 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).

⁶ 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 2012).

⁷ The sample included three strata: (1) districts that were part of the Intensive Agricultural District Program, an extension

household schedule provides data on demographics, assets, income, consumption, and economic activities at the household level, as well as detailed information on labor and nonlabor inputs used in agriculture and outputs from agricultural production. In case of splits, all of a household's successors in the same village were interviewed, and 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 percent in the 1980s, liberalization in the early 1990s resulted in much stronger economic growth, largely driven by the nonfarm 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 percent, reaching a point at which the nonfarm sector employs close to 40 percent of the country's rural workforce (Himanshu et al. 2013). Although land market imperfections clearly affect nonagricultural productivity (Duranton et al. 2015b), functioning of credit markets improved markedly since the 1990s (Duranton et al. 2015a). From 2006 to 2008, 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 2.1 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 (nonagricultural) 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 10-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 percent to 74 percent. The data also point toward a significant increase in off-farm labor opportunities, with the share of farmers who worked in nonfarm wage or self-employment increasing from low levels (6 percent each, respectively) to 20 and 18 percent, respectively, and the share of income from agriculture decreasing from 84 percent in 1982 to 61 percent in 2008.

Table 2.2 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 percent in 1982, 77 percent in 1999, and 81 percent 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 percent in 1982 to 34 percent in 2008 for small and 72 percent to 58 percent for medium farms. The decline was relatively bigger for the small and medium farms but changed only a little (from 94 percent to 81 percent) for the large ones.

Possibly in response to wage rises, 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 percent, respectively. The share of farms owning machinery, however, increased drastically over the survey periods, with rates of increase more pronounced for larger farm sizes. Although few small and medium farms (1 percent and 3 percent, respectively) owned machinery in 1982,

and input provision program placed in areas thought to have high potential for crop productivity growth; (2) districts covered by the Intensive Agricultural Area Programme; and (3) all other districts.

⁸ The 1999 round covered a total of 7,474 households, including all original households and their split-offs. The 2008 round (for the 2007–2008 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.

18 percent and 38 percent, respectively, did so in 2008; machinery ownership for large farms increased from 13 percent in 1982 to 33 percent in 1999 and then to 52 percent in 2008. A similar increase is observed in the rate of renting machinery, though the rate of increase is more similar across farm size groups.

Table 2.1 Household characteristics, 1982, 1999, and 2008

Household characteristic	1982	1999	2008
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 females	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 nonagriculture labor (%)	5.76	13.70	19.86
Engaged in nonagriculture. 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
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: Authors' computation from 1982, 1999, and 2008 ARIS-REDS surveys.

Table 2.2 Key characteristics by farm size group, 1982, 1999, and 2008

		1982			1999			2008	
Variable	Small	Medium	Large	Small	Medium	Large	Small	Medium	Large
Household characteristi	cs								
Head's age (years) Head's education	48.64	49.91	52.75	47.77	49.78	54.08	49.75	52.19	54.60
(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 females	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 ag	50.67	71.87	94.37	40.34	63.55	77.23	34.48	58.01	80.52
Labor use									
Using any labor for ag									
production (%)	96.42	99.13	99.32	92.87	99.88	99.55	93.46	99.89	99.80
Using family labor for ag	00.44	07.00	07.07	00.76	00.74	00.00	00.40	00.00	00.44
production (%) Using hired labor for ag	96.11	97.86	97.97	92.76	99.71	99.36	92.40	98.29	96.11
production (%)	52.96	67.30	79.72	58.36	78.18	83.92	65.88	84.16	91.30
Worked as ag labor (%)	47.66	29.77	11.22	34.28	17.28	6.09	29.50	17.44	6.14
Worked as nonag labor		20		01.20	17.20	0.00	20.00		0.11
(%)	7.17	6.65	4.39	20.55	11.98	5.81	29.21	15.20	8.90
Engaged in nonag self-									
employment (%)	7.63	5.70	6.49	9.13	6.52	7.99	18.33	17.71	15.56
Cultivation details									
Ag 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	44.08	49.64	54.50	55.30	66.78	65.03	69.63	76.30	79.43
machinery if yes, value	44.00	+3.U 4	54.50	55.50	00.70	05.05	09.00	10.30	13.40
(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
Nonag 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 6.0, between 2.0 and 6.0, and fewer than 2.0 acres of land, respectively.

3. CHANGES IN THE FARM SIZE-PRODUCTIVITY RELATIONSHIP AND FACTOR DEMAND OVER TIME

Nonparametric and parametric regressions suggest that the inverse relationship between farm size and productivity were attenuated over time to the point at which it had almost disappeared at the end of the 1992–2008 period. Possible mechanisms for this were explored. 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.

Evidence on Evolution of the Relationship

To illustrate the issue at hand, Figure 3.1 plots results from locally weighted polynomial regressions of total monetary yield against farm size for 1982, 1999, and 2008. The 95 percent confidence intervals are also displayed. Between 1982 and 1999, 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, but the relationship flattened out, largely due to differential performance by farms smaller than and larger than 1 hectare with productivity declining (though insignificant at 5 percent) for the former but increasing (and significant at 5 percent) for the latter.

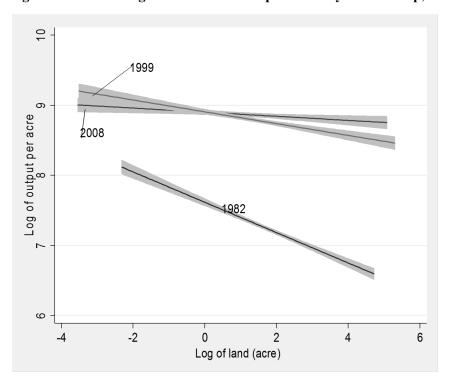


Figure 3.1. The changes in the farm size–productivity relationship, 1982, 1999, and 2008

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.

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

$$\ln y_{ijt} = \alpha_i + \alpha_2 \ln A_{ijt} + \alpha_3 H_{ijt} + \alpha_4 D_t + \varepsilon_{ijt}, \tag{1}$$

where y_{ijt} is value of crop output, 9 A_{ijt} is total cropped area including land quality indicators, H_{ijt} is a vector of household characteristics, D_t is a year dummy, α_i is household fixed effects, and ε_{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, as a result of unobserved production shocks that affect area cultivated, there is a possibility that $E(\ln A_{ijt}\varepsilon_{ijt}) \neq 0$, so that OLS would 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. For 1999–2008, lagged household characteristics that were not available in 1982 are also included as instruments. This strategy relies on the assumption of production shocks being independently distributed over time (Foster and Rosenzweig 2002). Overidentification 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 3.1 reports results of estimating equation 1 for the 1982–1999 (columns 1 and 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, ¹⁰ 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, though 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 later.

⁹ 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.

¹⁰ The Hansen J overidentification 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).

Table 3.1. Estimates of the farm size-productivity relationship, 1982–1999 and 1999–2008

	1982–1999		199	9–2008
	FE	FE-IV ^a	FE	FE-IV ^a
Variable	(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 16-64 years old	0.004	0.014	0.063**	0.039
•	(0.109)	(0.348)	(2.135)	(1.374)
No. of females 16-64 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. of observations	7,146	7,146	6,389	6,389
No. of 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 LMS 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 error is 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). *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.

Determinants of Factor Demand

To explore whether rising labor and other input costs may have affected 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

$$\ln X_{ijt} = \alpha_i + \beta_1 \ln A_{ijt} + \beta_2 \ln w_{it} + \beta_3 \ln P_{it} + \beta_4 H_{ijt} + \beta_5 D_t + \varepsilon_{ijt}, \tag{2}$$

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 nonagricultural 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 3.2 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 nonfarm 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). 11 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.

Table 3.2 Estimates of factor demand equations

	Hired labor	Family labor	Total labor	Machine use	Fertilizer
Variable	(1)	(2)	(3)	(4)	(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)
# members 14-64 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	<u>-</u> 0.153	-0.050	-0.07 4	0.144	-0.016
	(-0.756)	(-0.423)	(-0.752)	(0.911)	(-0.074)
Head's education	0.013	-0.005	<u>-</u> 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).

¹¹ In other words, regressions suggest that a 10.0 percent increase of the agricultural (nonagricultural) wage rate would result in an increase of machine use of 2.9 percent (3.1 percent), an increase in the use of fertilizer by 5.6 percent (1.0 percent but insignificant), and a 1.2 percent (9 percent) reduction in total labor use, respectively.

4. EXPLORING CHANGES IN AND DETERMINANTS OF LABOR MARKET FUNCTIONING

We test more directly whether labor market imperfections caused the observed relationship by exploring efficiency differences between family and hired labor, which is indicative of 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 nonagricultural labor demand. Although nonseparability is rejected in the aggregate, this is not the case in villages that have very little nonagricultural labor demand, supporting the importance of nonagricultural labor markets in attenuating the negative farm size—productivity relationship over time.

Differences in Efficiency Between Own and Hired Labor

Wage workers requiring supervision by family members have 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

$$\ln L_{iit} = \alpha_i + \gamma_1 \ln A_{iit} + \gamma_2 \ln w_{ii} + \gamma_3 \ln P_{it} + \gamma_4 H_{iit} + \gamma_5 D_t + \gamma_6 \ln M_{iit} + \varepsilon_{iit}, \tag{3}$$

where $\ln L_{ijt}$ is the logarithm of labor used in crop production. 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 labor and hired labor: a significantly negative (positive) value would imply that family labor is more (less) efficient than hired labor. We again estimate equation 3 for the early (1982–1999) and late (1999–2008) periods separately. Because data for the latter period allow us to differentiate preharvest from harvest labor, we did a robustness check by including both the preharvest family labor ratio and the postharvest family ratio as explanatory variables. In these regressions, household fixed effects allow us to control for unobserved household-level attributes, including ability. The family—to—total labor ratio, M_{ijt} , will likely be affected by endogenous household division and simultaneity bias, so that instrumentation is needed (Benjamin 1995). We use demographic variables as instruments and also instrument A_{ijt} by the amount of land inherited by a household.

Table 4.1 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 (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 allow for easier monitoring. Test statistics in the bottom of Table 4.1 suggest that, for the early period, IVs fail to pass the overidentification 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, 12 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 preharvest and harvest seasons (column 6). 13 A plausible explanation for this finding is that, with wage

12

¹² 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 4); separability holds for the late panel period but not the early panel.

¹³ F-tests fail to reject the null hypothesis that the ratios of family labor to total labor for preharvest and harvest seasons are

growth, the mix of inputs for agricultural production shifted from labor toward 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.

Table 4.1 Test for efficiency differences between own labor and hired labor

	1982–1999)		1999–2008		
	FE	FE (IV) ^a	FE	FE (IV) ^a	FE	FE (IV)a
Variable	(1)	`(2)	(3)	`(4)	(5)	`(6)
Log of land	0.348***	0.420***	0.364***	0.350***	0.341***	0.357***
3	(16.181)	(14.645)	(16.681)	(16.084)	(17.047)	(11.480)
Agricultural 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)
Nonagricultural 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	Ò.151** [*]	0.152** [*]	Ò.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 of males 14-64 years old	-0.033		-0.090		-0.106	
·	(-0.359)		(-0.865)		(-1.115)	
Share of females 14-64 years						
old	0.108		0.035		0.034	
	(0.934)		(0.284)		(0.299)	
Share of males > 64 years old	0.024		0.018		0.003	
·	(0.114)		(0.088)		(0.013)	
Share of 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					-0.045	-0.314
(preharvest)					(-1.273)	(-0.561)
Family-to-total labor ratio					0.006	0.359
(harvest season)					(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)
No. of 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
No. 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-value		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. ^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. FE = fixed effects; IV = instrumental variable.

jointly insignificant for the regressions without IVs (column 5) and with IVs (column 6) at any conventional levels.

Testing for Separability Between Labor Supply and Demand Decisions

With frictions in labor markets, households' labor endowment will be a key determinant of total labor use, resulting in nonseparability between farm households' consumption (labor supply) and production (labor demand) decisions. An alternative to the above is that farmers who may have been rationed out of nonagricultural labor markets in the first period may have been able to overcome barriers to their 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

$$\ln L_{iit} = \alpha_i + \delta_1 \ln A_{iit} + \delta_2 \ln w_{ii} + \delta_3 \ln P_{it} + \delta_4 H_{iit} + \delta_5 D_t + \delta_6 n_{iit} + \varepsilon_{iit}, \tag{4}$$

where $\ln L_{ijt}$ 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 nonfarm 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 elder male adult in household, and share of elder female adult in household; and α_i is household fixed effects. 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 preharvest labor demand from total labor demand. In 1982–1999, household size had a significant impact on total labor demand (Table 4.2). Although few other coefficients are individually significant at conventional levels, the joint test at the bottom of Table 4.2 points toward rejection of the hypothesis of household characteristics, including size, female headship, and age composition, not having affected use of labor for crop production in the preharvest stage and overall (columns 1 and 2).

The relationship disappears for the second period (columns 3 and 4 of Table 4.2), 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 4.2, suggest that the null hypothesis of all coefficients equaling zero can be rejected at 5 percent in the first but not the second period, supporting the notion that labor market functioning did improve over time.

Results for 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 nonagricultural wages. Where they are significant, coefficients on fertilizer and bullock prices have the expected positive signs. Demand for preharvest 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.

Table 4.2 Test for labor market efficiency for preharvest and total labor

	1982	2–1999	1999	1999–2008		
	Preharvest	Total	Preharvest	Total		
Variable	(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)		
Nonagricultural 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 of males 14–64 years old (α ₂)	-0.091	-0.161*	-0.094	-0.113		
	(-0.962)	(-1.696)	(-0.903)	(-1.196)		
Share of females 14–64 years old (α_3)	0.047	-0.007	0.035	0.036		
	(0.396)	(-0.061)	(0.280)	(0.317)		
Share of males > 64 years old (α ₄)	0.021	0.019	0.019	0.006		
	(0.099)	(0.089)	(0.091)	(0.030)		
Share of 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		
No. of farmers	4,925	4,925	4,439	4,439		
Joint F-test $(\alpha_1 + \alpha_2 + \alpha_3 + \alpha_4 + \alpha_5 = 0)$	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 error is adjusted for clustering effect at the village level. All prices are in logarithms.

Although results from estimating equation 4 allowed us to reject the hypothesis of labor market imperfections in 2008 for the total sample, the negative relationship between farm size and productivity did not disappear entirely in the second period (see Figure 3.1 and Table 3.1). If labor market imperfections caused this, it should be possible to find evidence of some imperfections. Given wide variation in nonagricultural labor demand across villages, we would expect this variation could cause differences in labor market performance across villages. To explore whether this 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 of earlier results along this dimension, using the 1999–2008 panel. ¹⁴ Results shown in Table 4.3 are for specifications in which we interacted total household size (columns 1 and 2) and other

¹⁴ 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.

demographic variables (columns 3 and 4) with indicator variables for high, medium, and low levels of village level activity in nonagricultural labor markets.

Table 4.3 Exploring heterogeneity in labor markets

	Preharvest labor	Total labor	Preharvest labor	Total labor
Variable	(1)	(2)	(3)	(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***
Lon of nonemia divinal ways	(-4.005)	(-4.514)	(-4.060)	(-4.582)
Log of nonagricultural wage	-0.088* (-1.786)	-0.197*** (-4.348)	-0.082* (-1.666)	-0.192*** (-4.230)
Price of fertilizer	0.155***	0.152***	0.153***	0.150***
THOSE OF ISTURES	(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
Log of bougghold size * High (v.)	(1.369)	(0.809)	(0.671)	(0.166)
Log of household size * High (γ ₁)	-0.039 (-0.735)	-0.044 (-0.916)	-0.053 (-0.883)	-0.076 (-1.360)
Share of prime male adult * Low (α₂)	(-0.733)	(-0.910)	-0.071	-0.138
chare of prime male addit. Low (d2)			(-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 (0.166)	0.067
Share of prime female adult * Medium (β ₃)			(–0.166) –0.023	(0.367) -0.036
Share of prime female addit Medidiff (ps)			(-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)
Share of elderly male adult * Medium (β ₄)			-0.134	-0.029
Observation of algorithms and a short think (c.)			(-0.471)	(-0.110)
Share of elderly male adult * High (γ ₄)			-0.213	-0.322 (-1.211)
Share of elderly female adult * Low (α ₅)			(–0.738) –0.118	-0.080
onate of clashy ternale addit. Low (43)			(-0.362)	(-0.266)
Share of elderly female adult * Medium (β ₅)			-0.295	-0.358
. ,			(-1.095)	(-1.438)
Share of elderly female adult * High(γ ₅)			0.033	-0.030
	7 400	7 400	(0.116)	(-0.115)
No. of observations	7,400 0.217	7,400 0.251	7,400 0.218	7,400 0.251
R-squared Joint F-test for low level	0.217	0.201	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.

Estimated coefficients provide strong support for the hypothesis that low levels of nonagricultural labor demand are a key driver of observed labor market imperfections and, most plausibly, of the inverse relationship between farm size and productivity. Although the interactions indicate that household size is insignificant in villages with more active nonagricultural labor demand, the coefficient on household size interacted with low level of nonfarm labor market activity is highly significant for preharvest and total labor (columns 1 and 2 of Table 4.3). The result is robust to the addition of more indicators of household composition (columns 3 and 4). In addition to one of these indicators (share of elderly male adults) being significant for the group with low level of nonfarm labor market activity, results from 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 nonagricultural productivity have long attracted interest in the literature (Hsieh and Klenow 2009), differences of equal magnitude exist in agricultural production (Adamopoulos and Restuccia 2014a). Though based on cross-sectional evidence, these differences point toward factor market imperfection, especially in land (Restuccia and Santaeulalia-Llopis 2015) and policy (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 nonagricultural demand provided the basis for separability in the second period, possibly by allowing farmers who had earlier been rationed out to participate in such markets.

Our main finding that, over the 25 years considered, the inverse farm size–productivity relationship in India weakened in response to better functioning of labor markets provides an economically meaningful explanation of an empirically robust phenomenon with clear policy implications. Where, with low capital and knowledge intensity of farming, redistributive land reform could be implemented, and it has helped increase agricultural productivity and investment by small farmers. However, if market imperfections, the nature and severity of which are likely to change over time, also have a significant impact, then a response that carefully considers historical context and weighs advantages and disadvantages of different options will be needed. 16

In the Indian context, although better labor market functioning can explain gradual disappearance of the inverse relationship, labor productivity diverges widely across producers, partly due to administrative restrictions on the operation of land rental markets that had been in place throughout the period. To the extent that our results relate to functioning of factor markets in general, measures to improve functioning of land markets by amending regulations that currently ban leasing in many Indian states will likely have a significant impact on agricultural productivity, ¹⁷ especially if combined with efforts to reduce the likelihood of fraud by increasing the currency of land ownership records.

¹⁵ 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, Songqing, and Yadav 2013) and land market functioning (Deininger, Jin, and Nagarajan 2008).

¹⁶ 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.

¹⁷ 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).

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