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Caste Discrimination and Transaction Costs in the Labor Market: Evidence from Rural North India

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Abstract

This paper is an empirical attempt to quantify caste-based discrimination in the labor market using household data taken from rural North India. In the regression analysis, transaction costs associated with entry into the labor market and reservation wages are estimated along with market wages. The estimation results provide evidence of the existence of transaction costs in the labor market and discrimination against backward classes with regard to access to regular employment. In line with previous studies, the results suggest that the achievements of India's reservation policy so far have at best been limited. In addition, a comparison between the estimates from the model employed in this paper and conventional (reduced-form) approaches shows that discrimination in labor market entry is likely to be underestimated in the conventional reduced-form approaches.

JEL classification codes: D23, J22, J24, J71
Keywords: regular employment, casual employment, labor market, India

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1. Introduction

For rural dwellers in developing countries, off-farm activities are becoming more important in determining their welfare. In rural India, where the labor market is relatively large, off-farm wage employment plays an important role as a source of income (Lanjouw and Shariff, 2004) and as insurance against agricultural risks (Kochar, 1999; Rose, 2001; Ito and Kurosaki, 2006).

While there are a number of studies on the labor market in India, only a few have investigated the role of labor market discrimination (see, e.g., Banerjee and Knight, 1985, on caste-based discrimination and Kingdon and Unni, 2001, on gender-based discrimination). The existence of discrimination may inhibit the expansion of the labor market which could help to reduce the dependence on agriculture. Moreover, the existence of discrimination may distort households' decisions not only with regard to labor allocation but also to human capital investment. In the rural context, both are issues of great importance. Examining labor market discrimination using rural household data therefore can make an important contribution to understanding how rural development can be achieved and rural poverty eliminated.

This paper is an empirical attempt to quantify caste-based discrimination in the labor market using household data from rural North India.[1] While several social reforms aiming to eradicate the caste system have been carried out,[2] the economic

[1] Castes (the traditional hereditary classes) in India consist of thousands of endogamous groups called *jatis* (the word literally means "birth"). Members of each *jati* are typically and traditionally engaged in the same occupation. Although the Indian caste system is originally based on Hinduism, not only Hindus but Muslims and members of other religious groups are also subject to it.

[2] One of the most important reforms is a series of policies called "reservation" in order to increase the economic opportunities available to members of the socially backward classes such as scheduled castes ("untouchables"), scheduled tribes and other backward castes. It is a type of affirmative action that reserves posts in educational and social institutions for these classes.

circumstances of socially backward and upper castes continue to differ substantially (Srinivasan and Kumar, 1999; Borooah, 2005). In the labor market, too, evidence of caste discrimination can be found and is reflected in market wages (Banerjee and Knight, 1985). However, an important question that has received little research attention so far is whether discrimination with regard to entry into the labor market exists, and if so, to what extent.

Several studies that have attempted to estimate discrimination in labor market access do exist. Banerjee and Knight (1985), for example, in their examination of wage differentials between scheduled and non-scheduled castes in the urban labor market based on the standard Blinder-Oaxaca decomposition (Blinder, 1973; Oaxaca, 1973), sought to distinguish two components: discrimination in wages within the same occupation ("wage discrimination") and discrimination in access to a well-paid job ("job discrimination"). They found that "unexplained" components account for a significant part of observed wage differentials, and that "wage discrimination" dominated "job discrimination." Using a similar method, Borooah et al. (2005) examined differences in employment rates between upper and backward castes. Their results also indicated that "job discrimination" against the backward classes explains only part of the observed differences.

However, their analyses on labor market participation are based on a reduced-form approach, so that the estimate of "job discrimination" virtually includes the effect of "wage discrimination." The aim of this paper is to address this shortcoming by employing a structural model in which the (fixed) transaction costs associated with entry into the labor market and reservation wages are estimated along with market

earnings from the observed individuals' labor supply responses.[3] If individuals knew all particulars about all available jobs, i.e., in the case of perfect information, there would be no need to devote time to looking for a suitable job. In reality, however, this is not the case. The costs of finding employment depend on an individual's characteristics and the features of the region he or she resides in. Moreover, if employers have discriminatory tastes toward certain social classes or utilize social class as supplementary information because of limited information on individual characteristics, the costs may differ from one class to another. The structural estimation employed in this study enables us to more clearly distinguish between discrimination in wages and in labor market entry. From a policy-making viewpoint, it is quite important to know whether labor market discrimination exists in the form of wage differentiation ("wage discrimination") or the range of jobs available ("job discrimination") or both.

The remainder of this paper is organized as follows. Section 2 proposes empirical models to estimate market/reservation wages and transaction costs simultaneously. Section 3 describes the dataset. The sample used in the analysis consists of working-age males in rural areas of Bihar and Uttar Pradesh in North India. Section 4 presents the empirical results. They provide evidence of the existence of transaction costs in the labor market and discrimination against backward classes with regard to access to regular employment, but no evidence of discrimination in wages from regular employment activities is found. Furthermore, a comparison between the

[3] For the studies on transaction costs associated with entry into the market, see, for instance, Cogan (1981) on married women's labor supply in the U.S., and Sadoulet et al. (1998) and Key et al. (2000) on corn producers' supply responses to labor and product markets in Mexico. These studies found that transaction costs matter for workers' or households' decisions on market involvement. In addition, Jacoby (1993) found evidence for the existence of transaction costs in the labor market in an analysis of farmers' labor supply in Peru, while Skoufias (1994) obtained a similar result with a sample of Indian rural households.

estimates from the structural model and the conventional (reduced-form) approaches shows that discrimination in labor market entry is likely to be underestimated when one employs the latter. Section 5 concludes the paper.

2. Empirical specification

Market labor can be broadly classified in terms of employment status into two types: casual and regular employment. In what follows, it is assumed that while there is no cost of entry into the *casual* labor market, transaction costs are involved in finding *regular* employment.[4] Transaction costs in the labor market represent fixed costs associated with labor market entry, such as actual expenditure and/or time spent traveling for finding employment. With transaction costs in monetary terms represented by C, the probabilities of being regularly employed and of being casually employed can respectively be expressed as follows:

$$\begin{aligned}\Pr(I_r = 1) &= \Pr(W_r - C > W_c,\ W_r - C > W^R) \\ &= \Pr(\ln W_r - \ln K > \ln W_c,\ \ln W_r - \ln K > \ln W^R),\end{aligned}$$

$$\begin{aligned}\Pr(I_c = 1) &= \Pr(W_c > W_r - C,\ W_c > W^R) \\ &= \Pr(\ln W_c > \ln W_r - \ln K,\ \ln W_c > \ln W^R),\end{aligned}$$

where I_k is an indicator variable that takes one if the individual is employed in a type k job ($k = r$: regular, c: casual), W_k is the wage earned in the job, W^R is the

[4] In general, specific skills or knowledge are not necessarily required in casual wage labor because this, for the most part, consists of physical labor. In addition, most casual laborers work within their villages, while regular workers often work in neighboring cities (outside the village). For these reasons, transaction costs are likely to be negligible and constant. Thus, the assumption not only simplifies the empirical model but is also plausible in the present context.

reservation wage,[5] and K is the relative value of the transaction cost defined by

$$K-1=\frac{C}{W_r-C}.$$

The functions of market wages in each type of employment, (relative) transaction costs, and reservation wages are respectively specified as

(1) $\ln W_k = X\beta_k + e_k, \quad e_k \sim N\left(0,\sigma_k^2\right), \quad k=r,c.,$

(2) $\ln K = Z_K\gamma_K + u_K, \quad u_K \sim N\left(0,\sigma_K^2\right),$

(3) $\ln W^R = Z_R\gamma_R + u_R, \quad u_R \sim N\left(0,\sigma_R^2\right),$

where X, Z_K, and Z_R are matrices of variables affecting market wages, transaction costs, and reservation wages; β and γ are vectors of coefficients to be estimated; and e and u are zero mean random error terms. Note that market wages can be observed only for market laborers; thus, we can only use the density $f\left(\ln W_k \mid I_k\right)$ when $I_k=1$. By applying Bayes' rule, $f\left(\ln W_k \mid I_k=1\right)=\Pr\left(I_k=1\mid \ln W_k\right)f\left(\ln W_k\right)/\Pr\left(I_k=1\right)$, and from standard conditional distribution results for joint normal random variables, we get the following equation:

$$\begin{aligned}\Pr\left(I_k=1\mid \ln W_k\right) &= \Pr\left(\eta_{1k}>\varepsilon_{1k},\eta_{2k}>\varepsilon_{2k}\mid \ln W_k\right)\\ &= \Phi_2\left(\frac{\eta_{1k}-\mu_{1k}}{\sigma_{1k}},\frac{\eta_{2k}-\mu_{2k}}{\sigma_{2k}},\rho_{12k}\right), \quad k=r,c.,\end{aligned}$$

where $\Phi_2(\cdot)$ is the bivariate standard normal cumulative distribution function, $\eta_{1r}=-\eta_{1c}=X\beta_r-Z_K\gamma_K-X\beta_c$, $\eta_{2r}=X\beta_r-Z_K\gamma_K-Z_R\gamma_R$, $\eta_{2c}=X\beta_c-Z_R\gamma_R$, $\varepsilon_{1r}=-\varepsilon_{1c}=-e_r+u_K+e_c$, $\varepsilon_{2r}=-e_r+u_K+u_R$, $\varepsilon_{2c}=e_c-u_R$. In addition, the means and elements in the variance-covariance matrix of the conditional distribution of ε

[5] As mentioned in the following section, the majority of rural workers in developing countries are engaged in self-employed farming and it is not necessarily assumed in our model that leisure is the sole alternative to wage labor. Thus, the term "reservation wage" is used in the broad sense that it determines whether the worker is engaged in wage work or not (including self-employment).

given e_k are, respectively,

$$\mu_{\ell,k} = E\left(\varepsilon_{\ell,k} \mid e_k\right) = \frac{Cov\left(\varepsilon_{\ell,k}, e_k\right) \cdot e_k}{\sigma_k},$$

$$\sigma^2_{\ell,r} = V\left(\varepsilon_{\ell,k} \mid e_k\right) = E\left(\varepsilon^2_{\ell,k}\right) - \left(\frac{Cov\left(\varepsilon_{\ell,k}, e_k\right)}{\sigma_k}\right)^2, \qquad k = r, c., \quad \ell = 1,2.$$

$$\rho_{12k} = \frac{Cov\left(\varepsilon_{1k}, \varepsilon_{2k} \mid e_k\right)}{\sigma_{1k}\sigma_{2k}} = \frac{Cov\left(\varepsilon_{1k}, \varepsilon_{2k}\right) - Cov\left(\varepsilon_{1k}, e_k\right) \cdot Cov\left(\varepsilon_{2k}, e_k\right) / \sigma_k^2}{\sigma_{1k}\sigma_{2k}},$$

where $E(\cdot)$, $Cov(\cdot)$, and $V(\cdot)$ represent operators taking the mathematical expectation, the covariance, and the variance of their arguments, respectively. Further, the probability of being a non-participant ($I_n = 1$) is

$$\begin{aligned}\Pr\left(I_n = 1\right) &= \Pr\left(\ln W_r - \ln K < \ln W^R, \ln W_c < \ln W^R\right) = \Pr\left(\eta_{1n} < \varepsilon_{1n}, \eta_{2n} < \varepsilon_{2n}\right) \\ &= \Phi_2\left(-\frac{\eta_{1n}}{\sigma_{1n}}, -\frac{\eta_{2n}}{\sigma_{2n}}, \rho_{12n}\right),\end{aligned}$$

where $\eta_{1n} = \eta_{1r}$, $\eta_{2n} = \eta_{2c}$, $\varepsilon_{1n} = \varepsilon_{1r}$, $\varepsilon_{2n} = \varepsilon_{2c}$, $\sigma^2_{\ell,n} = V\left(\varepsilon_{\ell,n}\right)$ ($\ell = 1,2$), and $\rho_{12n} = Cov\left(\varepsilon_{1n}, \varepsilon_{2n}\right) / \left(\sigma_{1n}\sigma_{2n}\right)$. Therefore, combining all of these equations we obtain the following log likelihood function:

$$\begin{aligned}LL = \sum \{ & I_r \cdot \ln\left[\Pr\left(I_r = 1\right) \cdot f\left(\ln W_r \mid I_r = 1\right)\right] \\ & + I_c \cdot \ln\left[\Pr\left(I_c = 1\right) \cdot f\left(\ln W_c \mid I_c = 1\right)\right] + I_n \cdot \ln\left[\Pr\left(I_n = 1\right)\right] \}\end{aligned}$$

$$\begin{aligned}(4) \qquad = \sum \Bigg\{ & I_r \left[\ln \Phi_2\left(\frac{\eta_{1r} - \mu_{1r}}{\sigma_{1r}}, \frac{\eta_{2r} - \mu_{2r}}{\sigma_{2r}}, \rho_{12r}\right) + \ln \phi\left(\frac{\ln W_r - X\beta_r}{\sigma_r}\right) - \ln \sigma_r \right] \\ & + I_c \left[\ln \Phi_2\left(\frac{\eta_{1c} - \mu_{1c}}{\sigma_{1c}}, \frac{\eta_{2c} - \mu_{2c}}{\sigma_{2c}}, \rho_{12c}\right) + \ln \phi\left(\frac{\ln W_c - X\beta_c}{\sigma_c}\right) - \ln \sigma_c \right] \\ & + I_n \cdot \ln \Phi_2\left(-\frac{\eta_{1n}}{\sigma_{1n}}, -\frac{\eta_{2n}}{\sigma_{2n}}, \rho_{12n}\right) \Bigg\},\end{aligned}$$

where $\phi(\cdot)$ is the (univariate) standard normal density function. Note that $\ln K$ and

$\ln W^R$ are estimated as latent variables in our model and hence one may think identification problems arise when estimating parameters γ_K (γ_R) separately from γ_R (γ_K) and β_k ($k = r, c.$). However, due to the assumption that there are no transaction costs involved in finding casual employment, but such costs are present in finding regular employment, identification of γ_K and γ_R are guaranteed in turn by β_k, which can be estimated consistently from the density $f(\ln W_k)$ in the log likelihood function. Therefore, parameters (β and γ) can be consistently estimated as long as the explanatory variables are orthogonal to the error terms.[6]

3. Data

3.1 Sample and key features

The data employed in this paper are from the *Survey of Living Conditions, Uttar Pradesh and Bihar*, which is one of the Living Standard Measurement Study (LSMS) surveys conducted in developing countries. Uttar Pradesh (UP) and Bihar are located in the Ganges Plain of North India and are known for their high poverty incidence. The survey was conducted in 1997/98, covering 1,035 households, 57 villages, and 13 districts in Bihar and 1,215 households, 63 villages, and 12 districts in UP. The sample used in the analysis comprises male household members aged between 15 and 60 and consist of 3,324 individuals.[7]

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[6] Furthermore, even in the case that all elements in X, Z_K and Z_R are identical, that is, there is no exclusion restriction on parameters, identification of β and γ remains guaranteed.

[7] To focus on labor market participation, students (351), disabled people (35) and members with missing information on work activities and/or other characteristics (227) were excluded from the sample.

[Table 1]

Information on work activities and wages is available for each household member from January 1997 to December 1997. Market wages are reported by employment status: regular (salaried) or casual employment. Table 1 shows summary statistics for labor market participation and monthly earnings of working-age males by caste. Monthly earnings from regular employment activities in addition to the base salary include other payments such as bonuses, while monthly earnings from casual employment activities are the average earnings in working months.

Obviously, the composition of the employment status and average monthly earnings differ from caste to caste. Individuals belonging to middle and backward castes are less likely to be in regular employment and more likely to be casually employed than those belonging to the upper or Muslim upper castes. Especially, among the scheduled castes, more than half of all male workers are casually employed. In addition, the average earnings of upper caste members are more than thirty percent higher than those of members of the other caste classes in both wage activities.

On the other hand, there are also large differences in the average schooling years between castes. As can be seen in Table 1, there is a negative correlation between the average years of education and the proportion of casual workers. Given that most casual workers are engaged in agricultural wage work, this may reflect the lack of response of agricultural wages to human capital (Kurosaki and Khan, 2006) and the stigma associated with working as an agricultural laborer in rural India. In the regression analysis, it is examined whether the differences in employment status and wage differentials between castes are mainly attributable to human capital

characteristics or other factors, namely caste-based discrimination.

3.2 Empirical variables

The empirical variables used in the regression analysis are summarized in Tables 2 and 3. The first four rows provide information on market earnings (*Log earnings-regular*, *Log earnings-casual*) and employment status (*Regular emp. dummy* and *Casual emp. dummy*). Out of the 2,000 individuals in the sample who did not participate in the labor market, 1,105 (55 percent) were engaged in agricultural self-employed work, 694 (35 percent) in non-agricultural self-employed work, and 201 (10 percent) in domestic work or were not working. Thus, strictly speaking, the "reservation wage" to be estimated using this dataset mainly reflects the marginal productivity in self-employment activities.[8]

[Table 2]

[Table 3]

Human capital characteristics affecting market earnings, transaction costs and reservation wages include educational attainment, job experience, and age. Actual job experience is not available from the dataset and is therefore measured using the following formula: age minus years of education minus 6 years (*experience*).[9] Age, of course, reflects not only job experience but also human capital accumulation over an individual's lifetime; however, to concentrate on the effect of human capital

[8] In this connection, because information on incomes from self-employment activities is not available, the production functions for those activities cannot be estimated directly.
[9] There are 11 cases (less than 0.4 percent of the sample) in which *experience* becomes a negative value (the minimum is -3), and these are replaced by zero.

accumulation after formal schooling, the variable *experience* is employed in the earnings equation. The effects of education and experience on market earnings are expected to be positive but are likely to depend on the employment status. The impact of education on the cost of entry is expected to be negative, while that of age is expected to be positive. The effect on the reservation wage is expected to be positive both in the case of education and age.

To control for the different roles of household members, other individual-level characteristics are taken into account by employing dummy variables for household heads and firstborn sons (*HH head* and *Firstborn son*) as well as the number of elder brothers (*No. of elder brothers*). Because these variables are likely to have no effects on market earnings and transaction costs, they are included only in the reservation wage function. Being the household head or the firstborn son of the household head may magnify the economic responsibilities of the person in the household, while having many elder brothers is likely to have the opposite effect. Thus, the expected signs are positive for *HH head* and *Firstborn son*, but negative for *No. elder brothers*.

Household level characteristics are family structure, farming assets, and caste membership. Variables for household structure are the number of working age and non-working age members (*No. of working age members* and *No. of non-working age members*). The number of working age (non-working age) members mainly captures the working (dependent) population in the household and the expected impact on reservation wages is negative (positive). On the other hand, the impact of these variables on market earnings and transaction costs is somewhat ambiguous. If household members' nutritional status is negatively correlated with household size, the effects of these variables on market earnings are expected to be negative and those on

transaction costs are positive.

As farming assets, the size of farmland owned by households (*Land size*), the share of farmland that is irrigated (*Irrigation ratio*), and the value of semi-fixed capital in agricultural production (*Agr. capital*) and livestock (*Livestock*) are employed.[10] These variables mainly capture household members' productivity in own-farm activities and the impact of these variables on reservation wages are therefore expected to be positive. Turning to the impact on market earnings and transaction costs, since household members' nutritional status is likely to be correlated with farming assets, they are likely to have a positive effect on market earnings and a negative one on transaction costs.

Caste dummies (with the upper caste as the reference group) capture "unexplained" differences between castes in market earnings, transaction costs and reservation wages. Although the inclusion of caste dummies in the reservation wage function may appear dubious, it seems in fact only natural to assume that workers belonging to a caste that has traditionally suffered from discrimination in the labor market have a negative attitude or low expectations with regard to finding a job. This being the case, the reservation wage of those belonging to an economically disadvantaged class is low in comparison with those belonging to an advantaged class. Table 1 suggests that the impact of belonging to one of the backward classes is negative on market earnings and positive on transaction costs. Therefore, a negative effect is expected on reservation wages.

As regional characteristics, the ratio of the landless within the village (*Ratio*

[10] Given that the "reservation wage" mainly captures the marginal productivity in agricultural/non-agricultural self-employed activities, variables affecting the productivity in non-farm enterprise production should be added. Unfortunately, data that would allow us to control for productivity in these activities are not available.

of landless), the average distance to the nearest bank, police station and secondary school (*Distance to facilities*), and a *UP state dummy* are employed. The *Ratio of landless* is a proxy for the number of potential workers, and its impact on market earnings is expected to be negative, although that on transaction costs and reservation wages is ambiguous. The *Distance to facilities* captures proximity to the nearest city and hence the degree of economic development of the village. It is expected that in villages far from an urban area, market earnings and reservation wages are low and the costs of finding regular employment are high. In addition to these variables, the ratio of workers in regular employment, a proxy for information exchange about jobs among villagers, is used only in the transaction cost function and is likely to lower transaction costs.

4. Estimation Results

Before embarking on the empirical investigation, some limitations of the analysis should be mentioned. First, several household level variables to control for family structure and farming assets are excluded from the market earnings and transaction cost equations because of a convergence problem. Therefore, there is a possibility that the estimates suffer from omitted variables bias. For instance, in casual employment activities, workers' nutritional status, which may be captured by these variables, is an important determinant of market wages in the Indian labor market (Deolalikar, 1988; Weinberger, 2003). In addition, it is possible that the estimated effect of human capital on market wages are biased (Heckman and Hotz, 1986; Kingdon, 1998). Unfortunately, the model that includes these variables in all functions failed to achieve a convergence of the likelihood function, but the effects of these variables on

market wages and transaction costs are likely to be negligible.

A second limitation is that there is also a possibility of bias due to other omitted variables. If individuals' schooling choices are determined by their ability (or the possibility of getting a job), this may cause schooling effects to be biased (i.e., there would be "ability bias") and the exclusion of students from the sample would make the problem worse. To examine this possibility, estimations using limited samples – including only those aged between 18 and 60, and including only those aged between 20 and 60 – were conducted. The coefficients were slightly different from the estimates using the full sample, but there were no systematic differences.[11] This suggests that there is no "ability bias" or, if there is, it is rather small.

Table 4 shows the estimation result of the log-likelihood function (4). As can be seen, the coefficient on most of the explanatory variables takes the expected sign. The null hypothesis of no transaction costs is tested by using a likelihood ratio test. The χ^2 statistic of 189.75 (p-value is 0.00) indicates rejection of the hypothesis.

4.1 Analysis of the market wage functions

Estimation results for the market wage functions are presented in the first and second columns of Table 4. The effect of human capital on the market wage differs sharply depending on the employment status. While the return to education is positive in the case of regular employment, it shows an inverted U-shape in the case of casual employment, indicating that marginal returns to education become negative at more than eight years of education. The impact of job experience peaks at 44 years for regular employment and at 20 years for casual employment. The fact that human capital does not contribute much toward improving productivity in casual employment activities

[11] These results are available on request.

may reflect the fact that the majority of casual workers are hired for unskilled, manual labor, such as agricultural work.

Looking at caste membership, the coefficients on caste dummies are negative in all cases in the casual labor wage functions and negative in most cases in the regular labor wage function. In casual employment activities, membership of an agriculture-based backward or scheduled caste significantly decreases wages.[12] This result is consistent with the findings from Table 1. The monthly earnings from casual employment activities for the agriculture-based backward castes are 33.5 percent ($\approx e^{0.289}-1$, 68 percent of the total wage differentials) lower than those for the upper castes, and those for the scheduled castes are 37.4 percent (67 percent) lower.

However, one should guard against any hasty interpretation of this result as evidence of wage discrimination in the casual labor market. In our analysis, workers' occupations are simply classified into two types (casual and regular employment), hence the diversity of occupations within the same status is ignored. An alternative interpretation, therefore, is that the socially backward classes are traditionally occupied in low-paying casual employment activities, so that the results possibly reflect occupational segregation rather than wage discrimination. Unfortunately, because of data limitations, it is impossible to investigate this issue further by disaggregating the sample by occupation or industry.

[Table 4]

[12] The result also shows that membership of a middle caste has a significant negative effect on wages from regular employment activities. But this result should be treated with caution because the sample contains only eight middle caste workers engaged in this type of job.

4.2 Analysis of the transaction cost and reservation wage functions

Estimation results of the transaction cost and reservation wage functions are presented in the third and forth columns of Table 4. Schooling years have a negative effect on transaction costs and a positive effect on reservation wages when evaluated at the sample mean (4.84 years), but the effect of *Schooling years* on transaction costs and the effect of *Schooling years squared* in both estimates is insignificant. Age has a significant positive effect on transaction costs but no significant effect on reservation wages.

Turning to caste dummies, individuals belonging to groups characterized as backward face significantly higher transaction costs than those belonging to the upper castes.[13] The results indicate that membership of one of the four backward classes increases transaction costs by 39 percent (*backward-agr.*) to 79 percent (*scheduled*), which may suggest discrimination in labor market entry.

However, the results obtained here should be interpreted with care. For instance, it is possible that caste membership captures the effects of caste networks (Munshi and Rosenzweig, 2005). If members of the upper castes are traditionally employed in regular employment activities and hiring through referrals is prevalent, then the negative effect of backward-caste membership may simply reflect the lack of such network referrals. There is another possible explanation as well. As already mentioned, workers belonging to groups that have traditionally experienced discrimination may hold low expectations of gaining a job. In this case, even in the

[13] To investigate the robustness of the impact of caste membership on transaction costs, an alternative specification is tested in which caste dummies are excluded from the market earnings and reservation wage equations. This results shows that the coefficients of dummies for the lower castes (except for the *Muslim backward dummy*) are positive and significant in the transaction cost equation. The results are available from the author on request.

absence of discrimination by employers based on their (or other employees') tastes or beliefs, these workers may not look as hard for a job as members of other castes. The higher transaction costs for members of socially backward castes may result from such self-fulfilling beliefs. However, even in these circumstances, there is no doubt that government policies to combat inequality in employment opportunities have not been successful in the study region. Our results show that members of socially backward castes indeed face greater difficulties in finding regular employment than members of the upper castes.

In the reservation wage function, on the other hand, all dummies have negative coefficients. This is consistent with the story mentioned in the previous section. However, this result may simply reflect the low marginal productivities in the self-employment activities that members of the socially backward castes are engaged in because they face discrimination even in these activities.

The effects of the other control variables are as expected. Among regional characteristics, the ratio of salaried workers in the same village significantly decreases the cost of entry. This indicates the importance of social networks in finding regular employment.

4.4 Effects of human capital and caste membership on participation

Thus far, the discussion has concentrated on the impact of the explanatory on the dependent variables. This subsection focuses on the impact of the explanatory variables on the probability of gaining regular employment. The marginal effects of human capital and caste membership on regular labor market participation, holding all other variables constant at their sample mean, are shown in the first column of Table 5. Furthermore, the effects are decomposed into three parts: those through market

earnings (in the second column), those through transaction costs (in the third column), and those through reservation wages (in the forth column).

[Table 5]

The table indicates that an additional year of education increases the probability of being in regular employment by approximately 1.0 percentage points when evaluated at the sample mean of schooling years (4.84 years). The breakdown of the effect of schooling on the probability of being in regular employment into its components shows that it is largely through the effect on market earnings.

The effect of caste membership through transaction costs lowers the likelihood of members of backward castes to be in regular employment by 6.5 percentage points (68 percent of the total difference in the employment rate) in the case of other backward castes, by 3.3 percentage points (32 percent) in the case of scheduled castes, and by 3.0 percentage points (39 percent) in the case of agriculture-based backward castes. However, these negative effects through transaction costs are to a great extent attenuated through earnings and reservation wages. This result implies that an analysis of the role of discrimination in labor market entry based on a reduced form approach is likely to underestimate the impact of discrimination. In fact, this is confirmed by the probit estimation (in the fifth column) and the estimation using the Blinder-Oaxaca decomposition procedure (in the sixth column). The comparison of the predicted mean probability and of the percent correctly predicted between our model and the probit model shows that our model predicts the labor participation decision equally as well as the probit model. Thus, when one employs a reduced form model, the

possibility of this underestimation should be recognized.

5. Conclusion

This paper examined caste discrimination in the labor market by estimating simultaneously market earnings, reservation wages, and the costs of finding regular employment. The estimation results suggest that socially backward castes do face disadvantages in finding regular employment in the sense that they face higher transaction costs associated with entry into the labor market. On the other hand, there is no evidence for wage discrimination in regular employment activities. Thus, these results suggested that caste-based discrimination takes the form of "job discrimination," which limits the range of available jobs, rather than "wage discrimination,"[14] However, this is at odds with Banerjee and Knight's (1985) findings, which indicated that "wage discrimination" explained a large part of the "unexplained" component of wage differentials. Although one should be careful about directly comparing their results with ours because the data they used was of 1,115 migrant laborers in Delhi in 1975-76, it should at least be noted that it is possibility that their analysis underestimated the role of "job discrimination." This paper has shown that the reservation wages of the socially backward castes are lower than those of the upper castes and consequently estimates of "job discrimination" in a reduced-form approach may capture not only discrimination in entry but also such self-selection effects.

Furthermore, it was found that the large differences in educational attainment between castes represent one important cause of inequality in employment status. Thus,

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[14] Similarly, studies on racial discrimination in the U. S. also found that discrimination typically takes the form of market segregation rather than price differentiation (see, e.g., Arrow, 1998).

as India's reservation policy has aimed to do, promoting opportunities in education and employment for socially backward castes can be an instrument in eliminating inequality in economic conditions between castes. At the same time, however, the results obtained here suggest that the reservation policy so far has had little effect on rural dwellers in the study region, Uttar Pradesh and Bihar, although it is not clear whether this is the results of some fundamental problems with this policy or reflects the specific circumstances of these two states, which are among the least developed in India. While it is beyond the scope of this paper to provide a detailed assessment of the policy, the fact that backward castes continue to be disadvantaged half a century after it was introduced suggests that its achievements have at best been limited.

Finally, several limitations of this paper should be mentioned. One possible problem is associated with the validity of the assumption introduced in the estimation model that transaction costs in the casual labor market are constant. As mentioned earlier, this assumption seems plausible in the context of rural India, and the comparison of results between our model and the probit model suggest that our model is valid. However, it is impossible to test the sensitivity of the model to the assumption since it is an essential assumption for parameter identification. But the most important limitation of this paper probably is that there is a possibility that estimated caste differences in market wages and participation probabilities capture other omitted differences between castes rather than discrimination. Although this appears to be a common problem in studies on labor market discrimination, the occupational classification in this paper (i.e., casual or regular employment) may exacerbate the problem. Essentially, wage rates and individual responses may vary not simply by employment status but by sector or type of occupation, but such occupational diversity

within the same employment status was ignored in our analysis. Consequently, estimated caste differences may to some extent reflect the effects of occupational segregation. Given that India's caste system is based on hereditary occupations, future analyses should try to employ data disaggregated by occupation or industry.

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Table 1: Labor supply of adult males by caste

Caste		Market labor		Average monthly earnings			Schooling years
	NOB	All	Regular	Casual	All	Regular	Casual	
ALL	3324	0.40	0.12	0.28	1293.0	2564.1	768.1	4.84
Upper	506	0.25	0.19	0.07	2804.4	3296.3	1373.4	9.25
Middle	82	0.20	0.10	0.10	1319.4	1745.8	893.1	6.73
Backward-agr.	963	0.33	0.11	0.21	1403.3	2398.3	869.4	5.12
Backward-other	645	0.38	0.09	0.28	1233.0	2359.8	863.6	4.09
Scheduled	818	0.59	0.09	0.51	847.4	2267.8	600.3	2.79
Muslim upper	109	0.43	0.18	0.25	1513.2	2217.3	991.7	4.44
Muslim backward	201	0.44	0.10	0.33	1187.3	2326.7	830.2	2.52

Note: (1) Employment status is based on workers' primary job classified based on their annual working days. Note that approximately 20 percent of workers were also engaged in secondary activities.
(2) The caste category in this table follows the classification of the *Survey of Living Conditions, Uttar Pradesh and Bihar*. Note that this does not necessarily correspond to the classification of the "reservation policy." For instance, "backward-other" in the table does not exactly correspond to "other backward castes (OBC)" in the reservation list.

Table 2: Definition of variables

Name	Definition	Expected Sign		
	Equation	(1) $\ln W$	(2) $\ln K$	(3) $\ln W^R$
A. Individual level variable				
Log earnings-regular	Log of monthly earnings paid to a regular worker			
Log earnings-casual	Log of average monthly earnings paid to a casual laborer			
Regular emp. dummy	Dummy variable for those mainly in regular employment			
Casual emp. dummy	Dummy variable for those mainly in casual employment			
Schooling years	Standardized years of completed education corresponding to the standard education system in India	+	−	+
Experience	Age minus standardized years of completed education minus 6 years	+	No	No
Age	Age of the person	No	+	+
HH head dummy	Dummy variable for the household head	No	No	−
Firstborn son dummy	Dummy variable for the firstborn son of the household head	No	No	−
No. of elder brothers	Number of elder brothers (incl. brothers-in-law)	No	No	+
B. Household level variables				
No. of working age members	Number of household members aged between 15 and 60	(−)	(+)	+
No. of non-working age members	Number of household members other than working-age members	(−)	(+)	−
Land size (10acres)	Land owned by the household	(+)	(−)	+
Irrigation ratio	Ratio of irrigated land to total land size owned by the household	(+)	(−)	+
Agr. capital (Rupees)	Value of fixed agricultural capital owned by the household	(+)	(−)	+
Livestock (Rupees)	Value of livestock owned by the household	(+)	(−)	+
Middle	Dummy variable for middle castes	−	+	−
Backward-agr.	Dummy variable for agriculture-based backward castes	−	+	−
Backward-other	Dummy variable for other backward castes	−	+	−
Scheduled	Dummy variable for scheduled castes	−	+	−
Muslim upper	Dummy variable for upper Muslim castes	−	+−	−
Muslim backward	Dummy variable for backward Muslim castes	−	+	−
C. Other				
Ratio of landless	Ratio of the landless in the village	−	+−	+−
Distance to facilities (Km)	Average distance to the nearest bank, police station, and secondary school from the village	−	+	−
Ratio of salaried workers (%)	Ratio of village-total workers in regular employment (other than household members) to village-total working-age people	No	−	No
UP state dummy	Dummy variable for Uttar Pradesh	+−	+−	+−

Table 3: Summary statistics of variables

Name	NOB	Mean	Std. Dev.	Min.	Max.
A. Individual level variables					
Log earnings-regular	387	7.60	0.74	5.30	9.62
(in Rupees)	387	2564.1	1863.3	200	15000
Log earnings-casual	937	6.36	0.73	3.65	8.72
(in Rupees)	937	768.1	678.2	38.3	6111.1
Regular emp. dummy	3324	0.12			
Casual emp. dummy	3324	0.28			
Schooling years	3324	4.84	5.20	0	20
Experience	3324	23.49	14.20	0	54
Age	3324	34.32	12.57	15	60
HH head dummy	3324	0.51			
Firstborn son dummy	3324	0.27			
No. of elder brothers	3324	0.39	1.03	0	10
B. Household level variables					
No. of working age members	3324	4.33	2.23	1	13
No. of non-working age members	3324	3.22	2.43	0	17
Land size (10acres)	3324	2.55	0.52	0	9.3
Irrigation ratio	3324	0.65	0.43	0	1
Agr. capital (Rupees)	3324	7134.48	30197.36	0	373600
Livestock (Rupees)	3324	7443.31	10692.68	0	150000
Middle	3324	0.02			
Backward-agr.	3324	0.29			
Backward-other	3324	0.19			
Scheduled	3324	0.25			
Muslim upper	3324	0.03			
Muslim backward	3324	0.06			
C. Other					
Ratio of landless	3324	0.39	0.21	0	0.99
Distance to facilities (Km)	3324	5.77	3.58	0.5	20
Ratio of salaried workers (%)	3324	5.45	4.27	0	26.23
UP state dummy	3324	0.55			

Table 4: Estimation results for market earnings, transaction costs and reservation wages

	Monthly earnings (Regular)	Monthly earnings (Casual)	Transaction costs	Reservation wages
Individual characteristics				
Schooling years	0.062(3.38)‡	0.060(2.63)‡	-0.044(1.45)	0.059(2.80)‡
Schooling years squared/100	0.114(1.04)	-0.383(1.76)*	0.349(1.48)	-0.301(1.54)
Experience	0.072(5.52)‡	0.014(0.85)†	-	-
Experience squared/100	-0.081(3.10)‡	-0.035(1.10)‡	-	-
Age	-	-	0.053(5.15)‡	0.003(0.21)
Age squared/100	-	-	-0.031(2.44)‡	-0.005(0.31)
HH head dummy	-	-	-	-0.020(0.53)
Firstborn son dummy	-	-	-	-0.051(1.16)
No. of elder brothers	-	-	-	0.000(0.02)
Household characteristics				
No. of working age members	-	-	-	0.008(0.70)
No. of non-working age members	-	-	-	-0.004(0.81)
Land size	-	-	-	0.054(0.79)
Irrigation ratio	-	-	-	0.147(1.66)*
Agr. Capital	-	-	-	0.032(2.13)†
Livestock	-	-	-	0.045(1.17)
Caste dummies				
Middle	-0.302(2.04)†	-0.812(1.44)	0.611(1.05)	-0.705(1.29)
Backward-agr.	-0.026(0.25)	-0.289(1.77)*	0.326(1.74)*	-0.403(2.54)†
Backward-other	-0.026(0.22)	-0.263(1.49)	0.357(1.69)*	-0.363(2.17)†
Scheduled	0.000(0.00)	-0.469(2.17)†	0.583(2.47)†	-0.759(3.82)‡
Muslim upper	-0.124(0.73)	-0.124(0.48)	0.007(0.02)	-0.272(1.19)
Muslim backward	0.143(0.83)	-0.276(1.38)	0.440(1.69)*	-0.388(2.08)†
Regional characteristics				
Ratio of landless	-0.374(2.00)†	-0.119(0.98)	-0.175(0.78)	-0.214(1.63)
Distance to facilities/10	-0.066(0.53)	-0.007(0.10)	0.007(0.05)	0.004(0.05)
Ratio of salaried workers	-	-	-0.020(2.08)†	-
UP state dummy	0.044(0.61)	0.115(2.05)†	-0.046(0.51)	0.045(0.67)
Intercept	5.588(19.4)‡	6.350(14.2)‡	-0.543(1.30)	6.562(14.4)‡
Standard error	0.687(14.5)‡	0.720(12.0)‡	0.169(1.58)	0.678(13.0)‡

Note: (1) Numbers in parentheses are z-values based on clustering robust standard errors using households as clusters.
(2) * Significant at 10%; † significant at 5%; ‡ significant at 1%.
(3) NOB = 3324; log-likelihood = -3887.19. H_0 : zero slope, LR $\chi^2(62)$ = 1223.70; H_0 : no transaction costs, LR $\chi^2(17)$ = 189.75.

Table 5: Marginal effects on participation in the regular labor market

Through:	Total (a)+(b)+(c)	Market earnings (a)	Transaction costs (b)	Reservation wages (c)	**Probit estimates** (Table A-1)	**Blinder-Oaxaca decomposition** (Table A-2)
Human capital						-
Schooling years	0.010	0.009	0.004	-0.004	0.010	
Age/10	0.019	0.152	-0.135	0.001	0.026	
Caste dummies						
Middle	-0.049	0.013	-0.074	0.012	-0.042	-0.054
Backward-agr.	-0.013	0.010	-0.030	0.007	-0.006	-0.019
Backward-other	-0.038	0.017	-0.065	0.011	-0.028	-0.032
Scheduled	-0.021	0.007	-0.033	0.005	-0.014	-0.022
Muslim upper	0.015	-0.009	-0.003	0.026	0.043	0.054
Muslim backward	0.004	-0.003	0.008	-0.001	0.024	0.014

Note: (1) Calculated figures in the table are the average marginal effect of each variable on the probability of being in regular employment. Note that in the case of binary variables, calculated figures are the average change in response probabilities when the variable changes from 0 to 1, and in general the whole effect does not equal the sum of (a), (b) and (c) because of the calculation method. For ease of comparison, however, it is adjusted in such a way that the sum equals the whole effect.
(2) The marginal effects through "Market earnings" are the sum of marginal effects through monthly earnings from regular and from casual employment activities.
(3) The dependent variable for the probit estimation presented in the fifth column is an indicator variable that takes one if a worker is engaged in a regular employment activity and zero otherwise (see Table A-1). While the predicted mean probability with the probit model (Table A-1) is 0.1164, that with our procedure (Table 3) is 0.1162 (the sample mean is 0.1164). In addition, the percentage of observations being correctly predicted with regard to the employment status (using the 50 percent rule) is 0.885 for the probit model and 0.884 for our model. Thus, our model appears to predict the labor participation decision equally as well as the probit model.
(4) In the calculation of the effects of schooling years and age, those through job experience are taken into account using the following relation between them: Experience = age – schooling years – 6.

Appendix 1: Probit estimation of the probability of being in regular employment

Table A-1: Estimation result

Dependent variable: Regular emp. dummy	Coef.	z-value
Individual characteristics		
Schooling years	0.050	(2.97)‡
Schooling years squared/100	0.056	(0.50)
Age	0.069	(3.82)‡
Age squared/100	-0.079	(3.43)‡
HH head dummy	-0.103	(0.80)
Firstborn son dummy	0.042	(0.37)
No. of elder brothers	-0.020	(0.44)
Household characteristics		
No. of working age members	0.032	(1.72)*
No. of non-working age members/10	-0.015	(1.03)
Land size	-0.089	(0.93)
Irrigation ratio	0.231	(2.80)‡
Agr. capital	-0.030	(2.02)†
Livestock	-0.117	(2.92)‡
Caste dummies		
Middle	-0.281	(1.34)
Backward-agr.	-0.036	(0.38)
Backward-other	-0.169	(1.57)
Scheduled	-0.081	(0.75)
Muslim upper	0.221	(1.31)
Muslim backward	0.131	(0.87)
Regional characteristics		
Ratio of landless	-0.101	(0.67)
Distance to facilities/10	-0.025	(0.25)
Ratio of salaried workers	0.049	(7.24)‡
UP state dummy	-0.005	(0.07)
Intercept	-3.127	(8.59)‡

Note: (1) See Table 4, Notes (1) and (2).

(2) NOB=3324; log-likelihood = -1064.56, pseudo R^2 = 0.110. H_0: zero slope, LR $\chi^2(23)$ = 262.45.

Appendix 2: Estimation of "job discrimination" à la Banerjee and Knight (1985)

The estimation procedure explained below is an application of the Blinder-Oaxaca decomposition method for the trichotomous occupational choice model (multinomial logit model). The probability that an individual i belonging to the g th social group ($g = R$: reference group, C: comparison group) will be a regular employee is given by

$$\text{(A-1)} \qquad \mathrm{P}_{i,r}^{g} = \frac{\exp X_i^{g} \beta_r^{g}}{\exp X_i^{g} \beta_r^{g} + \exp X_i^{g} \beta_c^{g} + \exp X_i^{g} \beta_n^{g}} \ ,$$

where X is a vector of variables affecting the labor participation decision and β_k is a vector of coefficients corresponding to the k th type of employment activity (r: regular employment, c: casual employment, n: not employed). Let $\tilde{\mathrm{P}}_{i,r}^{C}$ be the probability that an individual belonging to a comparison group (C) will be a regular employee if he belongs to a reference group (R) and take the average for each probability, $\mathrm{P}_r^{g} = \sum_{i=1}^{N^g} \mathrm{P}_{i,r}^{g} \Big/ N^{g}$, ($g = R, C.$) and $\tilde{\mathrm{P}}_r^{C} = \sum_{i=1}^{N^C} \tilde{\mathrm{P}}_{i,r}^{C} \Big/ N^{C}$. The difference between the two groups in the probability of being regularly employed is decomposed as follows:

$$\text{(A-2)} \qquad \mathrm{P}_r^{C} - \mathrm{P}_r^{R} = \left(\mathrm{P}_r^{C} - \tilde{\mathrm{P}}_r^{C}\right) + \left(\tilde{\mathrm{P}}_r^{C} - \mathrm{P}_r^{R}\right) \ .$$

The first component in this equation measures the "unexplained" difference due to differences in coefficients, while the second measures the "explained" difference due to differences in characteristics. Table A-2 shows the estimates for each component in equation (A-2).

Table A-2: Blinder-Oaxaca decomposition for participation rates in regular employment

: **Comparison group:**	Middle	Backward-agr	Backward-Other	Scheduled	Muslim Upper	Muslim Backward
Difference in means, $P_r^C - P_r^R$	-0.092	-0.075	-0.097	-0.102	-0.006	-0.085
Difference due to coefficients, $\left(P_r^C - \widetilde{P}_r^C\right)$	-0.054	-0.019	-0.032	-0.022	0.054	0.014
Difference due to characteristics, $\left(\widetilde{P}_r^C - P_r^R\right)$	-0.038	-0.057	-0.064	-0.080	-0.060	-0.099
Sample size	82	963	645	818	109	201

Note: (1) The reference group is the upper caste with a sample size of 506 individuals.