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Incomplete Contracts and Investment:

A Study of Land Tenancy in Pakistan

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Abstract

When contracts are incomplete, relationship specific investments may be underprovided due to the threat of opportunistic expropriation or holdup. This paper finds evidence of such underinvestment on tenanted land in rural Pakistan. Using data from households cultivating multiple plots under different tenure arrangements, the paper shows that land-specific investment is lower on leased plots. This result is robust to the possible effects of asymmetric information in the leasing market. Greater tenure security also increases land-specific investment on leased plots. Moreover, variation in tenure security appears to be driven largely by heterogeneity across landlords, suggesting that reputation may be important in mitigating the holdup problem.

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1 Introduction

Relationship specific investments are important in a wide variety of economic transactions. As has long been recognized, when contracts are incomplete, specific investments may be undersupplied due to the threat of opportunistic expropriation by one of the trading partners. The holdup problem underlies a number of prominent theories of institutions, in which particular organizational forms, such as firms or governments, are rationalized as means of curbing ex-post opportunism.[1] Che and Hausch (1999) have recently shown that 'cooperative' investments, those that directly benefit the other trading partner – as when a tenant undertakes an improvement of his landlord's property – are particularly susceptible to holdup. In this case, when commitment by the parties not to renegotiate ex-post is impossible, no contract may be able to protect the investor (see also Hart and Moore, 1999, on the issue of commitment).

Despite the centrality of holdup problems in economics, there is remarkably little direct evidence on their quantitative importance, or on the extent to which commitment mechanisms mitigate holdup.[2] This paper begins to fill these lacunae by examining cooperative investment within the classic principal-agent relationship, that between landlord and tenant. In this context, we ask two basic questions: (1) Does the threat of holdup significantly constrain specific investment? and (2) Is lack of committment ubiquitous, or are some relationships characterized by greater committment, and hence a lower holdup threat, than others?

Banerjee, et al. (2002) also emphasize contractual incompleteness in their investigation of the impact of tenancy reform on farm productivity in one Indian state. However, because they lack data on investments potentially subject to holdup, they cannot distinguish the investment channel from other effects of tenancy reform. There are, to be sure, micro-level studies showing that insecure property rights in land create a disincentive to invest (e.g., Besley, 1995; Jacoby, et al., 2002), but without information on the extent of committment this evidence says little about how land *tenancy* affects investment. The setting for our investigation is rural Pakistan, where land-leasing under both crop-sharing and fixed rental arrangements is pervasive. The empirical analysis compares land-specific

[1] See, e.g., Klein, Crawford, and Alchian (1978), Grossman and Hart (1986), Hart and Moore (1990), Williamson (1975), North and Weingast (1989).

[2] Joskow (1987) finds that longer term contracts are more common as the degree of relationship specific investment increases, but in a context where long-term contracting is feasible. His paper does not examine the investment decision *conditional* on contractual form. See also Chiappori and Salanié (2000) for a broad survey of empirical research on contract theory.

investment on owned versus leased land. If landlords cannot credibly commit to long-term contracts, then certain types of land specific investment will be underprovided by tenants, whereas investment in owned land is obviously immune to holdup.

This is not the first paper to compare farming practices on tenanted and owned land, but it is the first paper of its type to study investment behavior. Seminal work by Bell (1977), followed by Shaban (1987), is concerned with static efficiency; i.e., with moral hazard in current production effort that can arise in share-tenancy. Our approach, by contrast, is to examine the use of an input, farmyard manure, that enhances productivity over more than one agricultural season. While manuring improves soil quality over an extended period, it is an extremely labor intensive activity, one that, in rural Pakistan, is virtually never entrusted to a hired worker. Thus, manuring is for all practical purposes a non-contractible investment. Whether manure will be under-applied on tenanted land relative to owned land depends on the extent to which landlords can commit to rewarding the tenant for his investment. In a world of no commitment, tenants will apply manure only to the point were the marginal return in the *current* period (i.e., the period of the contract) equals the shadow price, and this dynamic inefficiency will be common to both share tenants and fixed rent tenants, even though the latter are full residual claimants.

Security of tenure does not appear to be entirely lacking in rural Pakistan. Despite the absence of enforceable long-term contracts, annual tenancy contracts are typically renewed. Consequently, tenants often stay with the same landlord for a number of years, although the duration of tenancy is highly variable. Our evidence will show that a large part of this tenure heterogeneity is generated by differences across landlords in retention policies. This fact allows us to ask whether differences in the degree of commitment lead to differences in investment across tenants. If all tenants are equally insecure, then the duration of tenancy would indicate nothing about a tenant's incentive to invest. The alternative hypothesis, against which we would like to test this null, is that the duration of a tenancy is a signal of the degree of commitment in a relationship, albeit an imperfect one. Longer durations should therefore be associated with greater investment.[3]

Our main econometric challenge lies in dealing with the endogeneity of tenurial arrangements. Past work has stressed the importance of controlling for unobserved characteristics of the tenant so as to avoid the problem of self-selection into different contract types. Thus,

[3] Laffont and Mattousi (1995) find that longer duration *share* contracts are associated with higher farm *output* in Tunisia. However, because they do not examine investment directly, nor (crucially) allow contract duration to affect output of fixed renters, one cannot use their results to distinguish the holdup problem from moral hazard in effort. Their empirical approach also makes more restrictive assumptions than ours in dealing with the endogeneity issues.

for example, tenants may invest less than landowners merely because they are less wealthy. The advantage of looking at land tenancy in India (Bell, 1977; Shaban, 1987)–and, as it happens, in Pakistan–is that a given household will often cultivate two or more plots of land under different contractual arrangements, allowing a comparison of behaviors across plots *within* the same household. A limitation of this approach is that it does not deal with the possibility that the decision to lease out a plot (or to lease it out under a fixed rent versus a sharecropping contract) may depend on the unobserved attributes of the plot, such as would be the case if there is adverse selection in the leasing market. We use instrumental variables to address this potential endogeneity problem. Our identification argument, developed below, is that the assets of the *owner* of a given cultivated plot, which our data set provides even if the plot is leased in, should determine the leasing status of the plot but not its unobservable type.

The next section of the paper provides the context for our study and sets out a simple two-period model of tenancy, moral hazard, and land-specific investment. Section 3 presents the estimation and identification strategy as well as our evidence for a leasing effect on investment. The analysis of the tenure security effect follows in section 4, along with conclusions and implications in section 5.

2 Tenancy and Land Specific Investment

2.1 Data and context

Ownership of agricultural land is highly concentrated in rural Pakistan, where about half the population is landless. As a result, land lease markets are quite active. According to the latest agricultural census (2000), about a third of total cultivated area was tenant operated, mostly (roughly two-thirds) under crop-sharing arrangements with the remainder under fixed rent. Land and tenancy reforms have been attempted at various times, but they have been largely ineffectual, leaving tenant cultivators with little legal recourse in the event of eviction.[4] Land ownership, however, is clearly established in rural Pakistan and the risk of expropriation by the state or by powerful individuals is negligible.

Our empirical analysis draws upon a new nationally representative rural household survey. The Pakistan Rural Household Survey (PRHS), which was completed in early 2002,

[4] Nabi (1986) reports data from a small 10 village survey in Pakistan's Punjab in which 46 percent of share-tenants said that their landlords could evict them 'easily' and 32 percent said that they or their relatives had experienced eviction within the last two years. At the same time, however, most of these tenants had been leasing land of the same landlord for more than 10 years.

collected data from about 2,800 households sampled across 17 districts and 150 villages. Roughly 60 percent of the households surveyed were farm households and a considerable fraction of these operated multiple plots. The survey was designed to provide detailed information at the plot level on land characteristics (soil type, irrigation, and so forth), land tenure (including characteristics of tenants and owners and details of the contracts), and production activities. These data were collected for the two main agricultural seasons, *kharif* (May-November) and *rabi* (November-May). The main cash crops, cotton, rice, and sugarcane, are grown in *kharif*, while the main food crop, wheat, is grown in *rabi*.

To provide evidence on the holdup problem, we want to focus on an activity that is, in the first place, at least partly an investment. Moreover, we would like this investment to be relationship specific and noncontractible. Last, but not least, it must be relatively easy to measure. Farmyard manure (FYM) meets all of these criteria.

FYM, composed largely of cattle dung, provides variable amounts of the three principal soil nutrients, nitrogen, phosphorous, and potassium. Equally, if not more, important than its role as a nutrient source, FYM improves the quality of the soil by increasing aeration, water retention, structure, and ability to retain nutrients (Government of Pakistan, 1997; Gaur, 1992). Further, these benefits of FYM are long-lasting.[5] Extended field trials in India cited by Gaur (1992) show that the marginal effects of FYM on grain yields persist for at least three years following the initial application.[6] Chemical fertilizers, meanwhile, are used extensively in Pakistan, but they leach from the soil relatively quickly and hence their productivity effects are essentially limited to the season of application.

Because FYM, once applied and incorporated into the soil, is not portable, it is a relationship specific investment in the purest sense. By contrast, investments like experimentation with new farming techniques or seed varieties, which have an aspect of general human capital, can be transferred, at least in part, to another landlord-tenant relationship. Still, there are other, larger scale, purely specific investments that farmers undertake in rural Pakistan, such as constructing irrigation and drainage canals, clearing land, and digging wells. However, these fixed investments are, for the most part, contractible; the tenant or some other party can be paid to do them for the landlord. It is, to be sure, an interesting question as to whether contractible investment is underprovided on tenanted land, but it is not one that speaks directly to the holdup problem.

[5] This feature of FYM has been exploited by Jacoby, Li, and Rozelle (2002) in their study of the incentive effects of land expropriation in rural China. Besley (1995) also includes manuring among the investments he examines.

[6] The marginal effect of FYM in succeeding cropping seasons averaged 50-63% of the effect in the inital season over the three years of the experiments.

The attraction of FYM, from our perspective, is that it is noncontractible. Farmers rarely purchase FYM, but rather typically collect it as a by-product of their own livestock, load it onto carts or donkeys, transport it to the plot and then spread and incorporate it into the soil.[7] This process is extremely labor intensive and thus costly to monitor. The extent and quality of FYM application is also not easily verified or observed ex-post, since weather and other sources of exogenous uncertainty make it difficult to extract this information from realized output alone.

Direct evidence on the noncontractibility of FYM application can be found in the farm wage employment section of the PRHS, where agricultural jobs held over the past year are enumerated by task. A miniscule 0.08 percent of all agricultural wage labor days fall under the category "collecting/spreading farmyard manure".[8] By way of comparison, about 7 percent of labor days are classified as "weeding", another seemingly hard to monitor task, whereas two-thirds of all paid labor days in agriculture are devoted to harvesting various crops. Harvesting work is more readily contractible because productivity per hour is relatively easy to observe.[9]

Noncontractibility in itself does not preclude an efficient level of investment. If the landlord is able to commit fully to retaining the tenant for the duration of the investment, a contract that makes the tenant the residual claimant to the returns on his investment would induce optimal investment. The landlord could then simply remove any surplus accruing to the tenant by appropriately increasing his rent. The classic holdup problem arises only when the investment is noncontractible *and* the landlord is unable to commit. The landlord stands to gain, in this case, by renegotiating the contract terms or contract renewal ex-post. The tenant, aware of this potential for opportunism on the part of the landlord, knows that he is unlikely to recoup all the fruits of his investment. It is thus privately optimal for the tenant to underinvest. The remainder of this section sketches out a simple model to formalize this intuition.

[7] Recommended quantities of FYM vary greatly by crop–at the high end, 10-15 cartloads (8000-12000 kg) for an acre of sugarcane, and even more for vegetables (Government of Pakistan, 1997).

[8] We do not have information from the survey that would allow us to compute the proportion of *family* labor used for this task. However, data from northeast China, which has an agricultural technology not disimilar from that of rural Pakistan, indicates that 8 percent of annual family farm labor goes into FYM application (Jacoby, Li, and Rozelle, 2002).

[9] Indeed, according to the data, harvesting jobs are nearly twice as likely as weeding jobs to pay piece rates, which rely on directly observing output or effort.

2.2 Land specific investment under full commitment

To provide an organizing framework for our empirical work, we incorporate investment into a standard limited liability model of land tenancy.[10] While limited liability on the part of the tenant is just one of several ways to obtain sharecropping as a possible optimal contract (risk aversion, ex-ante financial constraints, and double-sided moral hazard are others), the implications for investment are similar across these tenancy models.

Each landlord owns a single unit of land that he cannot self-cultivate and therefore must give over on lease to one out of a large population of tenants. The tenant is the sole provider of two inputs that are unobservable to the landlord, current production effort and investment. Although our empirical work considers a recurrent investment, we strip the model down to two periods. Investment, $m \in [0, \overline{m}]$, takes place only in the first period, but yields returns in both periods. Since these returns are embodied in the landlord's property, the investment is cooperative in the sense of Che and Hausch (1999). We ignore discounting and depreciation; given our restrictions on the technology, the latter assumption is innocuous. Effort, $e \in [0, \overline{e}]$, is undertaken in both periods. The production function $f(e, m)$ is increasing and concave in its arguments. Output, $Y = f(e, m) + \varepsilon$, depends on an additive shock ε with bounded support such that $Y \in [0, \overline{Y}]$. The tenant's cost functions $c(e)$ and $q(m)$ are increasing and convex in e and m.[11]

We take the landlord and tenant to be risk-neutral and consider linear contracts of the form $sY - r$, where s is the output share of the tenant and r is his fixed rental payment, which can be negative. Tenants have an exogenously given opportunity cost that determines their participation constraint and an exogenous pre-contract wealth that determines their limited liability constraint (i.e., the maximum they can be made to pay in any state of nature).

When the landlord can fully commit to a two period contract, the tenant's optimality conditions are

$$sf_e(e, m) = c_e(e) \qquad \text{and} \qquad 2sf_m(e, m) = q_m(m), \tag{1}$$

the factor of 2 arising in the second equation from the fact that the tenant reaps a return on his investment in both periods. Meanwhile, the first-best – the effort and investment

[10] Banerjee et al. (2002) also discuss this type of investment model. Other papers that have modelled tenancy in a limited liability setting include Shetty (1988), Basu (1992), and Mookherjee (1997), to name a few.

[11] For technical convenience, these costs functions are also assumed separable from one another. Otherwise, the model would have to be solved recursively starting from the second period.

levels that maximize total surplus $2[E(Y \mid e, m) - c(e)] - q(m)$ – solves

$$f_e(e,m) = c_e(e) \qquad \text{and} \qquad 2f_m(e,m) = q_m(m). \tag{2}$$

Clearly, the first-best is achieved only when $s = 1$; that is, when the lease is given on fixed rent. A landlord who can fully commit is essentially 'selling' the property rights on land to the tenant for the duration of the contract. As residual claimant, the tenant is fully incentivized and both moral hazard problems consequently disappear. However, it can be shown that the landlord will only offer a fixed rent contract to those tenants with sufficiently high wealth. If the tenant cannot afford the fixed rent, then the landlord faces a tradeoff between production efficiency and surplus extraction; as a result, he may offer the tenant a share contract, $s \in (0,1)$. Crop sharing gives rise to the familiar 'Marshallian' inefficiency, in which current production effort and, in our model, investment are provided below their first best level.

2.3 Tenure Insecurity and Investment

When the landlord can only commit to a one-period contract, the tenant's optimality conditions become

$$sf_e(e,m) = c_e(e) \qquad \text{and} \qquad sf_m(e,m) = q_m(m), \tag{3}$$

so the marginal return on investment is half as large as in the full commitment case. This is true regardless of whether the tenant is the residual claimant; a fixed rent contract does not eliminate the dynamic inefficiency. Under a share contract, however, the marginal return on investment is even lower than under a fixed rent contract. From these results it easily follows that

Proposition 1 *If investment is not itself contractible and the landlord cannot commit to a tenancy contract that lasts at least as long as the duration of the investment, then the tenant will undersupply land specific investment, even under a fixed rent contract.*

In practice, this holdup problem may be mitigated by reputational effects. Specifically, a landlord may be reluctant to milk his reputation by reneging on his tenant, realizing that if he does so he will be 'punished' by never getting optimal investment on his land. Of course, reputational equilibria are sensitive to a number of assumptions, such as the extent of information costs, on which we need not dwell (see, e.g., Kreps, 1990). The point is that the existence of reputational heterogeneity provides us with an additional testable implication. Returning to the model, assume that the tenant, knowing his landlord's reputation, believes his contract will be renewed in the second period with probability θ. Full commitment can then be interpreted as the case where $\theta = 1$, and no commitment as the case where $\theta = 0$. Clearly, $\theta > 0$ will increase investment relative to the no commitment case. We thus have

Proposition 2 *As the degree of tenure uncertainty increases, the tenant will reduce his land specific investment.*

There are other potential sources of heterogeneity in tenure security besides the retention policies of landlords. Tenants may have different search or moving costs or face different distributions of outside opportunities. In these cases, we may find that some tenants underinvest relative to others even if all landlords can fully commit to long term contracts. It is, in effect, the tenant here who cannot commit to staying long enough to recoup his investment. However, in the environment that we study, heterogeneity in turnover rates induced by tenant characteristics is unlikely to be of great importance. We return to this issue in the next section.

3 The Leasing Effect

3.1 Econometric specification and identification

Our test of Proposition 1 is based on a regression of per-acre FYM use, M_{ci}, by cultivator c on plot i

$$M_{ci} = \alpha L_{ci} + \beta\prime X_{ci} + \nu_c + \varepsilon_{ci}, \tag{4}$$

where L_{ci} is an indicator of whether the plot is leased and X_{ci} is a vector of exogenous plot characteristics.[12] The mean zero error term ν_c captures the effects of unobserved factors common to a given cultivator; e.g., prices, wealth, access to credit, risk aversion and the discount rate, farming knowledge, average land quality, and, importantly, the household's available stock of FYM and other farm assets. The plot-specific error term ε_{ci} reflects

[12] Shaban (1987) states the precise restrictions on technology necessary to move from the tenant's first-order conditions under different tenure types to a regression of input intensity on tenure type. We suppress an analogous discussion here for the sake of brevity.

measurement error in FYM use and, potentially, unobserved attributes of the plot.[13]

Most models of contractual choice in agriculture would imply that L_{ci} depends on at least one of the abovementioned factors captured by ν_c. To deal with this endogeneity problem, we follow earlier work in the literature by restricting attention to owner-cum-tenant (OCT) households. Suppose, for the sake of exposition, that we have a sample consisting of two-plot households, one plot of which is leased in ($L_{ci} = 1$) and the other which is owned ($L_{cj} = 0$). In this case, the first-differences estimator and the fixed effect estimator, the one that we actually use in the empirical work, are numerically identical. Differencing equation 4 across plots within a household, yields

$$\Delta M_c = \alpha L_{ci} + \beta\prime \Delta X_c + \Delta\varepsilon_c \tag{5}$$

where Δ is the difference operator (note: $\Delta L_c = L_{ci}$). The OLS estimate of α from this regression is consistent provided that $E\left[L_{ci}\Delta\varepsilon_c\right] = 0$. Note, the fact that owner-cum-tenant households might be a selective sample (in the sense that $E\left[\nu_c|OCT\right] \neq 0$), does not affect the estimates of 5 because ν_c differences out of this equation.

Given a sufficiently rich set of observed plot characteristics X_{ci}, the identifying assumption $E\left[L_{ci}\Delta\varepsilon_c\right] = 0$ should hold to a reasonable approximation. However, plot quality or fertility is not easy to assess in survey data. Unobserved plot fertility could be correlated with leasing choice due to adverse selection in the leasing market. If a component of plot fertility is private information to the landowner, then the landlord may have an incentive to lease out, rather than cultivate, plots of a certain type.[14] Endogeneity bias induced by the presence of unobserved plot quality has not, to our knowledge, been addressed in past work of this genre. Shaban (1987), for example, controlling for essentially the same plot characteristics that we do (plot value, irrigation, and soil type), assumes that leasing decisions are uncorrelated with plot-specific unobservables.[15]

[13] We eventually also include in the regression a dummy for whether the plot is leased in on fixed rent versus on share, but this does not introduce any new econometric issues at this stage.

[14] A similar effect could arise even if plot fertility is not private information. Suppose that certain plots are more sensitive to tenant abuse (e.g., soil degradation) than others. Landlords may be reluctant to lease out their more sensitive plots (see Allen and Lueck, 1992, for a tenancy model along these lines). If plot sensitivity is also correlated with FYM use, then we again have an endogeneity problem. All the econometric issues that arise with adverse selection apply with equal force to this moral hazard problem.

[15] Besley's (1995) study of property rights and land-specific investment in Ghana does use a household fixed effects/instrumental variables estimator similar to ours, but the instruments are intended only to deal with measurement error in land rights. He, therefore, also assumes that plot quality is fully captured by a limited set of observable characteristics. Likewise, Pender and Fafchamps (2001), who use instrumental variables to test for Marshallian inefficiency in Ethiopia, do not consider the implications of asymmetric information in the leasing market.

What is needed are instruments correlated with contractual choice, but uncorrelated with unobserved attributes of the plot, and which also vary across plots within the same household. Such variables are typically hard to come by in farm household surveys. In our data set, though, tenant cultivators report the characteristics of their landlord, including among other things his total landholdings. So, we have data on the plot owner's landholdings regardless of whether the plot is leased in or owner-cultivated. Since farmers endowed with relatively more land tend to lease out a greater proportion of it, a given plot is more likely to be leased in if it belongs to a larger landowner. Our principal instrument for $\Delta L_c = L_{ci}$ in equation 5 is therefore $\Delta A_o = A_{oi} - A_{oj} = A_{oi} - A_{cj}$, where A_{oi} denotes the total landholdings of the owner of plot i. For owner-cultivated plots ($L_{ci} = 0$) this will just be the landholdings of the cultivator himself (i.e. $A_{oj} = A_{cj}$).[16] Identification also requires that $E\left[\Delta A_o \Delta \varepsilon_c\right] = 0$. In other words, the difference in the total landholdings of the owner-cum-tenant and the landlord of his leased plot should not be correlated with differences in the unobserved attributes of the owner-cultivated and leased plots. We show why this should be the case in the next subsection.

Using the same argument, additional instruments can be constructed from other assets of the landowner (tubewells and tractors) as well as by allowing the impact of relative land endowments on the leasing decision to vary by the landholdings of the tenant-cultivator (interacting ΔA_o with the land endowment of the cultivator A_c). Note that the *direct* effect of the cultivator's landownership on his input use has already been removed from equation 5 along with the fixed effect ν_c.

3.2 Adverse selection and landholdings

To reiterate, our identification strategy requires zero correlation between the unobserved quality of a plot and the landholdings of its owner, conditional on the household fixed effect. Adverse selection per se is not a problem for us in this regard. If a plot's quality is private information to its owner, and some fraction of each landowner's plots are 'good' and the rest 'bad' quality, then bad plots would be leased out and good plots self-cultivated, as the latter could never command a rent commensurate with their productivity.[17] While leased-in plots would thus be of lower average quality than owner-cultivated

[16] Note that, while landowner characteristics may be endogenous in the first-stage contractual choice regression due to landlord-tenant matching (as argued recently by Ackenberg and Botticini, 2002), a correlation between instruments and error term in a *first-stage* regression does not invalidate identification of the second stage regression parameters.

[17] Generalizing to a continuous distribution of quality, there will be a criticial quality level above which the owner will not want to lease out the plot (see Wilson, 1980, for an exposition).

plots, this quality differential would not depend on the difference in landholdings across the plots' owners; all leased plots would be of the same (bad) quality.

What could create an identification problem is that the *degree* of adverse selection in the leasing market might depend on the land endowment of the lessor. Suppose, in particular, that there are diseconomies of scale in self-cultivation over the relevant range of landholdings due, for example, to the cost of supervising hired labor. At the margin, it may pay for a landowner to lease out some of his good plots in this case, even at the disadvantageous market equilibrium rent. Since larger landowners would find it necessary to lease out a greater *proportion* of their good plots, there could be a positive correlation between a plot's quality and the landholdings of its owner.

The question, of course, is whether this argument is empirically relevant. We can directly check this because we have data on the total area cultivated by each tenant's landlord. In the nearly 800 tenancies for which these data are complete, two-thirds of the landlords cultivate no land at all and about 90 percent cultivate less than half of their land. The average landlord cultivates 11 percent of his owned area, a figure that rises to just 15 percent for 'small' landlords (owning less than 100 *kanals* = 12.5 acres). Such uniformly limited landlord cultivation leaves little scope for variation in the degree of adverse selection.

Nevertheless, if the degree of adverse selection does in fact vary systematically with landholdings, then land rents must adjust in equilibrium; this is because a landlord's holdings are easily observed by potential tenants. Thus, as these tenants realize that small landowners lease out worse land *on average* than large landowners, the rent that they would be willing to pay to a small landowner for an observationally equivalent plot must decline. Land rent should then be positively related to the landholdings of the landlord, conditional on the information that the tenant has about the plot.[18]

Table 1 presents an analysis of land rent (i.e., log of cash + value of in-kind rent) based on the 198 fixed rent tenants in our sample with complete information. We control for plot characteristics, including the value of the plot reported by the *tenant*, as well as for village fixed effects to capture local leasing market conditions. Neither the log of the landlord's landholdings, in column (1), nor the small landowner dummy, in column (2), are significantly related to land rents. Thus, we find no evidence that the degree of adverse selection in the leasing market varies with our key instruments. Which is not to say that adverse selection is unimportant; we still need to correct for its possible influence

[18] The same argument would apply to the terms of sharecropping contracts, but empirically it is much harder to measure the 'rent' accruing to the landlord in this case.

in equation 5. But now we can proceed to our test of Proposition 1 confident of our identifying assumptions.

3.3 Data preliminaries

Starting from the nearly 2400 plots on which there was at least some cultivation during the 2000-01 agricultural year, we drop those of households cultivating a single plot. As shown in Table 3, this leaves us with 1508 plots operated by 593 households for the fixed effects analysis of FYM use. Of these plots, 474 are cultivated by 184 owner-cum-tenant households; i.e., households with at least one owner-cultivated plot and one leased plot. It is exclusively these plots that will identify any contracting effect. Since we want to control for a number of plot characteristics as well, we retain the remaining 1,034 plots to maximize the efficiency of the estimates. The 337 leased plots in this subsample also help in estimating the tenancy duration effect.

Before proceeding, a point about our measure of FYM, which is in kilograms used over the year per *cultivated* acre of the plot. Although all plots in our sample had some cultivation during the year, parts of some plots were left fallow or in some cases a whole plot was uncultivated during one season (usually for lack of irrigation water rather than to rejuvenate the soil). In principle, farmers might apply FYM to fallow areas for the benefit of future crops, though, given that the highest return on manuring is realized in the season of application, this strategy seems inefficient. Nevertheless, to check whether such behavior varies systematically by tenancy status, we experimented with FYM use scaled by the *total* area of the plot, regardless of cultivation. Since the estimated leasing effects differ only at the third decimal place across these two ways of defining the dependent variable, we just report the results based on FYM per cultivated acre in this paper.

The descriptive statistics in Table 3 foreshadow some key results. Among the owner-cum-tenant households, owned plots receive nearly twice as much FYM on average as leased plots, and owned plots are considerably more likely to receive any FYM in the first place. Furthermore, there is no discernible difference in FYM use between sharecropped plots and those leased in under a fixed rental arrangement, although the use rate is somewhat lower on rented plots. While these findings are in line with the predictions of our holdup model, it is premature to draw conclusions based on simple mean comparisons.

Two additional features of the data on FYM use emerge from Table 3, each with implications for our estimation procedure. First, the data are heavily censored at zero; almost two-thirds of plots receive no FYM (partly this is attributable to the carryover effect, which may make it uneconomical to apply every year). Second, for those plots on which FYM is applied, there is a great range of variation in the amount per acre. We thus experimented with various transformations of the dependent variable. Among those that can handle zero values, we found that a simple one, $\log(M + k)$, works reasonably well in terms of fitting the data. Unfortunately, in the presence of zeros, there is no way to let k be a free parameter in the estimation, so the choice of k is largely arbitrary (we set it to 0.1).[19] In addition, the estimated coefficients are not invariant to the choice of k, which means that these coefficients cannot be used to compute, say, the implied percentage differential in FYM use on leased versus owned plots. Instead of making such quantitative statements, our focus, for now at least, is on hypothesis testing. For this purpose, the logarithmic transformation is perfectly adequate since the t-statistics are essentially invariant to the choice of k.[20]

Our analysis controls for the following plot characteristics: area, location (inside/outside village), access to year-round or seasonal canal irrigation, access to groundwater of varying quality, topography, soil quality, and self-assessed plot value.[21] Details of these variables, means by tenancy status, and coefficient estimates for a particular specification are reported in appendix Table A.1. Note that we do not have data on the exact distance between the farmer's house and the plot; we only know whether the plot is outside his village or not. This could be a concern, given that both the cost of manure application on a given plot and tenancy status might vary by distance to the homestead. However, in our sample, 92% of all cultivated plots are located within the village (93% of owned plots compared to 90% of leased plots). Given the size of most villages in Pakistan, the vast majority of plots are therefore, at most, only a kilometer or two from the farmer's home.[22] Within this radius, transport cost differentials are likely to be minimal. More-

[19] We choose k so that it is at least an order of magnitude less than the minimum value of the dependent variable. Maximum likelihood estimation of k would fail, since the log Jacobian of the transformation of y is $-\log(y + k)$. At $y = 0$, the optimization algorithm would always try to set k as close to zero as possible.

[20] For example, the t-value for the leasing coefficient in row (3) of Table 4 is 5.91. The corresponding t-value is 5.92 when $k = 0.01$ and 5.93 when $k = 0.001$, even though the associated coefficients vary widely in magnitude.

[21] Both owners and tenant cultivators are asked to report plot value in the survey. One might worry that these values partly reflect the extent of past manuring, although this seems quite unlikely given that the effects of FYM are far from permanent. In any event, the exclusion of this variable from the regressions has a negligible impact on the estimates of interest.

[22] We can be more precise by examining data from the 1991 IFPRI survey, which did collect information on distance to plot (but not on whether the plot was outside the village). Virtually all of the households in this survey were included in the 2001 PRHS sample as part of a panel. Among the 914 plots recorded in the 1991 survey, mean distance between household and plot is only 600 meters (median 300 meters).

over, when we exclude the 118 plots located outside the farmer's village from our analysis, the estimates do not change perceptibly, suggesting that the lack of distance information for these plots is not driving our results.[23]

3.4 Main results: The leasing effect

All the FYM specifications in Table 4 include household fixed effects. The leasing coefficient estimate from the basic model in row (1), which assumes that leasing choices are uncorrelated with unobserved plot attributes, is negative and highly significant (p-value < 0.0000). Thus, less FYM per acre is used on a leased plot than on an owned plot with the same observed characteristics and cultivated by the same household. This evidence against full committment is not an artifact of the sample restrictions required by the household fixed effects procedure, namely the exclusion of single-plot households. When we estimate the corresponding household random effects specification on the full sample of 2375 plots, and include an interaction term between the leasing dummy and a dummy for multi-plot households, the coefficient on this interaction term is insignificant (p-value $= 0.483$). While the leasing coefficient estimated from this random effects specification may itself be biased (because $E[L_{ci}\nu_c] \neq 0$ in equation 4),[24] unless the degree of this bias differs across the multi-plot and single-plot household samples, the interaction term coefficient is an unbiased estimate of the *difference* in leasing effects across the two samples.

Returning to Table 4, we next use Honore's (1992) fixed effect tobit estimator to assess the ramifications of severe censoring of FYM use at zero. While this tobit estimator makes limited distributional assumptions (e.g., it does not impose normality), it does not allow us to treat leasing choices as endogenous conditional on the fixed effect. To get estimates comparable to those in row (1) – i.e., marginal effects – we multiply the fixed effect tobit coefficients and standard errors by one minus the observed censoring rate.[25] The results in row (2) of Table 4 show that ignoring corner solutions leads us to understate the evidence against the null, but not by very much; the t-statistic on the leasing coefficient rises from

Moreover, leased plots are, on average, 88 meters *nearer* to the farmer's homestead than owned plots, though this difference is not significantly different from zero.

[23] Using the specification in row (3) of Table 4 as a comparator, we obtain a t-value on the leasing coefficient of 6.10 instead of 5.91 when these 118 plots are excluded.

[24] It is, in fact, biased. The household random effects estimate of the leasing coefficient on the multi-plot household sample is -1.50 (0.277), which (though still significantly negative) is significantly different from its fixed effect counterpart in Table 4 (Wu-Hausman test p-value $= 0.018$). Note that the random effects specifications reported here also include a full set of village dummies.

[25] Strictly speaking, unless one imposes distributional assumptions, the fixed effect tobit does not yield predicted marginal effects.

5.78 in row (1) to 6.48 in row (2).[26]

The data also exhibit enormous variation in the size of farm plots, with the variance of FYM per acre being much higher on the smaller plots. We therefore depart from the homoskedasticity assumption by allowing $var(\varepsilon_{ci}) = \sigma^2 a_{ci}^{-\eta}$, where a_{ci} is plot area and η is estimated along with the other parameters (by maximum likelihood). The results of reestimating specification (1) with η free are reported in row (3). The leasing coefficient is basically unchanged, though the standard error is marginally lower. Even though we can reject homoskedasticity (p-value $= 0.0001$) and the residual variance does, as expected, decline with plot area (i.e., $\widehat{\eta} > 0$), heteroskedasticity is not severe enough in this logarithmic specification to make much of a difference. Nonetheless, we continue to leave η unrestricted for the remainder of our estimation.

Next we relax the assumption that the leasing decision is uncorrelated with unobserved plot characteristics. As discussed earlier, such a correlation could arise if there is asymmetric information in the leasing market leading either to adverse selection or to moral hazard (see fn.14). Either way, the impact on our estimate of the leasing effect depends on whether leased plots are of high or low fertility on average and on how plot fertility influences FYM use. For example, if low fertility plots tend to be leased out and low fertility reduces the return to manuring, then there will be a negative correlation between the leasing dummy and the error term, whereas this correlation will be positive if low fertility plots have a higher return to manuring. In short, the endogeneity bias in the leasing coefficient could work in either direction.

In addition to the plot owner's landholdings, our instrument set includes a dummy for whether the owner is a small landowner (see Table 1), dummies for whether he owns a tractor and a tubewell, and interactions of these four variables with each other as well as with the total owned area of the plot cultivator. The joint explanatory power of these 12 instruments, conditional on the household fixed effects and the plot characteristics, is high ($F(12, 891) = 89.6$); they also easily pass the overidentification test (see Table 4).[27]

Comparing rows (3) and (4) of Table 4 shows the negligible effect of instrumenting on the the leasing coefficient and, above all, that the IV estimate remains significantly negative. The Wu-Hausman exogeneity test has a p-value of only 0.507, an especially

[26] While most of the variation in FYM use is at the extensive margin, there is also considerable information in the quantity data conditional on use. The estimated leasing coefficient from a fixed effect logit model based on the use/not use decision alone has a t-value of 5.08. However, if we reestimate specification (1) dropping all cases of zero FYM use, we still obtain a t-value on the leasing effect of 3.14.

[27] To be conservative, throughout this paper we use the heteroskedasticity robust version of the overidentification test (see Wooldridge, 1995).

compelling result given the high power of the test.[28] Failure to reject exogeneity of the leasing choice suggests either that asymmetric information in the leasing market is unimportant or that unobserved plot attributes do not significantly affect the returns to FYM use. As we will see, this finding also has implications for our interpretation of the effect of tenancy duration.[29]

4 The Tenure Security Effect

4.1 Econometric considerations

To test Proposition 2, we relax the restriction in equation 5 that the impact of leasing on investment is the same for all tenants (i.e., a common α). Thus, let μ_{oc} be a latent variable representing the degree of tenure security, the dual subscripts indicating that this variable is specific to an owner (landlord)-cultivator (tenant) match. In terms of the model in section 2.3, we may think of μ as the objective probability of contract renewal, in contrast to the subjective (on the part of tenant) probability θ; in other words, θ is the tenant's prior on μ. For μ to matter for investment behavior, the tenant's prior must be informative to some extent.

For owner-cultivated plots, on which tenure security is presumably absolute and unvarying, we set $\mu_{oc} = \mu_{cc} = 1$ without loss of generality. We also make the innocuous normalization $E\left[\mu_{oc}|L_{ci} = 1\right] = 0$, from which it follows that $E\left[\mu_{oc}L_{ci}\right] = 0$. Augmenting equation 5 gives

$$\Delta M_c = \alpha L_{ci} + \gamma L_{ci} \times \mu_{oc} + \beta\prime\Delta X_c + \Delta\varepsilon_c. \tag{6}$$

Our null hypothesis is that $\gamma = 0$; i.e., tenure security does not influence the tenant's investment on his leased plot *vis a vis* his owned plot. The alternative hypothesis is that investment incentives are stronger in more secure tenancies, so that $\gamma > 0$.

[28] In particular, based on an inverse power function calculation (Andrews, 1989), we can be 95% certain that we would have rejected exogeneity if the true leasing coefficient was less than 0.96 in absolute value (instead of 1.95). Notice that, even if the leasing coefficient were this small, we would still have been able to reject the null that it is zero.

[29] Since recommended amounts of FYM vary greatly by crop, it is also worth asking how much of the leasing effect is due to differences in the crops grown on tenanted versus owner-cultivated plots. Controlling in specification (3) for the proportion of cultivated plot area devoted to the eight most important crops (wheat, rice, cotton, sugarcane, maize, *rabi* fodder, sorghum fodder, and vegetables) only reduces the leasing effect to -1.65 (0.352), which suggests that tenants respond to the investment disincentive mainly by using FYM less intensively on every crop.

The test requires information on μ_{oc}, which is not directly observed. But notice that the ongoing duration of a tenancy, d_{oc}, is a (noisy) indicator of the underlying insecurity of tenure. In Jovanovic's (1979) job-matching model, for example, the hazard rate of job separation in the presence of specific investment is a function of *elapsed* job duration and of match quality. If workers (tenants) have different costs of search and firms (landlords) have different retention policies, then the separation hazard also depends on these additional exogenous sources of turnover. Since the hazard rate (which is one minus the contract renewal probability) uniquely defines the distribution of tenancy duration, we may write $\log(d_{oc}) = E[\log(d_{oc})|\mu_{oc}] + \xi_{oc}$, where ξ_{oc} is random 'luck'. Taking a linear approximation to the conditional expectation delivers

$$\log(d_{oc}) = \mu_{oc} + \xi_{oc}, \tag{7}$$

which is exact if elapsed duration is distributed as a Weibull. Equation 7 suggests putting (demeaned) log duration interacted with the dummy variable L_{ci} into equation 6 as a proxy for $L_{ci} \times \mu_{oc}$. It also implies that doing so leads to an errors-in-variables problem, because of ξ_{oc}. Since the resulting estimate of γ is biased toward zero, we will tend to find weaker evidence against the null (that $\gamma = 0$) than if we observed μ_{oc} directly. Of course, this is a moot point if we actually end up rejecting the null.[30]

Tenancy duration may also be endogenous in equation 6 for reasons other than measurement error, as we discuss in detail in subsection 4.3.

4.2 Evidence from landlords

For our test of the tenure security effect to have power, there must be a reasonable amount of variation in tenancy duration. But variation in d_{oc} does not necessarily imply variation in μ_{oc}, or in θ for that matter. To see why, take the case where θ, the subjective probability of retention, does not vary across tenants. Since actual contract renewal is a stochastic process, we would still observe a nondegenerate distribution of elapsed tenancy durations, but no relationship between duration and investment. It is thus important to assess the extent to which variation in tenancy duration reflects heterogeneity in the objective factors underlying μ_{oc}, like tenant mobility and landlord behavior.[31]

To do so, we use data collected from the landlords in our sample. For each plot that

[30] Note that, since $E[\mu_{oc} L_{ci}] = 0$, the estimate of α is robust to measurement error in μ_{oc}.

[31] Match quality, which by definition is not specific to the landlord or to the tenant, cannot be distinguished from the error term ξ_{oc}.

was leased out, the survey asks its owner about the characteristics of the tenant, including his principal assets, and about the duration of the tenancy. For reasons that will become apparent shortly, we focus only on those landlords who report at least two leased plots, which gives us a sample of 345 tenanted plots owned by 127 landlords in 52 villages. The median elapsed tenure in this sample is 4 years and the mean is 6 years.

Column (1) of Table 2 reports a regression of log duration on an array of tenant, landlord, and plot characteristics, and on geographical dummies for each of the 18 *tehsils* (administrative units below a district). Less than a third of the total variance in log(duration) is explained by these covariates. Tenants with larger land holdings appear to be less mobile, staying with a given landlord longer. But tenants' ownership of a tractor or plow has no effect on duration, nor do any of the landlord asset variables.

The regression in the second column of Table 2 includes landlord fixed effects. All the landlord characteristics and *tehsil* dummies consequently drop out. Remarkably, this regression explains 80 percent of the variance in log duration. This means that a landlord-specific unobservable, which is uncorrelated with landlord assets, accounts for at least half of the variation in tenancy durations in our sample.[32]

Evidently, a large portion of the variation in μ_{oc} consists of heterogeneity in the behavior of landlords toward all their tenants taken as a group. One interpretation of this heterogeneity is that there are those who can acquire reputations as 'good' landlords who do not opportunistically evict their tenants and others who cannot. If so, then not only can we use duration data to test proposition 2, but we can interpret any tenure security effect as arising largely from landlord commitment problems.

4.3 Main results: The tenure security effect

Among the 563 tenanted plots in our sample, the median number of years that the same tenant has cultivated the plot is 7; 8 years for share-tenants and 4 years for fixed-rent tenants. Results reported in rows (5) and (6) of Table 4 strongly support a tenure security effect. The negative impact of leasing on FYM use declines significantly with tenure on the plot, whether leasing itself is treated as exogenous (p-value = 0.0015) or as

[32]Note, however, that part of the landlord fixed effect includes landlord means of *unobserved* plot and tenant characteristics. Since these unobservables are likely to have high intra-village correlation, we also compared the fit of a regression with *village* fixed effects to that of one with landlord fixed effects on a sample of 263 plots (95 landlords in 20 villages) for which there were at least two landlords per village. The R^2's are, respectively, 0.351 and 0.784. This suggests that unobserved plot and tenant characteristics that vary across villages account for very little of the landlord effect.

endogenous (p-value = 0.0019). More durable relationships between landlord and tenant, which the evidence in section 4.2 suggests are associated with 'better' landlords, yield greater investment. So, we do not seem to be in a world of no commitment.

A possible alternative explanation for this result, however, is that tenants stay longer on plots of higher quality, and on these plots the return to FYM use also happens to be higher. Why would tenants stay longer on better plots?[33] Consider the adverse selection model again, but with a continuum of unobservable (to the tenant) plot quality. Initially, the tenant does not know whether he has leased in a particularly good or bad plot, although he may know the *average* quality of leased plots. Presumably, the longer the tenant cultivates the land the more he learns about its underlying fertility. Assume, further, that landlords cannot commit to more than one-year contracts. Thus, after the first year of tenancy, the landlord and tenant renegotiate. Suppose that the landlord offers the tenant the same rent as in the previous year. The tenant will only stay on for a second year if his posterior estimate of plot quality is at least as high as his prior (assuming zero cost of changing landlords). Therefore, *holding rent constant*, tenants will remain longer on better plots. Of course, there is no reason to suppose that the landlord will offer the same rent, since he knows that the tenant has learned something about the plot's quality. If the landlord has a relatively good plot, then he would want to raise the rent, but not to the point where the incumbent tenant would leave. This 'informed' tenant is more valuable to him than an uninformed tenant and needs to be given some incentive to stay on. The opposite is true for a landlord with a below average quality plot; he would want to replace the incumbent tenant with an ignorant one. In sum, the model implies: (1) Tenants remain longer on better plots; and (2) Rent rises with tenure.

Our earlier evidence already belies the empirical relevance of this story. Failure to find *negative* endogeneity bias in the leasing coefficient (recall Table 4, row (4)), despite powerful instruments, suggests that the necessary conditions – asymmetric information in the land leasing market and returns to FYM use that increase in unobserved plot fertility – are not *jointly* present in rural Pakistan. Therefore, tenancy duration should not be acting as a proxy for unobserved plot quality in the FYM regressions.

Nevertheless, to deal with the issue directly, we now instrument tenancy duration. Because landlord assets explain little of the observed variation in duration (see Table 2),

[33] Clearly, any plot attribute that is observable to both landlord and tenant prior to contracting should not affect length of tenure. This is indeed borne out in the data. When we run a regression analogous to specification (1) of Table 2, but using the log duration and other variables reported from the tenant's rather than landlord's side, and based on the full sample of tenants ($N = 954$), plot value, canal irrigation, tubewell irrigation, plot topography, and soil type are not jointly significant (results available upon request).

we augment our original instrument set with the leave-out mean of log duration within the village.[34] Average duration in a village should be correlated with plot-specific duration to the extent that the duration of tenancies are determined by village leasing market conditions. Moreover, village-average duration should not be correlated with the unobserved (to the tenant) quality of a given plot; in particular, any plot quality information reflected in village-average duration should be common knowledge in the village. Row (7) of Table 4 reports IV estimates that treat both leasing and tenancy duration as endogenous. Compared to row (6), there is no evidence of *upward* bias; if anything, the reverse seems to be the case, as would be consistent with the measurement error story in subsection 4.1.

We can also check whether implication (2) passes empirical muster using the per acre land rent regression in Table 1. When we introduce the log of tenancy duration into this regression, it attracts a negative but insignificant coefficient (column (3)). Thus, there is no evidence that rents are *increasing* in tenancy duration conditional on the observable characteristics of the plot (and village fixed effects). Using instead a dummy for tenure longer than 3 years (two-thirds of fixed rent tenancies in our sample have lasted at least four years), we again find, in column (4), no significant difference in rent per acre paid by short and long duration tenants. These findings, and additional evidence to be presented later, suggest that tenancy duration does not, to a significant degree, reflect an incumbent tenant's informational advantage regarding plot quality.

4.4 Fixed rent versus share contracts

As we have emphasized, the potential for holdup is present regardless of whether the tenant has taken the land on fixed rent or on a share contract. Although the fixed rent tenant, unlike the sharecropper, is full residual claimant, as long as he cannot be assured of the full return on his land-specific investment, he will also invest less than a plot owner. The final specifications in Table 4 include a dummy for whether the plot is leased on a fixed rent basis. The coefficient on this dummy captures the effect of fixed rental on FYM use *over and above* the average leasing effect. Thus, if holdup is an issue, we expect the *sum* of the leasing and fixed rent coefficients to be significantly less than zero. Further, as we have seen, theory predicts that the fixed rent coefficient will be positive; fixed rent tenants would provide the first-best level of manuring in a world of full commitment, but, regardless of the degree of commitment, share-tenants would always underprovide

[34] We also add the six instruments to be discussed in the next subsection. Village mean duration must be interacted with these 18 instruments in order to contribute to the estimation of the household fixed effects model.

this investment because of moral hazard (the tenant's output share is no greater than 50 percent in the vast majority of our cases).

To help explain the decision to take a plot on fixed rent versus on share, and thus to improve the precision of our IV estimates, we add 6 instruments to our earlier set of 12 (specifically, interactions between two of the plot owner asset variables and three village-level variables: proportion of land sharecropped, proportion of land rented, and gini coefficient for landownership).[35] But even this augmented instrument set does not explain the fixed rent decision very well, relatively speaking; the F-statistic for the excluded instruments in the fixed rent equation is only 27.3, compared to 60.6 in the leasing equation. Nevertheless, whether we use instruments (row (9)) or not (row (8)), we can reject with near certainty the hypothesis that plots leased on fixed rent receive the *same* amount of investment as owned plots (p-value $= 0.0002$ in row (9); < 0.0000 in row (8)). Far less certain is the comparison between fixed renters and share tenants. The uninstrumented results in row (8) suggest that rented plots actually receive less FYM than sharecropped plots (p-value $= 0.022$). After instrumenting, the fixed rent coefficient is practically unchanged but becomes insignificant. If we ignore the endogeneity of leasing decisions (as seems justified), and return to the fixed effect tobit estimator discussed earlier, we obtain the results reported in row (10). Here we again find a negative but insignificant fixed rent effect (p-value $= 0.303$).

The curious finding that share-tenants do not invest any less than fixed rent tenants, *conditional on tenancy duration*, might be explained by a monitoring argument. Given moral hazard in current production effort, it may pay for a landlord to be involved in and to supervise his share-tenant's production activities. Indeed, share-tenants in Pakistan *are* typically heavily supervised (see Nabi, 1986; Jacoby and Mansuri, 2004), either directly by their landlord or by hired labor managers (*kamdars*). If there are economies of scope in supervision, then the cost of monitoring a tenant's investment effort (at least to some degree) is lower in sharecropping than in fixed rental arrangements. Provided that this monitoring effect outweighs the moral hazard effect, share tenants may invest more than fixed rent tenants, although an econometric investigation of this hypothesis is beyond the purview of the present paper. Our main result here is that tenants of either type, unequivocally, invest less than owners.

[35] The first two of these village level variables are computed directly from the full survey sample, whereas the land gini is based on a census of landownership for the entire population of each village. Since these 6 new instruments hardly add any explanatory power to the first-stage regression for the overall leasing decision, we did not use them in earlier IV specifications.

Importantly, the coexistence of fixed rent and share tenancy contracts in rural Pakistan allows us to address a lingering concern about the tenure security effect, which is that it might be an artifact of tenant eviction upon failing to meet an output standard. If landlords use eviction as a threat to elicit higher effort (see, e.g., Banerjee, et al., 2002) or as a way to weed out bad tenants, then longer duration tenants may be precisely those who tend to shirk less. Longer term tenants may then use more FYM merely because it is a complement to their (greater) effort in production.

But this eviction story applies only to share-tenants, as fixed-rent contracts are not contingent upon observed output. Thus, if the tenure security effect in Table 4 arises solely from punitive eviction, it must be due to variation in the duration of share-tenancies alone. In other words, allowing the coefficient on log duration to differ by type of tenancy should result in a coefficient of zero for fixed rent tenants. What we find is just the opposite. Estimating an unconstrained version of specification (8) in Table 4 leads to a log duration coefficient of 0.508 (0.253) for share-tenants and 0.837 (0.416) for fixed-rent tenants, both of which are significantly different from zero, but not from each other (p-value = 0.494). The same modification of the fixed effect tobit specification (10) yields similar conclusions (the corresponding estimates are 1.056 (0.329) and 1.529 (0.644), p-value = 0.468). So, it does not appear that the positive relationship between FYM use and tenancy duration is being driven by landlords using eviction as a sharecropper discipline or screening device.

4.5 Further tests

Our final set of tests concerns the consequences of landlord-tenant matching. One might object to our evidence for a tenure security effect on the grounds that whatever it is that ties the tenant to his landlord could influence not only his FYM use, but his other input decisions as well. In Jovanovic's (1979) model, for example, heterogeneity in match quality determines both job duration and productivity. Long-lived tenant-landlord matches may be more productive in the sense that the marginal products of a whole range of inputs may be higher than those of short duration tenancies. In this case, long duration tenancies may be associated with more intensive use of FYM *and* of other inputs even in the absence of committment problems. Unobserved plot quality, discussed in subsection 4.3, is a special case, implying, as we have seen, that tenants stay longer (and perhaps use more FYM) on more productive plots. Unlike the case of eviction, though, this positive tenure-productivity effect applies with equal force to sharecropping and fixed rental arrangements and is therefore consistent with the findings in the last subsection.

To assess the relevance of the matching argument, we examine chemical fertilizers. Since these fertilizers have negligible carryover across seasons, their use should not be directly affected by investment incentives and consequently by variation in tenure security. A positive relationship between tenancy duration and chemical fertilizer use would thus be consistent with a match duration-productivity link, casting doubt on our interpretation of the analogous finding for FYM use as a tenure security effect.[36]

The two most important chemical fertilizers in Pakistan are nitrogen and phosphate. As shown in Table 3, both of these chemicals are applied to more plots in a given year than is FYM, but mean quantities per acre hardly differ across owned and leased plots. Our econometric approach is essentially the same as in Table 4. Since we find no evidence of heteroskedasticity (i.e., $\widehat{\eta}$ is not significantly different from zero) for both nitrogen and phosphate, we impose the restriction $\eta = 0$ throughout. Also, for the IV estimation, using the same 18 instruments as in Table 4, we reject the overidentifying restrictions for both fertilizers. To get instruments that pass this test, we drop 8 of the interaction terms from the original set, which gives us, in the case of nitrogen, specification (3) of Table 5. In the case of phosphate, no subset of the original instruments manages to pass the overidentification test, so we do not report any IV results for this fertilizer.

Not a single coefficient associated with the form of tenancy or its duration is statistically significant in Table 5; this includes the specifications that deal with censoring of chemical fertilizer use at zero (rows (2) and (5)) and with the endogeneity of leasing choices (row (3)). Thus, we cannot reject the hypothesis that owned, rented, and sharecropped plots all receive the same amount of nitrogen and phosphate fertilizer per acre. This finding is not surprising given that, in most cases, fertilizer costs are shared between landlord and share-tenant at the same rate as is output and, hence, there is no direct Marshallian inefficiency (Shaban, 1987, obtains similar results for fertilizer use).[37] For our purposes, the major finding is that there is no significant association between tenancy duration and the use of chemical fertilizers. As argued, this evidence lends credence to the tenure security interpretation in the case of FYM, since the degree of tenure security

[36] There is, however, a potential indirect effect of tenancy duration on chemical fertilizer use. If chemical fertilizer and FYM are complements (substitutes) in production, then chemical fertilizers would be used more (less) intensively on owned plots and on leased plots with a high degree of tenure security. Given that FYM enhances the efficacy of chemical fertilizers, it is not clear whether these two inputs are, on balance, substitutes or complements. In any case, this indirect effect is unlikely to be very large.

[37] Allen and Lueck (1992) point out that leasers, especially those on fixed rent, may have an incentive to *overapply* chemical fertilizers relative to owners because they do not fully internalize the long-term damage to soil. Evidently, this effect is not important in Pakistan, perhaps because use of chemical fertilizers, at least across the range of variation seen in our data, is still relatively low by international standards.

should have no direct effect on a static input like chemical fertilizer.

5 Conclusion

Although committment and holdup problems lie at the core of incomplete contract theory, their quantitative significance has rarely been assessed directly. This paper finds striking evidence of a holdup threat in a setting where theory suggests it should be prevalent, namely where investment is cooperative (in our case, the tenant's investment benefits the landlord). In rural Pakistan, specific investment is lower on leased land than on owned land cultivated by the same household, even after accounting for the potentially confounding effects of asymmetric information in the leasing market. Further corroboration is found in the fact that land taken on fixed rent receives as little, if not less, investment than land taken on a share basis, a result that is inconsistent with an explanation based on static moral hazard.

We also find, after ruling out numerous alternative interpretations of the evidence, that security of tenure, and hence the incentive to invest, varies considerably across tenancies. This means that the mere potential for holdup does not necessarily lead to an unravelling of all investment. What is the mechanism for restraining ex-post opportunism? Our data show important differences across landlords in their propensity to retain their tenants, suggesting that some landlords might value reputation more than others. One can imagine an equilibrium in which 'good' landlords, who can credibly commit to not appropriate their tenants' investment returns, coexist with 'bad' landlords on whose land tenants never invest. From the policy perspective, the finding that the degree of commitment varies across tenancies means that there is scope for effective tenancy reform. Put differently, if tenure insecurity arises merely from inherent differences in tenant mobility, not from landlord behavior, then legislation binding landlords to long term contracts would do little to encourage investment.

Finally, it is worth asking about the economic significance of underinvestment on tenanted land. Of course, the full extent of the dynamic inefficiency cannot be assessed without accounting for all the different types of land-specific investment. This paper focuses on just a single investment, albeit an important *noncontractible* one, the use of farmyard manure. Nevertheless, it is still instructive to ask, based on our estimates, how yields would be affected by giving land ownership rights to tenants. Fortunately, we have access to results from a large number of agricultural experiments in India compiled by

Gaur (1992) showing the effect of FYM use on the yields of major crops;[38] these should also be approximately valid for neighboring Pakistan. In the year of application, one metric ton of FYM per acre would increase the yields of any of the four major crops (wheat, rice, cotton, and sugarcane) by 1.5-1.75 percent. Based on the leasing coefficient from a linear model otherwise identical to specification (3) in Table 4, leased plots received on average about 0.6 metric tons less FYM per acre than owned plots in the 2000-01 agricultural year. Therefore, by converting a leased plot into an owned plot, major crop yields would rise by about one percent in the first year. Taking into account the carryover effects of FYM to future years (see fn. 6) would give a cumulative yield gain of around 2-2.5 percent. For poor tenant households in rural Pakistan, even modest gains such as these will have important income effects.

To be sure, these are just gross impacts; they do not account for the cost of FYM application, but they do give some sense of the magnitudes involved. Banerjee et al. (2002) find that a tenancy reform in India, one that fell far short of giving share-tenants full ownership rights, increased crop yields on the order of 50-60 percent. At least part of this increase, they argue, can be attributed to higher investment due to improved tenure security. Our findings suggest that investment – more precisely, noncontractible investment – may only be a small part of the story. However, we reiterate, this conclusion rests critically on what other noncontractible investments (e.g., land maintenance activities) tenants can potentially undertake and how responsive these investments are to changes in the security of tenure. Filling in these gaps in our knowledge is an important area for future research.

[38] These estimates are more accurate than could ever be obtained from economic data because they are quite literally taken from experiments, in which everything else is held constant. The experimental effects reported are for FYM use on top of a standard application of chemical fertilizer.

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Table 1: Determinants of Land Rents

	(1)	(2)	(3)	(4)
Log(landlord's landholdings)	-0.022	–	–	–
	(0.048)			
Landlord is small landowner (<12.5 acres)	–	-0.082	–	–
		(0.159)		
Log(duration of tenancy)	–	–	-0.137	–
			(0.083)	
Tenancy duration > 3 years	–	–	–	-0.125
				(0.154)
Log(plot value per acre)	0.247*	0.249*	0.229*	0.232*
	(0.104)	(0.103)	(0.095)	(0.100)
Village dummies: p-value	0.0000	0.0000	0.0000	0.0000
R^2	0.712	0.712	0.721	0.714

Notes: Standard errors adjusted for household-level clustering in parentheses (* denotes p-value < 0.05). The dependent variable is the log of rent (cash + value of in-kind) per acre. Each regression also include the plot characteristics described in Table A.1 (except plot area) and dummies for whether plot is cultivated in kharif or rabi season only. N=198 plots in 64 villages.

Table 2: Determinants of the Duration of Tenancy

	Mean (std. dev.)	(1)	(2)
Tenant characteristics:			
Age	44.9	0.0038	0.0020
	(11.5)	(0.0069)	(0.0053)
Landholdings (acres)	2.62	0.0190*	0.0347*
	(10.4)	(0.0054)	(0.0160)
Own a tractor/plow	0.22	0.030	0.088
		(0.201)	(0.191)
Landlord characteristics:			
Landholdings (acres)	18.9	0.0045	–
	(34.9)	(0.0058)	
Own a tractor	0.081	-0.122	–
		(0.367)	
Own a tubewell	0.087	-0.173	–
		(0.258)	
Landlord related to tenant	0.084	0.141	0.666
		(0.247)	(0.432)
Landlord and tenant in same caste or clan	0.261	-0.410*	-0.782*
		(0.186)	(0.199)
Plot characteristics: p-value		0.0040	0.1824
Landlord fixed effects		No	Yes
R^2		0.308	0.804

Notes: Standard errors in parentheses, corrected for clustering on landlord in column (1) (* denotes p-value$<$ 0.05). The dependent variable is the log of tenancy duration. All specifications include plot characteristics as described in Table A.1. Column (1) also includes tehsil dummies. N= 345 plots (127 landlords).

Table 3: Samples and Descriptive Statistics on Input Use

Sample	N	FYM	Nitrogen	Phosphate
Plots of multi-plot households	1508	1173	38.7	15.8
		(3048)	(35.6)	(20.6)
		[36.7]	[85.5]	[67.6]
Plots of owner-cum-tenants	474	1069	42.6	16.7
		(2721)	(38.5)	(20.5)
		[37.1]	[87.9]	[73.0]
Of these plots:				
Owned	248	1388	43.1	17.1
		(3083)	(39.8)	(23.0)
		[46.2]	[87.3]	[71.8]
Leased	226	714	41.9	16.4
		(2204)	(37.1)	(17.4)
		[26.7]	[88.5]	[74.3]
Of these plots:				
Sharecropped	127	702	32.7	12.9
		(2425)	(32.0)	(15.2)
		[29.9]	[83.5]	[67.7]
Fixed rental	99	730	53.7	20.7
		(1895)	(39.8)	(19.1)
		[23.2]	[95.0]	[82.8]

Notes: Mean (std. dev.) [% nonzero]. All quantities in kilogram per cultivated acre.

Table 4: Household Fixed Effects Estimates of Plot-Level FYM Use

	Leased	Leased on fixed rent	Leased × log(*duration*)	η	Overidentification test (p-value)
(1)	-1.98*	–	–	0	–
	(0.342)			–	
(2)[a]	-2.47*	–	–	–	–
	(0.381)				
(3)	-1.95*	–	–	0.152*	–
	(0.329)			(0.040)	
(4)[b]	-2.15*	–	–	0.152*	0.556
	(0.446)			(0.040)	
(5)	-1.76*	–	0.684*	0.153*	–
	(0.333)		(0.216)	(0.040)	
(6)[b]	-1.91*	–	0.687*	0.153*	0.573
	(0.426)		(0.221)	(0.039)	
(7)[c]	-1.68*	–	0.911*	0.149*	0.261
	(0.352)		(0.453)	(0.040)	
(8)	-1.15*	-1.35*	0.595*	0.159*	–
	(0.426)	(0.587)	(0.219)	(0.039)	
(9)[b]	-1.41*	-1.29	0.592*	0.160*	0.208
	(0.620)	(1.06)	(0.232)	(0.039)	
(10)[a]	-1.84*	-0.844	1.15*	–	–
	(0.579)	(0.819)	(0.287)		

Notes: Standard errors in parentheses (* denotes p-value< 0.05). The dependent variable is log(kg FYM per acre + 0.1). All specifications include 12 plot characteristics (see Table A.1).

[a]Fixed effect tobit with quadratic loss function. Coefficient and s.e. multiplied by 1 - censoring rate.

[b]IV estimates (leasing decisions endogenous).

[c]IV estimates (leasing decisions and tenancy duration endogenous).

Table 5: Household Fixed Effects Estimates of Plot-Level Chemical Fertilizer Use

	Leased	Leased on fixed rent	Leased × log(*duration*)
Nitrogen			
(1)	-0.340	-0.038	-0.029
	(0.311)	(0.430	(0.156)
(2)[a]	-0.487	-0.080	-0.039
	(0.327)	(0.468)	(0.191)
(3)[b]	-0.765	0.332	-0.067
	(0.479)	(0.927)	(0.170)
Phosphate			
(4)	0.187	-0.516	0.181
	(0.400)	(0.553)	(0.200)
(5)[a]	0.104	-0.528	0.185
	(0.484)	(0.524)	(0.205)

Notes: Standard errors in parentheses (* denotes p-value< 0.05). The dependent variable is log(kg fertilizer per acre + 0.01). All specifications include 12 plot characteristics (see Table A.1).
[a]Fixed effect tobit with quadratic loss function. Coefficient and s.e. multiplied by 1 - censoring rate.
[b]IV estimates. Sig. test of excluded IVs in first-stage leasing equation, $F(10, 895) = 87.5$, for leasing on fixed rent, $F(10, 895) = 27.2$. Overidentification test p-value $= 0.299$.

Appendix

Table A.1: Descriptive Statistics and Results for Plot Characteristics

	Means (Std. Dev.)		Regression Coefficients (Std. Errors)		
	Owned Plots ($N = 945$)	Leased Plots ($N = 563$)	FYM[a]	Nitrogen[b]	Phosphate[c]
log(total area)	3.23	3.45	0.427*	0.335*	0.817*
	(1.23)	(1.07)	(0.172)	(0.123)	(0.159)
Outside village	0.068	0.096	-1.56*	-0.180	-1.11*
			(0.51)	(0.380)	(0.489)
Flat topography	0.797	0.886	0.620	-0.258	-0.009
			(0.485)	(0.353)	(0.454)
log(value/acre)	9.49	9.44	0.699*	0.251	0.498
	(1.03)	(1.03)	(0.332)	(0.231)	(0.297)
Canal Irrigation:					
Perennial	0.222	0.455	2.42*	0.531	0.516
			(1.02)	(0.730)	(0.939)
Seasonal	0.210	0.195	2.22*	-0.684	-0.494
			(0.96)	(0.677)	(0.871)
Groundwater:					
Good quality	0.415	0.322	2.52*	-0.170	0.216
			(0.854)	(0.619)	(0.797)
Brackish	0.051	0.060	0.402	1.15	0.374
			(1.40)	(1.00)	(1.29)
Very brackish	0.030	0.080	-1.30	1.51	2.71
			(2.03)	(1.51)	(1.95)
Soil type:					
Sandy	0.189	0.210	-1.42*	-0.692	-0.026
			(0.656)	(0.467)	(0.601)
Maira	0.311	0.213	-0.666	-0.572	-0.052
			(0.736)	(0.531)	(0.684)
Chikni	0.218	0.265	-0.535	-1.21*	0.155
			(0.772)	(0.560)	(0.721)
p-value			0.0000	0.0571	0.0003

Notes: * denotes p-value< 0.05. Omitted categories: canal irrigation = none; groundwater = none; soil type = clay.

[a] From row (9), Table 4.

[b] From row (1), Table 5.

[c] From row (4), Table 5.