

Networks and Misallocation: Insurance, Migration, and the Rural-Urban Wage Gap[†]

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We provide an explanation for the large spatial wage disparities and low male migration in India based on the trade-off between consumption smoothing, provided by caste-based rural insurance networks, and the income gains from migration. Our theory generates two key empirically verified predictions: (i) males in relatively wealthy households within a caste who benefit less from the redistributive (surplus-maximizing) network will be more likely to migrate, and (ii) males in households facing greater rural income risk (who benefit more from the insurance network) migrate less. Structural estimates show that small improvements in formal insurance decrease the spatial misallocation of labor by substantially increasing migration. (JEL G22, J31, J61, O15, O18, R23, Z13)

The misallocation of resources is widely believed to explain a substantial proportion of the variation in productivity and income across countries. Past work has documented both differences in productivity across firms (e.g., Restuccia and Rogerson 2008; Hsieh and Klenow 2009) and the misallocation of resources across sectors; most notably the differences in (marginal) productivity between agriculture and nonagriculture (Caselli 2005; Restuccia, Yang, and Zhu 2008; Vollrath 2009; Gollin, Lagakos, and Waugh 2014). While this literature has devoted much attention to the relationship between misallocation, at the firm or sectoral level, and cross-country income differences (e.g., Parente and Prescott 1999; Lagos 2006; Buera and Shin 2013), relatively little is known about the determinants of the misallocation itself.

In India, the rural-urban wage gap, corrected for cost-of-living differences, is greater than 25 percent and has remained large for decades, as we document in this paper. One explanation for this large wage gap is that underlying market failures prevent workers from taking advantage of arbitrage opportunities. A second explanation, based on a recent paper by Alwyn Young (2013) is that the large wage

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gap solely reflects differences in skill between rural and urban workers. In Young's framework, there is perfect intersectoral mobility and the size of the wage gap is completely determined by differences in the skill intensity of production between the rural and urban sectors. It follows that a country with an exceptionally large wage gap, such as India, will be characterized by an exceptionally large flow of workers sorting on skill. In contrast with this prediction, and indicative of misallocation, we will see that internal migration is very low in India, both in absolute terms as well as relative to other countries of comparable size and level of economic development.

The rural-urban wage divide is not the only symptom of spatial labor misallocation in India. Rural wages differ substantially across Indian villages and districts, and studies of rural wage determination have shown that shifts in local supply and demand affect local wages, which would not be true if labor were spatially mobile (Rosenzweig 1978; Jayachandran 2006). It is not that spatial mobility in India is generally low. Almost all women leave their native village upon marriage (Rosenzweig and Stark 1989). The question is why rural *male* workers have not taken advantage of the substantial economic opportunities associated with spatial wage differentials in India to move permanently to the city.

The explanation we propose, in the spirit of Banerjee and Newman (1998), is based on a combination of well-functioning rural insurance networks and the absence of formal insurance, which includes government safety nets and private credit. In rural India, informal insurance networks are organized along caste lines. The basic marriage rule in India, which recent genetic evidence indicates has been binding for 1,900 years, is that no individual is permitted to marry outside the subcaste or *jati* (for expositional convenience we will use the term *caste*, interchangeably with *subcaste*, throughout the paper). Frequent social interactions and close ties within the caste, which consists of thousands of households and spans a wide area covering many villages, support very connected and exceptionally extensive insurance networks (Caldwell, Reddy, and Caldwell 1986; Mazzocco and Saini 2012).

Households with migrant members will have reduced access to rural caste networks for two reasons. First, migrants cannot be as easily punished by the network, and their family back home in the village now has superior outside options (in the event that the household is excluded from the network). It follows that households with migrants cannot credibly commit to honoring their future obligations at the same level as households without migrants. Second, an information problem arises if the migrant's income cannot be observed. If the household is treated as a collective unit by the network, it always has an incentive to misreport its urban income so that transfers flow in its direction. If the resulting loss in network insurance from migration exceeds the income gain, then large wage gaps could persist without generating a flow of workers to higher-wage areas. Just as financial frictions distort the allocation of capital across *firms* in Buera, Kaboski, and Shin (2011), the absence of formal insurance distorts the allocation of labor across *sectors* in the model that we develop below. This distortion is paradoxically amplified when the informal insurance networks work exceptionally well because rural households then have more to lose by sending their members to the city.

One way to circumvent these restrictions on mobility would be for members of the rural community to move to the city (or another rural location) as a group. Members of the group could monitor each other and enforce collective punishments,

solving the information and commitment problems described above. They would also help each other find jobs at the destination. The history of industrialization and urbanization in India is indeed characterized by the formation and the evolution of caste-based urban networks, sometimes over multiple generations (Morris 1965; Chandavarkar 1994; Munshi and Rosenzweig 2006). A limitation of this strategy is that a sufficiently large (common) shock is needed to jump-start the new network at the destination, and such opportunities occur relatively infrequently (Munshi 2011). Thus, while members of a relatively small number of castes with (fortuitously) well established destination networks can move with ease, most potential migrants will lack the social support they need to move.

A second strategy to reduce the information and enforcement problems that restrict mobility is to migrate temporarily. Seasonal temporary migration has, in fact, been increasing over time in India (Morten 2013). The principal limitation of the temporary migration strategy is that it will not fill the large number of jobs in developing economies in which there is firm-specific or task-specific learning and where firms will set permanent wage contracts.

Both strategies discussed above will be used by rural households and castes to facilitate mobility. However, the central hypothesis of this paper is that most men will nevertheless be discouraged by the loss in insurance from migrating and the labor market will not clear, giving rise to the large spatial wage gaps *and* the low male permanent migration rates that motivate our analysis.¹ Previous studies have also made the connection between insurance networks and migration in India. Rosenzweig and Stark (1989) show that marital migration by women extends network ties beyond village boundaries. Morten (2013) links opportunities for temporary migration to the performance of rural networks. Both of these studies take participation in the network as given, whereas we hypothesize that permanent male migration can result in the exclusion of entire households from the network. The simplest test of the hypothesis that this potential loss in network services restricts mobility in India would be to compare migration rates in populations with and without caste-based insurance. This exercise is infeasible, given the pervasiveness of caste networks. What we do instead is to look *within* the caste and theoretically identify which households benefit less (more) from caste-based insurance. We then proceed to test whether it is precisely those households that are more (less) likely to have migrant members.

When an insurance network is active, the income generated by its members is pooled in each period and then distributed on the basis of a prespecified sharing rule. This smooths consumption over time, making risk-averse individuals better off. The literature on mutual insurance is concerned with *ex post* risk-sharing, taking the size of the network and the sharing rule as given.² To derive the connection between

¹ While we provide a specific risk-based mechanism to explain large rural-urban wage gaps in India, the literature on international migration merely postulates the existence of “migration costs” to explain the persistence of global wage inequalities (e.g., Chiquiar and Hanson 2005; McKenzie and Rapoport 2010).

² With complete risk-sharing, the sharing rule is independent of the state of nature, generating simple statistical tests that have been implemented with data from numerous developing countries. The general result is that high levels of risk-sharing are sustained, but complete risk-sharing is rejected (e.g., Townsend 1994; Grimard 1997; Ligon 1998; Fafchamps and Lund 2003; Angelucci, De Giorgi, and Rasul 2015). These empirical regularities have led, in turn, to a parallel line of research that characterizes and tests state (and history) dependent sharing rules under partial insurance (Coate and Ravallion 1993; Udry 1994; Ligon, Thomas, and Worrall 2002). The benchmark

networks and permanent migration, however, it is necessary to take a step back and model the *ex ante* participation decision and the optimal design of the income sharing rule. In our framework, households can either remain in the village and participate in the insurance network or send one or more of their members to the city, increasing their income but losing the services of the network. The sharing rule that is chosen in equilibrium determines which households choose to stay.

With diminishing marginal utility, the total surplus generated by the insurance arrangement can be increased by redistributing income so that relatively poor households consume more than they earn on average. This gain from redistribution must be weighed against the cost to the members of the network from the accompanying decline in its size, since relatively wealthy households will now be more likely to leave and smaller networks are less able to smooth consumption. We are able to show, under reasonable conditions, that the income sharing rule will nevertheless be set so that there is some amount of redistribution in equilibrium. This implies that relatively wealthy households within their caste benefit less from the network and so will be more likely to have migrant members *ceteris paribus*, providing the first prediction of the theory.

Our analysis is related, yet distinct in important respects, from Abramitzky (2008) who studies redistribution and exit in Israeli kibbutzim. For an exogenously determined (equal) income-sharing rule, he shows that exit rates are decreasing in communal wealth (which is forfeited upon exit) and that those with superior outside options are more likely to leave. In our model, the wealthy do not have superior outside options, wealth is private and is not forfeited, and the decision to participate and the income-sharing rule are endogenously and jointly determined. In a second model, Abramitzky uses diminishing marginal utility, as we do, to motivate redistribution. However, the sharing rule is chosen such that there is no *ex post* exit once individuals' abilities and outside options are revealed. Genicot and Ray (2003), in contrast, endogenize the size of the risk-sharing arrangement, but assume that all individuals are *ex ante* identical, which implies an equal sharing rule by construction. Our model endogenizes both the size of the network (and complementary migration) as well as the sharing rule, in a framework with heterogeneous households that builds naturally on existing models of *ex post* risk-sharing.³

While women's migration at marriage diversifies the income of the network, migration by a male household member diversifies the household's income and so is typically assumed to lower the income risk that the household faces (e.g., Lucas and Stark 1985). The implicit assumption in our framework is that in the Indian context, the loss in network insurance when an adult male from the household migrates dominates this gain from income diversification. It follows that households who face

sharing rule, in the initial period and when the participation constraint does not bind, continues to be exogenously determined in these models.

³Other studies in the migration literature, e.g., McKenzie and Rapoport (2007) and Stark and Taylor (1991) also consider the relationship between relative wealth and migration. We focus on the effect of wealth inequality in the origin community on migration, whereas McKenzie and Rapoport study how migration changes inequality in the sending community. Stark and Taylor study how wealth inequality determines migration, as we do, but their theoretical predictions are driven mechanically by an unverifiable assumption about individual preferences, which is that relatively poor (deprived) households in the sending community have a greater incentive to migrate as a way of closing the wealth gap with their neighbors. Their prediction is at odds with the data, since we find (consistent with our theory) that relatively wealthy households are more likely to have migrant members.

higher rural income risk and who, therefore, benefit more from the network *ceteris paribus*, will be less likely to have male migrant members. This second prediction is especially useful in distinguishing our theory from alternative explanations for large rural-urban wage gaps and low migration in India. One alternative explanation for the lack of mobility is that individuals cannot enter the urban labor market without the support of a (caste) network at destination. There are also alternative explanations (discussed below) available for redistribution within the caste and the increased exit from the network by relatively wealthy households. However, none of these explanations imply that households facing greater rural income risk should be less likely to have migrant members.

We begin the assessment of the theory by showing that there is substantial redistribution of income within castes, using data from the Indian ICRISAT panel surveys and from the most recent (2006) round of the Rural Economic and Development Survey (REDS), a nationally representative survey of rural Indian households that has been administered by the National Council of Applied Economic Research at multiple points in time over the past four decades. Following up on this result, we show (using data from a census of villages covered in the 2006 REDS) that *relatively* wealthy households within their caste are significantly more likely to report that one or more adult male members have permanently left the village.⁴ The literature on migrant selection, e.g., McKenzie and Rapaport (2007, 2010), Munshi (2011), indicates that migrant networks at destination support the movement of weaker—less able, less educated, less wealthy—individuals. In our analysis, insurance networks at the origin disproportionately *discourage* the movement of (relatively) less wealthy individuals. Highlighting the role that rural income risk plays in the migration decision, we also find that households with a higher coefficient of variation in their (rural) income—who benefit more from the rural insurance network—are less likely to have migrant members.⁵

Having found evidence consistent with the theory, we proceed to estimate the structural parameters of the model. Migration and the income-sharing rule are determined endogenously in the model. Our estimates of the income-sharing rule indicate that there is substantial redistribution within the caste, consistent with the descriptive evidence and the tests of the theory. Counterfactual simulations that quantify the effect of formal insurance on migration, leaving the rural insurance network in place, indicate that a 50 percent improvement in risk-sharing for households with migrant members (which is still some way from full risk-sharing) would more than double the migration rate, from 4 to 9 percent. In contrast, (nearly) halving the rural-urban wage gap, from 18 percent to 10 percent, without any change in formal insurance, would reduce migration by just 1 percentage point.

⁴We subject this result to robustness tests that (i) use alternative measures of income and independent datasets, and (ii) that examine the relationship between the household's relative wealth and its participation in the caste-based insurance network. The latter test allows us to verify a key assumption of our model, and that of Banerjee and Newman (1998), which is that migration should be associated with a loss in network services.

⁵We assume in the model that entire households do not migrate, consistent with evidence provided below, and that households with migrant members are treated by the network as a single collective unit. If entire households did migrate, or if individual migrants and the family members they left behind were treated independently by the network, then we would expect rural income risk to be *positively* associated with migration.

I. Descriptive Evidence

This section begins by documenting the exceptionally large rural-urban wage gap in India and its exceptionally low migration rates. We subsequently describe the role played by rural caste networks in providing insurance for their members. The theory developed in the next section is based on the premise that migration is accompanied by a loss in these network services, connecting rural caste networks to the low mobility, and accompanying labor misallocation, we have documented. This connection will be subjected to greater scrutiny in the empirical analysis that completes the paper.

A. Rural-Urban Wage Gaps and Migration

An important indicator of spatial immobility is the rural-urban wage gap. To measure the rural-urban wage gap in India we use the Government of India's sixty-first National Sample Survey (NSS) covering the period July 2004–June 2005. Schedule 10 provides, for a given week during the survey period, the total number of days each person worked and, for workers classified as regular salaried employees or casual wage laborers, their wage and salary earnings both in cash and in kind. Based on this information, we computed a daily wage for each rural and urban worker.⁶ Column 1 of Table 1 reports the mean of these wages for rural and urban workers with less than primary education. We focus on this group to avoid the confounding effects of differences in the returns to education in rural and urban labor markets. Workers with little education will perform similar—menial—tasks in both markets, and so wage gaps for them can be interpreted as an arbitrage opportunity. The gap that we compute is very large; the urban wage is over 47 percent higher than the rural wage. As a basis for comparison, Figure 1 provides the percentage rural-urban wage gap in two large developing countries—China and Indonesia—computed from the 2005 Chinese mini Census and the Indonesia Family Life Survey (IFLS) 4 (2007), respectively.⁷ As can be seen, the wage gap for India, at over 45 percent, is much higher than the corresponding gap for the other two countries, which is about 10 percent.

One reason that urban wages are higher than rural wages is that the cost of living may differ across rural and urban areas. If the same bundle of goods consumed in urban areas costs more in rural areas, then the wage gap in column 1 of Table 1 may overstate the real gain in earnings from migration. To adjust the wages for purchasing parity, we used the consumption information provided in Schedule 1.0 from the same NSS. Schedule 1.0 provides the value and quantity for durable and nondurable goods consumed by rural and urban households, enabling the computation of rural and urban unit prices. Table 1, column 2 reports the urban wage deflated by the Laspeyres index (rural or origin base) and thus the real rural-urban wage gap.

⁶The NSS, as do other Indian datasets, defines the urban population to include residents of cities and towns that exceed a population-size threshold. This threshold has changed over time, as discussed below.

⁷The wage for Indonesia is the hourly wage based on payments and wage work in the week preceding the survey for male wage workers aged 25–49 with less than primary school completion. Forty-eight percent of rural male workers were in that schooling category. The cross-sectional weights with attrition were used to compute the urban and rural means. The hourly wage for China is also for men aged 25–49 in the same educational category.

TABLE 1—RURAL-URBAN WAGE GAPS IN INDIA IN 2004

Sector	Wage		
	Nominal (1)	PPP-adjusted (rural consumption) (2)	PPP-adjusted (urban consumption) (3)
Urban	62.66	54.05	57.58
Rural	42.54	42.54	42.54
Percent gain	47.30	27.06	35.35

Notes: Wages are measured as daily wages for individuals with less than primary education. PPP-adjustment is based on rural and urban consumption bundles, respectively, for those individuals.

Source: NSS

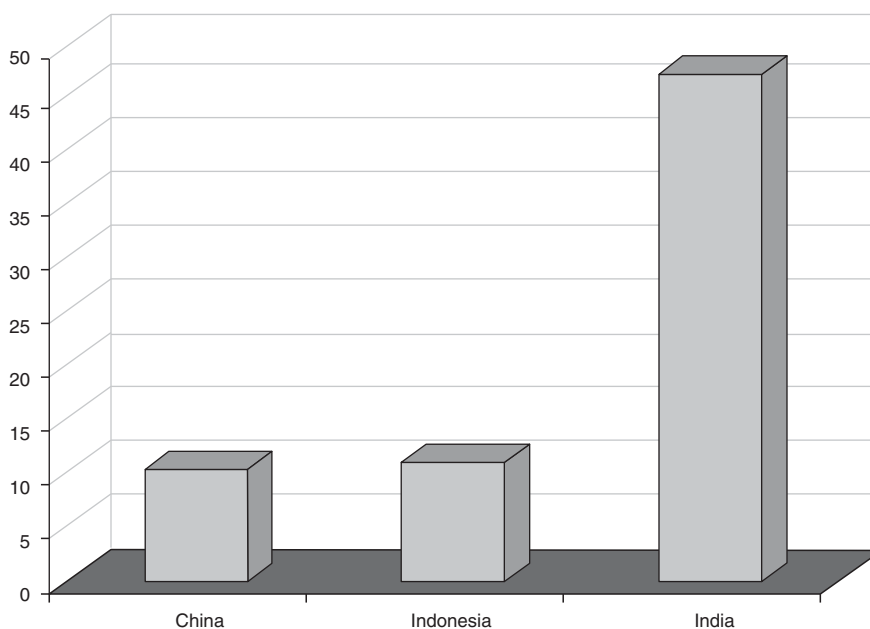


FIGURE 1. RURAL-URBAN WAGE GAP, BY COUNTRY

Sources: Chinese mini-census 2006, IFLS 2007, and NSS 2004

The PPP-adjusted urban wage is the nominal urban wage, multiplied by the value of the consumption bundle of rural households whose heads have less than primary education and then divided by the value of the same bundle based on urban prices. As can be seen, while this correction for standard of living substantially cuts the earnings advantage from shifting from rural to urban employment, there is still a real wage gap of over 27 percent. To assess the sensitivity of our results to the choice of consumption bundle, we used the corresponding urban consumption bundle, appropriately priced for rural and urban areas, to deflate the nominal urban wage. Using this destination-based deflator (the Paasche index), the real wage gap is

even higher at over 35 percent.⁸ Although the Chinese and Indonesian data we use to construct the wage gaps in Figure 1 do not allow us to construct the corresponding PPP-adjusted statistics, the nominal gaps provide us with an upper bound on the real gaps since urban wages will always be higher than rural wages. It follows that the real wage gap in India is at least 16 percentage points larger than it is in China and Indonesia.

It is possible that the 2004–2005 year was peculiar. To gauge how the real wage gap has changed over time in India we use the nominal rural and urban wages estimated from the NSS rounds for 1983–1984, 1993–1994, 1999–2000, 2004–2005, and 2009–2010 by Hnatkowska and Lahiri (2014) to compute the real urban and rural wages. First we apply our PPP correction to the urban wage series using the rural consumption bundle and unit prices from the 2004–2005 NSS. We then apply the agricultural-worker CPI series and the industrial-worker CPI series to the PPP-adjusted rural and urban wage series, respectively, to obtain an inflation- and PPP-adjusted real wage series. Appendix Table A1 provides the nominal wages, the CPI figures, and the deflated wages by year for rural and urban workers. Figure 2 plots the movements in these wages over time. As can be seen, the real wage gap in 2004–2005 actually underestimates the average wage gap over the period 1983–2009. After a sharp decline between 1999 and 2004, the wage gap remains stable from 2004–2005 through 2009–2010 at over 20 percent. This stability contrasts once again with changes over time in other countries. Using successive rounds of the IFLS, and adjusting for inflation, the nominal wage gap in Indonesia declined from 72 percent in 1993 to 11 percent in 2007. This is what we would expect as infrastructure improved with economic development, facilitating increased migration over time. Based on the NSS statistics reported in Appendix Table A1, the inflation-adjusted nominal wage gap in India declined by much less, from 59 percent in 1993 to 30 percent in 2009, and most of this change can be accounted for by the decline in the wage gap between 1999 and 2004.

The change in the wage gap between 1999 and 2004 has two potential causes; a change in the definition of “urban” and the general-equilibrium effect of increased rural-to-urban migration. Hnatkowska and Lahiri conclude that almost all of the change in the gap is due to the changing criteria for urbanization. By reclassifying some rural populations as urban, one would expect that the average urban wage would decrease but with possibly little effect on average rural wages. This is exactly what we see in Figure 2; when there is a decline in the wage gap, it is almost wholly due to a sharp urban wage decline. If the decline in the wage gap was due to rural-urban migration, then urban wages would decline *and* rural wages would

⁸ As originally pointed out in Harris and Todaro (1970), migration responds to the expected wage; that is, the potential migrant takes into account the probability of employment. Although in that article the emphasis was on unemployment in urban areas, unemployment in rural areas potentially matters as well. The NSS elicited, in Schedule 10, information on employment and unemployment in the past year for all workers. The survey provides for each worker the number of months without work and whether, if without work, the worker made any efforts to get work on some or most days. From this information we computed the fraction of the year a worker was employed and/or unemployed for both rural and urban workers. Interestingly, but perhaps unsurprisingly given the seasonality of agriculture, nonemployment and unemployment rates are higher in rural than in urban areas. We weighted real wages (where the nominal urban wage is deflated using the rural consumption bundle) by the rate of employment (fraction of the year employed) and by the fraction of days not unemployed, respectively. The expected earnings gain from migration using these figures is higher than the employment-unadjusted real wage gap (column 2), lying between 32 percent and 35 percent.

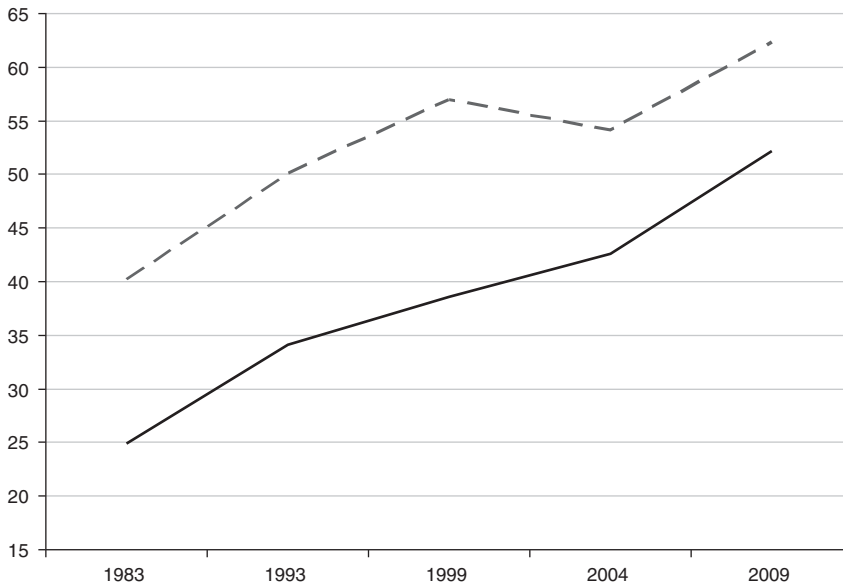


FIGURE 2. REAL RURAL AND URBAN WAGES IN INDIA

Source: NSS 1983–2009

increase. To provide additional support for the claim that the decline in the wage gap between 1999 and 2004 is not being driven by migration, we report migration rates based on decadal population censuses over the 1961–2001 period. Following Foster and Rosenzweig (2008), migration rates are computed for the cohort of males aged 15–24 (who are most likely to move for work) within each decade by comparing their numbers, residing permanently in rural and urban areas, at the beginning and the end of the decade.⁹ These migration rates are plotted in Figure 3, where no spike in migration is visible in the 1991–2001 period. Despite the persistently large (real) wage gaps that we have documented, rural-urban migration in India has remained low for decades, reaching a maximum of 5.4 percent in the earliest period and dropping below 4 percent in recent decades.¹⁰

It is possible that the wage gap we quantify (conditional on education) merely reflects sorting on unobserved skill, and a large difference in the skill intensities of production between rural and urban areas of India, as suggested by Young's (2013) model. We do not think sorting on skill explains the large wage gap in India. First, agriculture became more skill-intensive as a result of the Green Revolution in many

⁹This method requires that mortality rates are similar across urban and rural populations. In the age group 15–24, mortality is very low. The method also assumes that definitions of rural and urban remain constant across the decade. The urbanizing of the population by redefinition, as described above, will inflate the migration rates computed using the cohort method. The rates that are computed are thus likely to be upper bounds on true migration. The 2001 census indicates that movement due to marriage by women accounts for roughly 45 percent of all permanent migration in India, while employment, business, and the movement of entire families accounts for just 39 percent of migration (similar statistics are obtained in the 1991 round). We consequently focus on male out-migration when measuring the spatial mobility that is associated with the rural-urban wage gap.

¹⁰Although the detailed information needed to compute the migration rate from 2001 to 2011 is currently unavailable, provisional figures from the latest 2011 census indicate that the proportion of the population that is urban rose by only 3.8 percentage points between 2001 and 2011, to 31.6 percent (Office of the Registrar General and the Census Commissioner 2011).

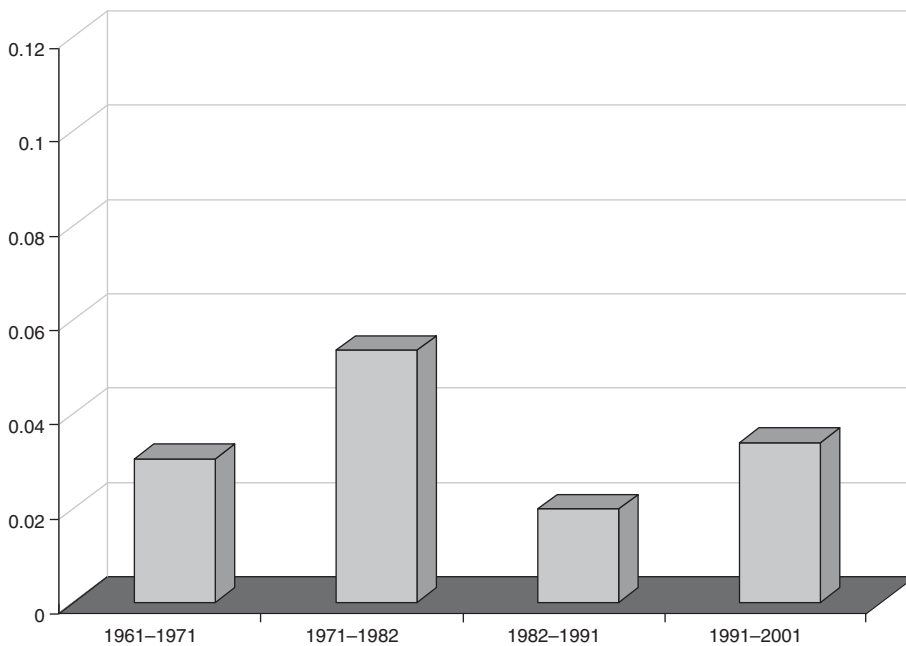


FIGURE 3. CHANGE IN RURAL-URBAN MIGRATION RATES IN INDIA, 1961-2001

Source: Indian Population Census 1961-2001

parts of India starting in the 1970s and prior to the economic reforms of the 1990s (Foster and Rosenzweig 1995). In contrast, TFP growth in manufacturing was close to zero or even declining during this period (Balakrishnan and Pushpangadan 1994; Saha 2014). Young's model would predict that the wage gap would therefore have declined in that period. It did not. Second, Young's model implies that migration rates from rural to urban and from urban to rural areas should both be high where wage gaps are high to achieve the appropriate mix of skills in both sectors. But in India, both urban and rural out-migration rates are low. An independent measure of migration can be constructed from the nationally representative India Human Development Survey (IHDS) conducted in 2005, which covers both rural and urban areas. The survey provides information on the number of years that each sampled household has been residing in the current location. We assume that a household has in-migrated if it has resided in that location for less than ten years. Based on this definition, and restricting attention to households with male heads aged 25-49, the IHDS can be used to compute urban-rural and rural-urban migration rates. These statistics are 1.06 percent and 6.48 percent, respectively. Using the same definitions applied to the male subsample of the 2005 Indian Demographic and Health Survey (DHS), the rates are 5.55 and 5.34 percent. There is thus no evidence that the exceptionally large wage gap in India is accompanied by a commensurate flow of workers, in either direction, refuting the counterargument that these gaps simply reflect differences in (unobserved) skill.¹¹ Even with the DHS statistics, which are substantially

¹¹ Young (2013) reports balanced urban-rural and rural-urban migration rates above 20 percent in his sample of 65 countries. He uses DHS data and pools information on men and women. Men make up 10 percent of the

higher than the corresponding IHDS statistics, migration rates are much lower in India than in countries of similar size and levels of economic development. For example, the 1997 Brazil DHS, which also includes a male sample, reports that urban-rural and rural-urban migration rates are 4.55 percent and 13.9 percent. The rural-urban migration rate, in particular, is more than twice as large as India.

India's unusually low mobility is also reflected in its urbanization rates. Figure 4 plots the percent of the adult population living in the city, and the change in this percentage over the 1975–2000 period, for four large developing countries: Indonesia, China, India, and Nigeria (UNDP 2002). Urbanization in all four countries was low to begin with in 1975 but India falls far behind the rest by 2000. Deshingkar and Anderson (2004) show that rates of urbanization in India are lower, by 1 full percentage point, than countries with similar levels of urbanization, and that the fraction of the population that is urban in India is 15 percent lower than in countries with comparable GDP per-capita. The exceptionally low mobility in India, despite the apparent benefit from moving to the city, demands an explanation. This is what we turn to next.

B. Rural Insurance Networks

In this section we show that transfers (gifts and loans) from caste members are important and preferred mechanisms through which consumption is smoothed in rural India. Much of the evidence is based on the 1982 and 1999 REDS rounds, which covered 259 villages in 16 major Indian states. Table 2 reports the percentage of households in the two survey rounds who gave or received caste transfers, which include gift amounts sent and received as well as loans originating from or provided to fellow caste members, in the year prior to each survey. The table shows that even in a single year, participation in the caste-based insurance arrangement is high—25 percent of the households in the 1982 survey and 20 percent in the 1999 round.¹² We would expect multiple households to support the receiving household when it is in need of help and consistent with this view, sending households contribute 5–7 percent of their annual income on average whereas the corresponding statistic for receiving households is 20–40 percent.¹³

A variety of financial instruments are used to smooth consumption within the caste, with caste loans accounting for just 23 percent of all within-caste transfers by value. Nevertheless, the 1982 survey data in Table 3 indicate that although banks are the dominant source of rural credit, accounting for 64.6 percent of all loans by value, caste members are the dominant source of informal loans, making up 13.9 percent of the total value of loans received by households in the year prior to the survey.¹⁴ This is more than the amount households obtained from moneylenders

sample. This is evidently unsatisfactory for India where 88 percent of women move outside their village when they marry (IHDS 2005). These women are not moving to clear the labor market, and the same problem arises in all other patrilineal societies in his sample. This is why we focus on male migrants in the discussion above.

¹²The statistics in Table 2 are weighted using sample weights and thus are population statistics.

¹³Some of these differences arise because sending households have higher income on average than receiving households, indicative of redistribution within the caste that will play an important role in the discussion that follows. Nevertheless, it is easy to verify that the amount sent per household is less than the amount received.

¹⁴We restrict attention to the 1982 survey because the classification of activities that loans are used for is much coarser in 1999; in particular, consumption expenses do not appear as a separate category.

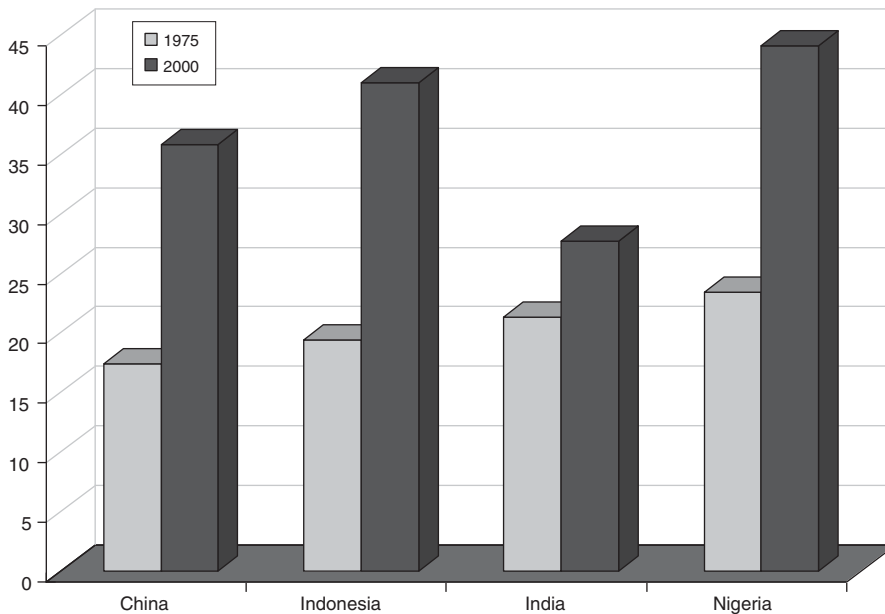


FIGURE 4. CHANGE IN PERCENT URBANIZED, BY COUNTRY, 1975–2000

Source: UNDP 2002

TABLE 2—PARTICIPATION IN THE CASTE-BASED INSURANCE ARRANGEMENT

Survey year	1982 (1)	1999 (2)
Households participating (percent)	25.44	19.62
Income of senders	5,678.92 (7,617.55)	19,956.29 (22,578.95)
Percent of income sent	5.28	8.74
Income of receivers	4,800.29 (4,462.63)	10,483.84 (13,493.68)
Percent of income received	19.06	40.26
Observations	4,981	7,405

Notes: Standard deviations in parentheses. Participation in the insurance arrangement includes giving or receiving gifts and loans. Participation measured over the year prior to each survey round. Income is measured in 1982 rupees.

Source: REDS 1982 and 1999

(7.9 percent), friends (7.8 percent), and employers (5.6 percent). Table 3A reports the proportion of loans in value terms, both by source and purpose, using data from the 1982 REDS. As can be seen, caste loans are disproportionately used to cover consumption expenses and for meeting contingencies such as illness and marriage. For example, although loans from caste members were 14 percent of all loans in value, they were 23 and 43 percent, respectively, of the value of all consumption

TABLE 3A—PERCENT OF LOANS BY PURPOSE AND SOURCE: REDS 1982

Purpose	Investment	Operating expenses	Contingencies	Consumption expenses	All
	(1)	(2)	(3)	(4)	(5)
<i>Share</i>	0.15	0.60	0.13	0.07	1.00
<i>Source</i>					
Bank	64.11	80.80	27.58	25.12	64.61
Caste	16.97	6.07	42.65	23.12	13.87
Friends	2.11	11.29	2.31	4.33	7.84
Employer	5.08	0.49	21.15	15.22	5.62
Moneylender	11.64	1.27	5.05	31.85	7.85
Other	0.02	0.07	1.27	0.37	0.22
Total	100.00	100.00	100.00	100.00	100.00

Notes: Statistics are weighted by the value of the loan and sample weights. Statistics computed using 982 loans received in the year prior to the 1982 survey round. Investment includes land, house, business, etc. Operating expenses are for agricultural production. Contingencies include marriage, illness, etc. Other category not reported for purpose.

Source: REDS 1982

TABLE 3B—PERCENT OF LOANS BY PURPOSE AND SOURCE: IHDS 2005

Purpose	2005 IHDS				
	Investment	Operating expenses	Contingencies	Consumption expenses	All
	(1)	(2)	(3)	(4)	(5)
<i>Share</i>	0.25	0.46	0.21	0.04	1.00
<i>Source</i>					
Bank	46.79	62.49	18.78	19.82	46.70
Caste	7.82	4.11	19.64	14.24	9.12
Friends	6.01	3.33	8.28	7.09	5.38
Employer	3.31	0.54	1.11	1.85	1.23
Moneylender	20.69	12.82	46.80	53.65	24.67
Other	15.38	16.71	5.39	3.35	12.90
Total	100.00	100.00	100.00	100.00	100.00

Notes: Statistics are weighted by the value of the loan and sample weights. Statistics computed using 12,066 rural loans received in the year prior to the 2005 IHDS. IHDS 2005 reports loans received from relatives rather than caste. Investment includes land, house, business, etc. Operating expenses are for agricultural production. Contingencies include marriage, illness, etc. Other category not reported for purpose.

Source: IHDS 2005

and contingency loans.¹⁵ In contrast, bank loans are by far the dominant source of finance for investment and operating expenses, but account for just 25 percent and 28 percent of loans received for consumption expenses and contingencies.

Are the statistics in Table 3A, representing the rural population of India in 1982, comparable to the current period? Table 3B describes loans by source and purpose

¹⁵ Caldwell, Reddy, and Caldwell (1986) surveyed nine villages in South India after a two-year drought and found that nearly half (46 percent) of the sampled households had taken consumption loans during the drought. The sources of these loans (by value) were government banks (18 percent), moneylenders, landlord, employer (28 percent), relatives and members of the same caste community (54 percent), emphasizing the importance of caste loans for smoothing consumption.

using the IHDS (2005). This survey, conducted on a representative sample of rural households throughout the country, reports loans received over the five years preceding the survey by source. Unfortunately the survey does not use caste-group as a category, although it does identify loans from relatives, which we will assume are within-caste loans. Although some caste loans will now be assigned to other categories (if they are provided by caste members not directly related to the recipient), the basic patterns reported from the 1982 survey round in Table 3A remain unchanged. Loans from relatives make up 9 percent of all loans by value, more than both friends and employers. Looking across purposes, we see once again that informal caste loans are most useful in smoothing consumption and meeting contingencies. Overall, lending patterns have remained fairly constant over the two decades covered in Tables 3A and 3B.¹⁶

We argue in this paper that caste networks restrict mobility because comparable arrangements are unavailable, particularly for smoothing consumption and meeting contingencies. Table 4 shows that loan terms are substantially more favorable for caste loans on average. It is quite striking that of the caste loans received in the year prior to the 1982 survey, 20 percent by value required no interest payment and no collateral. The corresponding statistic for the alternative sources of credit was close to zero, except for loans from friends where 4 percent of the loans were received on similarly favorable terms. The IHDS does not provide information on collateral but does report whether a loan was interest-free. We see in Table 4, column 5 that caste (extended family) loans are substantially more likely to be interest-free than loans from other sources, matching the corresponding statistics from the 1982 REDS in column 1.¹⁷

Tables 3 and 4 establish that loans from caste members are important for smoothing consumption and meeting contingencies, and continue to be advantageous to borrowers compared with loans from major alternative sources of finance in rural India. It is important to reiterate that these caste loans account for a small fraction of all within-caste transfers by value. The cost of losing the services of the network is evidently substantial and may explain why individuals continue to marry within their subcaste, which is a prerequisite for membership in the caste network, today.

Figure 5 reports rates of out-marriage (i.e., marriage between members of different castes) in rural India for the children and siblings of household heads over the 1950–1999 period, based on retrospective information collected in the 1999 REDS round. Out-marriage is just above 5 percent of all marriages, closely matching other sample surveys conducted in urban and rural India (IHDS 2005; Munshi and Rosenzweig 2006; Luke and Munshi 2011), and has remained stable over time. Recent genetic evidence indicates that binding restrictions on out-marriage were put in place 1,900 years ago and that the Indian population today consists of 4,635 distinct genetic groups (Moorjani et al. 2013).¹⁸ These groups consist of thousands

¹⁶NGOs and credit groups, which have received a great deal of attention in the economics literature in recent years are included in the “Other” category in the IHDS. However, these sources together account for less than 2.1 percent of all loans by value received by rural households.

¹⁷Regression results with 1982 REDS data, reported in Appendix Table A2, indicate that caste loans are significantly more likely to be interest-free than loans from banks, employers, and moneylenders. They are also significantly more likely to be collateral-free than loans from banks.

¹⁸These genetic groups are not restricted to the Hindu population. Muslims marry within *biradaris* and Christians continue to marry within their original (pre-conversion) subcastes or *jatis*. In our dataset, Muslim households report their *biradari* and Christian households report their *jati*.

TABLE 4—PERCENT OF LOANS BY TYPE AND SOURCE

Loan type:	REDS 1982			IHDS 2005
	Without interest (1)	Without collateral (2)	Without collateral or interest (3)	Without interest (4)
<i>Source</i>				
Bank	0.57	23.43	0.38	0.00
Caste	28.99	60.27	20.38	44.62
Friends	9.35	91.72	3.89	21.5
Employer	0.44	65.69	0.44	10.75
Moneylender	0.00	98.71	0.00	0.27

Notes: Statistics are weighted by the value of the loan and sample weights. Columns 1–3 computed using 982 loans received in the year prior to the 1982 survey round. Column 4 computed using 12,066 rural loans received in the 5 years prior to the 2005 IHDS. IHDS 2005 reports loans received from relatives rather than caste.

Sources: REDS 1982 and IHDS 2005

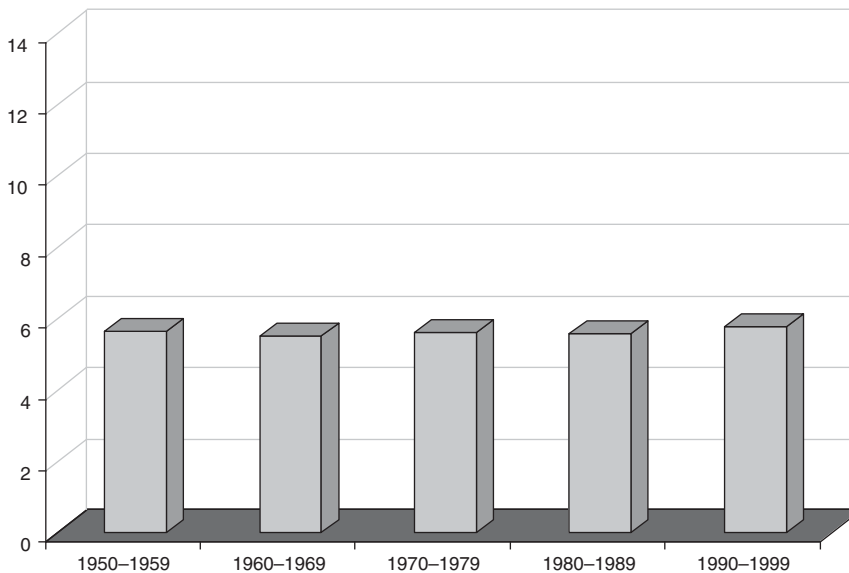


FIGURE 5. CHANGE IN OUT-MARRIAGE PERCENT IN RURAL INDIA, 1950-1999

Source: REDS 1999

of households. Marital endogamy, together with the fact that women typically marry outside their natal village, allows caste networks to span wide areas, while maintaining their connectedness. This connectedness across villages is complemented by strong local ties, which arise as a consequence of the spatial segregation by caste within villages. Households that renege on their obligations will thus be punished locally (in the neighborhood) and in the wider caste community. Information will also flow very smoothly through this interlinked community. The analysis that follows examines the effect of these exceptionally well-functioning caste networks on mobility and the rural-urban wage gap.

The central assumption in our analysis is that men migrating independently (and permanently) to the city cannot be monitored effectively by their rural communities and so will be excluded from rural-based insurance networks. By the same argument, caste networks will not be able to function effectively if their members are spread thinly over a very wide rural area. However, we also note that migration can be sustained without the loss of network insurance if members of a caste move together as a group. The group can then monitor its members in the city. A caste could use an analogous strategy to support cooperation and reduce information problems when its members are spread over a wide area. A single caste will not have a presence in each village, but instead will cluster in select villages. This clustering shows up clearly in the 2006 REDS census, where the mean number of castes per state is 64, while the mean number of castes per village is 12. With 340 households on average in a village, this implies that a caste will have about 30 households in those villages where it is represented.¹⁹

II. The Theory

Our theory describes how the existence of well-functioning rural insurance networks can lead to low migration. The theoretical structure we develop will be taken to the data, allowing us to quantify the magnitude of the mobility restrictions. It will also be used to generate testable predictions that distinguish it from alternative explanations for the low mobility in India.

A. Income, Preferences, and Risk-Sharing

The basic decision-making unit is the household, which consists of multiple earners. The household belongs to a community within which all its social activities take place. Each household derives income from its local activities. Income varies independently across households in the community and over time. In addition, one or more members of the household receive a job opportunity in the city. The key decision is whether or not to send them to the city.

We assume that the household has logarithmic preferences. This allows us to express the expected utility from consumption, C , as an additively separable function of mean consumption, M , and normalized risk, $R \equiv V/M^2$, where V is the variance of consumption:²⁰

$$EU(C) = \log(M) - \frac{1}{2} \frac{V}{M^2}.$$

¹⁹The pattern of spatial clustering we have documented has theoretical foundations. Jackson, Rodriguez-Barraquer, and Tan (2012) examine reciprocity in societies where any two individuals interact too infrequently to support exchange but where the possible loss of multiple relationships (in the event of default) can be used to support cooperation. They show that robust networks in such settings are social quilts: tree-like unions of completely connected subnetworks. Based on the statistics reported above, caste networks appear to exhibit precisely these properties.

²⁰This expression is obtained by evaluating log consumption at mean consumption, M , and ignoring higher-order terms. For the Taylor expansion to be valid, with CRRA preferences, consumption must lie in the range $[0, 2M]$. This implies that its coefficient of variation must be less than 0.31. The panel data that we use, described below, satisfies this condition for 90 percent of households with food consumption and 70 percent of households with overall consumption (which includes durables).

Rural incomes vary over time and so risk-averse households benefit from a community-based insurance network to smooth their consumption. Because our interest is in the ex ante decision to participate in the rural insurance network, we assume that complete risk-sharing can be maintained ex post (once the arrangement has formed). The advantage of this assumption is that it allows us to derive closed-form solutions for the mean and variance of consumption with insurance that lead, in turn, to a simple migration decision-rule. This simplifies the theoretical analysis and later allows us to estimate a parsimonious structural model. As noted in the introduction, this assumption is, moreover, consistent with evidence from all over the developing world, including India, documenting extremely high levels of ex post risk-sharing.

The ex post commitment that is needed to support these high levels of risk-sharing is maintained by social sanctions, which take the form of exclusion from social interactions within the community when a participating household reneges on its obligations. These sanctions are less effective when someone from the household has migrated to the city.²¹ With full risk-sharing, the household is either in the network, receiving a fixed fraction of the income generated by the membership in each state of the world, or out of the network. We assume that households with migrants cannot commit to reciprocating at the level needed for full risk-sharing and so will be excluded from the network.

Individuals migrate independently (and permanently) in our model. Their urban income is private information. If a household with migrants is included in the insurance network, it will thus have an incentive to overreport the value of its urban income ex ante, as a way of increasing its income share. Once the risk-sharing arrangement is in place, however, it will have an incentive to underreport its income realizations ex post, claiming a series of negative shocks, as a way of channeling transfers in its direction. Partial insurance, which ties transfers to income realizations, will reduce the cost to the network from this information problem, but it will not change the household's incentive to misreport its income. This "hidden income" problem is potentially more important than the commitment problem in explaining why households with migrants will be excluded from the network. Each household thus has two options. It can remain in the village and participate in the insurance network, benefiting from the accompanying reduction in the variance of its consumption, or it can send one or more of its members to the city and add to its income but forgo the services of the rural network.

B. *The Participation Decision*

Let M_A , V_A denote the mean and variance of the household's income (which is the same as its consumption in autarky) when all its members remain in the village. Denote the mean and variance of its consumption if it participates in the insurance network by M_I , V_I , respectively. If one or more members move to the city, its mean

²¹ While the community could punish the remaining members of the household, this is not as effective as punishing all members. One potential solution to this commitment problem would be for the remaining members to separate themselves from the migrants. This is not a credible strategy, however, because unobserved remittances can continue to flow within the household. The remaining household members also have better outside options (through their urban connection) which reduces their ability to commit.

income will increase to $M_A(1 + \tilde{\epsilon})$, where $\tilde{\epsilon}$ denotes the gain in income from urban wages net of any loss in rural income due to their departure. This gain in income must be traded off against the increased consumption risk that the household will face. With network insurance, (normalized) consumption risk is denoted by $R_I \equiv V_I/M_I^2$. When the household sends migrants to the city, it loses the services of the network and the corresponding risk is βR_A , where $R_A \equiv V_A/M_A^2$. The standard presumption is that the income diversification that accompanies migration will reduce the income risk that the household faces. Then $\beta < 1$ even if a household with migrants has no alternative mechanism through which it can smooth its consumption. As formal (nonnetwork) insurance becomes available, the risk-parameter β will decline even further. However, we continue to assume that migration increases the consumption risk that the household faces, $R_I < \beta R_A$. This is the wedge that restricts mobility and allows a wage gap to be sustained in our theory. Note that this key insight of our theory would apply with any model of ex post risk-sharing, as long as the reduced access to the network resulted in increased consumption risk for households with migrants.

With logarithmic preferences, the household will thus choose to participate in the rural insurance network and remain in the village if

$$(1) \quad \log(M_I) - \frac{1}{2} \frac{V_I}{M_I^2} \geq \log(M_A) - \frac{1}{2} \beta \frac{V_A}{M_A^2} + \epsilon,$$

where $\epsilon \equiv \log(1 + \tilde{\epsilon})$.²² Given the standard assumption in models of mutual insurance that there is no storage and no savings, full risk-sharing and log preferences imply that each household's consumption will be a fixed fraction of the total income, $\sum_i y_{is}$, that is generated by the N households in the insurance network in each state s of the world. Let mean rural income, M_A , be the same for all households to begin with. The income gain from migration, ϵ , is assumed to be uncorrelated with rural income and is private information, so it follows that total income will be distributed equally among the members of the network.

Taking expectations, or variances, over all states, the equal-sharing rule implies that

$$(2) \quad M_I = E\left(\frac{1}{N} \sum_i y_{is}\right) = \frac{1}{N}(NM_A) = M_A$$

$$(3) \quad V_I = V\left(\frac{1}{N} \sum_i y_{is}\right) = \frac{1}{N^2}(NV_A) = \frac{V_A}{N}.$$

Mean consumption with insurance, M_I , is equal to mean consumption under autarky, M_A . However, the variance of consumption with insurance, V_I , is less than the variance of consumption under autarky, V_A , for $N \geq 2$.

²²If the terms in inequality (1) describe per-period utility, then both sides of the inequality would be multiplied by $1/(1 - \delta)$ for an infinitely lived household with discount factor δ . This would have no effect on the results that follow.

C. Equilibrium Participation

Based on the decision rule specified by inequality (1), participation will depend on the gain from mutual insurance, $1/2\beta R_A - 1/2R_I$, versus the income gain from migration, which is ϵ since $\log(M_I) = \log(M_A)$. The key feature of equation (3) is that it implies that the gain from insurance depends on the endogenously-determined number of network participants, N , since V_I and, thus, R_I , is decreasing in N .

Because the gain from insurance depends on the decisions of other households in the community, the number of network participants, N , is the solution to a fixed-point problem. To determine the fraction of the population that participates in equilibrium, we first derive the threshold ϵ_I at which the participation condition holds with equality. Let the ϵ distribution be characterized by the function $F(\epsilon)$. We then set the fraction of the community that participates, $F(\epsilon_I)$, to be equal to N/P ,

$$(4) \quad \frac{N}{P} = F(\Delta M + \Delta R),$$

where P is the population of the community, $\Delta M \equiv \log(M_I) - \log(M_A)$, $\Delta R \equiv 1/2\beta R_A - 1/2R_I$. ΔR is a function of N from equation (3) and so equilibrium participation, N^* , can be derived from equation (4).

We make the following assumptions about the distribution of ϵ :

ASSUMPTION 1: The left support is equal to zero. This assumption implies that average income must increase with migration, highlighting the trade off between moving and staying.

ASSUMPTION 2: The right support of the distribution is unbounded.

ASSUMPTION 3: The density of the distribution, f , is decreasing in ϵ . This assumption says that superior urban opportunities occur less frequently in the population.

Given these distributional assumptions,

LEMMA 1: Equilibrium participation is characterized by a unique fixed point, $N^ \in (0, P)$.*

$\Delta M = 0$ because $M_I = M_A$. $\Delta R > 0$ by assumption. This implies, from Assumption 1, that $F(\Delta M + \Delta R) > N/P$ at $N = 0$. Assumption 2 implies that $F(\Delta M + \Delta R) < N/P$ at $N = P$. $F(\Delta M + \Delta R)$ is increasing in N because R_I is decreasing in N (hence, ΔR must be increasing in N). By a continuity argument, a fixed point N^* at which equation (4) is satisfied must exist. We show in the Appendix that Assumption 3 implies that $F(\Delta M + \Delta R)$ is strictly concave, ensuring that this fixed point is unique.

D. Participation and Income Sharing with Inequality

We now characterize equilibrium participation and the income-sharing rule with heterogeneous rural incomes. By introducing this realistic feature of communities,

we are able to derive an important implication of our theory, which is that relatively wealthy households within the community will be more likely to send members to the city. To derive the new equilibrium, we take advantage of the fact that the ratio of marginal utilities between any two households participating in the network must be the same in all states of the world with full risk-sharing. Dividing the community into K income classes of equal size, P_k , this implies, given log preferences, that $C_{ks}/C_{Ks} = \lambda_k$, where C_{ks} , C_{Ks} denote the consumption of households in income class k and K (the highest income class) in state s of the world.

Aggregating over all households who choose to participate in the network— N_k in each income class k —each household in income class k consumes a fraction $\lambda_k / \sum_k \lambda_k N_k$ of the total income, $\sum_i y_{is}$, that is generated by the insurance network in each state of the world. Note that we normalize so that λ_K equals one. Following the same steps as in equations (2) and (3), expressions for the mean and variance of consumption with insurance in each income class k are derived as follows:

$$(5) \quad M_{Ik} = \left(\frac{\lambda_k}{\sum_k \lambda_k N_k} \right) \sum_k N_k M_{Ak} \quad V_{Ik} = \left(\frac{\lambda_k}{\sum_k \lambda_k N_k} \right)^2 \sum_k N_k V_{Ak}.$$

Because total income is pooled and then redistributed with full risk-sharing, consumption in each income class is now a function of the number of participants, N_k , and the income-sharing rule, λ_k , in every income class. However, equation (5) implies that the normalized risk, $R_I \equiv V_{Ik}/M_{Ik}^2$ is the same for all income classes and is independent of λ ,

$$(6) \quad R_I = \frac{\sum_k N_k V_{Ak}}{(\sum_k N_k M_{Ak})^2}.$$

Participation in the network continues to be derived as the solution to a fixed-point problem, but this problem must now be solved for each income class. Equilibrium participation will satisfy the following conditions, corresponding to equation (4), for each income class k :

$$(7) \quad \frac{N_k}{P_k} = F(\Delta M_k + \Delta R_k),$$

where $\Delta M_k \equiv \log(M_{Ik}) - \log(M_{Ak})$, $\Delta R_k \equiv 1/2\beta R_{Ak} - 1/2R_I$.

If we knew the income-sharing rule, λ_k , we could substitute expressions from equations (5) and (6) in equation (7) to solve simultaneously for N_k in all K income classes. The more challenging problem that we face is that the sharing rule λ_k and participation N_k must be derived simultaneously. To derive the sharing rule that is chosen by the community, we assume that its objective is to maximize the surplus that is generated by the insurance network. This surplus is the utility from participation in the network minus the utility in autarky, summed over all income classes. Within each income class, k , the total number of participants is determined by a

threshold $\epsilon_{Ik} = \Delta M_k + \Delta R_k$. Households with $\epsilon > \epsilon_{Ik}$ would send members to the city regardless of whether or not the insurance network was in place. They can thus be ignored when computing the surplus generated by the network. If $\beta < 1$, and given that $\epsilon > 0$, households with $\epsilon < \epsilon_{Ik}$ will always send members to the city when the network is absent. Total surplus can then be described by the expression,

$$W = \sum_k P_k \int_0^{\epsilon_{Ik}} \left\{ \left[\log(M_{Ik}) - \frac{1}{2} R_I \right] - \left[\log(M_{Ak}) - \frac{1}{2} \beta R_{Ak} + \epsilon \right] \right\} f(\epsilon) d\epsilon.$$

Noting that $N_k = P_k \int_0^{\epsilon_{Ik}} f(\epsilon) d\epsilon$ and collecting terms, the surplus function reduces to²³

$$(8) \quad W = \sum_k N_k \epsilon_{Ik} - P_k \int_0^{\epsilon_{Ik}} \epsilon f(\epsilon) d\epsilon.$$

Equilibrium participation and the income-sharing rule can be jointly derived by maximizing W with respect to λ_k , subject to the fixed point conditions in equations (7), after substituting in the expressions for M_{Ik} , R_I from equations (5) and (6). We now use this theoretical framework to identify which households benefit less (more) from the network and who should therefore be more (less) likely to have migrant members.

E. Relative Wealth, Rural Risk, and Migration

If participation in the network were fixed, the community could increase the surplus generated by the network by redistributing income from richer households to poorer households (given diminishing marginal utility). If households can select out of the network, however, the sharing rule must be attentive to the possibility that increased exit by households who subsidize the rest of the network will make it smaller, reducing its ability to smooth consumption. We nevertheless obtain the following result.

PROPOSITION 1: *Some redistribution is socially optimal, which implies that (relatively) wealthy households in the community should ceteris paribus be more likely to have migrant members.*

To derive this result in the Appendix, we consider the case with two income classes, $k \in \{L, H\}$, of equal size, $P_L = P_H$, where $M_{AH} > M_{AL}$. Recall that the threshold ϵ in each income class, $\epsilon_{Ik} = \Delta M_k + \Delta R_k$, and that the number of participants, $N_k = P_k F(\epsilon_{Ik})$. To ensure that differences in participation across income classes do not arise for other reasons, we assume that $R_{AL} = R_{AH}$, which implies that $\Delta R_L = \Delta R_H$, and that the ϵ distribution, characterized by the F function, is the same for both income classes. Without income redistribution, mean consumption equals mean income for each household and so $\Delta M_L = \Delta M_H = 0$. It follows that

²³If $\beta > 1$, then there exists a threshold ϵ , $0 < \epsilon_{Ak} < \epsilon_{Ik}$, below which households do not send migrants to the city even when the network is absent. The second term in square brackets in the preceding equation is then replaced by $\int_0^{\epsilon_{Ak}} \log(M_{Ak}) - 1/2\beta R_{Ak} + \int_{\epsilon_{Ak}}^{\epsilon_{Ik}} \log(M_{Ak}) - 1/2\beta R_{Ak} + \epsilon$. We would then integrate from ϵ_{Ak} to ϵ_{Ik} , rather than from zero to ϵ_{Ik} , in equation (8). This would not, however, change any of the results that follow.

participation and, hence, migration rates will be the same in both income classes without redistribution.

Denote the ratio of consumption between low-income and high-income households in each state of the world by λ . Without income redistribution, λ is the ratio of mean-incomes of the two classes, M_{AL}/M_{AH} . With equal income sharing, λ is equal to one. In general, $\lambda \in [M_{AL}/M_{AH}, 1]$. The sharing rule λ^* that is chosen in equilibrium cannot be derived analytically. What we do instead is to focus on the (only) income-sharing rule without redistribution, $\lambda = M_{AL}/M_{AH}$. We show that an increase in λ , evaluated at that sharing rule, unambiguously increases the surplus, even after accounting for the effect on participation. This implies that there must be some redistribution in equilibrium. Migration rates do not vary across income classes in the absence of redistribution, by construction. With redistribution, relatively wealthy households benefit less from the network and so are more likely to have migrant members.

The theory also has implications for how variation in rural income risk affects migration and redistribution within the network. The decision rule specified in equation (1) indicates that the gain from network insurance, $\beta R_A - R_I$, is larger for a household facing greater rural income risk, R_A . This implies that the threshold ϵ_I , above which it will send members to the city is larger, and so it is more likely to participate in the network. However, we must once again account for potential redistribution and its consequences for participation. In this case, redistribution will favor safe households at the expense of households facing greater income risk. We are nevertheless able to derive the following general result.

PROPOSITION 2: *Households that face greater rural income risk are ceteris paribus less likely to have migrant members.*

This result is derived in the Appendix. Income classes, $k \in \{L, H\}$, are now replaced by risk-classes, $k \in \{R, S\}$, where $R_{AR} > R_{AS}$. To rule out redistribution for other reasons, mean rural incomes are assumed to be the same in both risk-classes, $M_{AR} = M_{AS}$. The ϵ distribution is also assumed to be the same in both classes. Relabel λ to be the ratio of consumption between high-risk and low-risk households in each state of the world. Without redistribution, $\lambda = M_{AR}/M_{AS} = 1$. If the two risk-classes are of equal size, $P_R = P_S$, then the number of network participants will be greater in the risky class, $N_R > N_S$, because $\Delta M_R = \Delta M_S = 0$, $\Delta R_R > \Delta R_S$. The benefit of redistribution is that a dollar taken from each participating risky household will be divided among a smaller number of safe households. At the same time, the number of households that benefit is smaller than the number who lose and this will be accounted for when computing the surplus. The effect of redistribution on overall participation, with its consequences for consumption smoothing, must also be considered.

If there are net gains from redistribution, nevertheless, then λ will decline. However, since the gains from redistribution arise because $N_R > N_S$, λ must be bounded below at a level $\underline{\lambda}$ at which participation is the same in both risk classes; $\lambda \in [\underline{\lambda}, 1]$. To prove Proposition 2 we focus on the (only) income-sharing rule with equal participation, $\lambda = \underline{\lambda}$, and show that an increase in λ evaluated at $\underline{\lambda}$, unambiguously increases the surplus. This implies that $\lambda^* > \underline{\lambda}$ and, hence, that

households facing greater rural income risk have higher participation rates (lower migration rates) in equilibrium even with redistribution. In contrast, if networks are absent and we maintain the standard risk-diversification assumption, $\beta < 1$, then households facing greater rural income risk will be more likely to send migrants to the city.²⁴

III. Testing the Theory

The theory generates three testable predictions: (i) income is redistributed in favor of poor households within the caste; (ii) relatively wealthy households who, therefore, benefit less from the insurance network should be more likely to have migrant members; and (iii) households facing greater rural income risk who benefit more from the network should be less likely to have migrant members. These tests shed light on the central hypothesis that insurance provided by rural networks inhibits mobility. Additional tests validate the key assumption that permanent male migration is associated with a loss in network services. These results, taken together, can be used to distinguish between our explanation for large wage gaps and low migration in India and alternative explanations that do not require a role for rural insurance networks.

One explanation for low migration and large wage gaps in India is based on the existence of urban caste-based labor market networks. While the members of a relatively small number of castes with well-established urban networks will enjoy high wages in the city, most potential migrants moving independently will be shut out of the urban labor market. Past research, e.g., Munshi and Rosenzweig (2006) and Munshi (2011), indicates that caste networks continue to be active in Indian cities. However, this does not preclude the coexistence of our theory, in which the loss in rural insurance reduces individual migration, with this alternative explanation in which migrants must move as a group, which results in lower overall mobility. Two distinguishing features of our theory are (i) that households facing greater rural income risk are less likely to have migrant members, and (ii) that migration is associated with a loss in network services.²⁵

There also can be alternative explanations for the first two predictions of our theory, but no alternative that we are aware of delivers all three predictions. For example, it is possible that communities provide other types of public goods financed by a progressive payment scheme, also resulting in redistribution and increased exit by relatively wealthy households.²⁶ Moreover, there may be other reasons why higher

²⁴The available evidence supports the assumption that $\beta < 1$. As noted in Section I, the NSS data indicate that there is lower unemployment in urban versus rural areas in India. Everything else equal, this implies that income risk declines with migration. β will certainly be less than one in that case, and this is what we obtain when we estimate the structural parameters of the model. Without rural insurance networks, a household will not send migrants to the city if $\log(M_A) - \frac{1}{2}R_A \geq \log(M_A) - \frac{1}{2}\beta R_A + \epsilon$. If $\beta < 1$, and we continue to assume $\epsilon > 0$, then all households will send migrants, which is inconsistent with Proposition 2. Once we introduce a migration cost, K , there exists a threshold $\epsilon^* = K - \frac{1}{2}(1 - \beta)R_A$, above which households send migrants. ϵ^* is decreasing in R_A , establishing a positive relationship between rural income risk and migration in the absence of rural insurance networks that runs counter to Proposition 2 once again.

²⁵Another explanation for low mobility, as in the literature on kin-tax in Africa, e.g., Platteau (2000), is that origin networks tax migrants heavily. However, this would not explain why greater rural income risk is associated with lower migration.

²⁶Recent evidence from developing countries suggests that while payment schemes in rural communities are indeed redistributive, they are regressive rather than progressive (Olken and Singhal 2011). The empirical evidence thus runs counter to this alternative explanation.

household income, which is positively correlated with relative income within the community, will be associated with higher out-migration. Neither of these theories would explain why households facing greater rural income risk are less likely to have migrant members.

A. Evidence on Redistribution within Castes

We first empirically assess the extent of redistribution within castes. We begin with data from the 2005–2011 Indian ICRISAT panel survey, which provides information on household incomes over a seven-year period and consistent consumption data for the first four of those years, for a sample of households in six villages in the states of Andhra Pradesh and Maharashtra. The panel data enables us to compute the theoretically-relevant intertemporal mean values for consumption and income for each household.²⁷

We divide up the households in each caste into quintiles of the within-caste income distribution to compute mean consumption and mean income in each income class. Restricting the sample to castes with at least 20 members represented in the data, we have seven castes among 552 households in the six villages. Table 5, column 1, reports relative income, measured by the ratio of average income in the income class to average income in the highest income class, averaged across all castes. Relative income is increasing across income classes by construction. Column 2 reports the corresponding statistics for relative consumption. A comparison of column 1 and column 2 indicates that there is substantial redistribution within castes. The consumption ratio exceeds the income ratio for each income class, with the consumption-income ratio in column 3, or more correctly the ratio of ratios, close to four for the lowest income class.

With just 500 households and 7 castes, the ICRISAT sample is too small to test the second prediction of the model, which is that relatively wealthy households within their caste should be more likely to have migrant members. The 2006 Rural Economic Development Survey (REDS), collected information from over 119,000 households residing in 242 villages in 17 major Indian states on the migrant status of each household; i.e., whether any adult male (father, brother, or son of the head) had permanently left the village in the preceding five years, the income of each household in the prior year, and its subcaste affiliation.²⁸ In the data, permanent migrants are defined as those who are no longer members of the local household.²⁹

²⁷These data provide the value of all foods and nonfoods consumed, including self-produced and purchased items measured at various times over the year, that can be summed to obtain an annual total. The Indian CPI for agricultural laborers is used to compute real consumption values expressed in 2005 rupees. Average real (inflation-adjusted) annual income is also computed for the same households over the entire seven-year period, including wages, salaries, and farm and nonfarm income, but excluding any transfers and remittances.

²⁸The selection of villages was meant to provide a representative sample of rural Indian households. Any sample of villages will not yield a representative sample of castes, unless castes are distributed evenly across villages. For the castes represented in the data, however, the income distribution derived from the randomly sampled villages will be representative of the caste-level income distribution.

²⁹We cannot determine whether any of the departed household members were formerly the household head. There are few instances of entire households migrating in India—data from the 1999 and 2006 REDS indicates that less than 10 percent of rural households that were present in the 1999 round could not be located in the same village in 2006. The comparable statistic for Indonesian households from the first two waves of the IFLS is 18 percent over just four years (Thomas, Frankenberg, and Smith 2001). The Indonesian data also disentangle migration,

TABLE 5—INCOME AND CONSUMPTION WITHIN THE CASTE

	ICRISAT			REDS 2006			
	Relative income	Relative consumption	Consumption-income ratio	Relative income	Relative consumption	Consumption-income ratio	Migration
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>Relative income class</i>							
1	0.119	0.460	3.871	0.316	0.843	2.665	0.032
2	0.281	0.625	2.224	0.416	0.854	2.052	0.034
3	0.373	0.626	1.680	0.513	0.871	1.697	0.051
4	0.510	0.673	1.319	0.627	0.887	1.413	0.046
5	1.000	1.000	1.000	1.000	1.000	1.000	0.051

Notes: Income classes are defined by quintiles within each caste. Income and consumption are measured relative to the highest (fifth) income class. REDS 2006 income and consumption are inputted from ICRISAT data. REDS data consists of 100 castes, while ICRISAT data consist of 7 castes. Sample-size restriction is at least 30 households per caste with REDS data and 20 households per caste with ICRISAT data.

Nonresident household members who are temporary migrants are included in the household roster, as is standard in most household surveys.

To test how relative income affects migration, we construct a measure of the household's average income over time. This will depend on its wealth (productive assets) as well as the number of earners. A shortcoming of the REDS data is that it provides incomes only in the year preceding the survey and includes transfers. To address this limitation, we impute average income for each household using the ICRISAT panel dataset and a vector of household and village-level variables that are common to both the REDS and ICRISAT datasets. Both datasets provide household-level information on total land area (together with a binary variable indicating whether the household is landless), irrigated area, soil type (red, black, sandy) and soil depth, household size, the number of earners, and the occupation of the household head. Each dataset also provides, at the village level, a time-series of rainfall; daily rainfall for all seven years for the ICRISAT survey and monthly data over an eight-year period starting in 1999 for the REDS, from which we construct village-level mean annual rainfall and the variance of annual rainfall.

The land characteristics, taken together, determine the value of land owned by the household. Although land wealth accounts for 85 percent of household wealth in rural India (Rosenzweig and Wolpin 1993), land sales are extremely infrequent (Foster and Rosenzweig 2002). This implies that land wealth is largely inherited and can be treated as predetermined, at least from the perspective of current household members. The household's permanent labor income is determined by the number of earners and the occupation of the household head.³⁰ When imputing average income for REDS households with permanent male migrants, we included those migrants among the earners.

We first estimate, using ICRISAT data, the relationship between average annual household income over all seven years excluding all transfers and the vector of

which accounts for two-thirds of the missing households, from attrition. Assuming that two-thirds of missing Indian households also migrated, this implies that the annual rate of permanent household migration is less than 1 percent.

³⁰Occupational choices are very limited in rural India. In the 2006 REDS census, 34 percent of the heads were cultivators, i.e., landowners, and 38 percent were agricultural laborers and laborers not classified. The next most popular occupation—shopkeepers—accounted for just 2 percent of the household heads. Not surprisingly, the most important determinants of household income are landholding size and number of male earners.

(common) household and village characteristics, including all the household characteristics interacted with mean rainfall. The vector of regressors have sufficient predictive power, given that this is a cross-sectional regression, with an R^2 around 0.3. The coefficients obtained from the regression estimated with ICRISAT data are then used to impute average income for each of the REDS households, based on their characteristics. As noted, the ICRISAT villages are located in Andhra Pradesh and Maharashtra. For comparability, we restricted the REDS sample to four geographically contiguous and broadly similar South Indian states—Maharashtra, Andhra Pradesh, Tamil Nadu, and Karnataka—when imputing incomes.³¹ Average consumption is also imputed for the REDS households from ICRISAT consumption data using the same method and the same first-stage specification that was used to impute average income.

Table 5, columns 4–6, replicate the computations carried out with the ICRISAT data for the REDS sample. A comparison of column 4 and column 5 indicates that, as is true for the ICRISAT households (where no values are imputed), there is substantial redistribution within castes. The consumption ratio exceeds the income ratio for each income class, with the consumption-income ratio in column 6, or more correctly the ratio of ratios, close to three for the lowest income class.³² Finally, column 7 reports migration rates by income class. Consistent with the theory and the redistribution documented in Table 5, we see that migration rates are increasing in relative income. A household's *relative* position within its caste's income distribution will be positively correlated with its *absolute* income. The statistics reported in column 7 do not account for the direct effect of (absolute) household income on migration, which would dampen the increase in migration across relative income classes if wealthier households are less likely to migrate for other reasons. The regression analysis that follows will control for the direct effect of the household's income on migration.

B. *Reduced-Form Estimates*

Proposition 1 links the household's *relative* income to its decision to send migrants to the city. However, household income could directly determine migration due to credit constraints or risk-aversion (Bryan, Chowdhury, and Mobarak 2014). Higher income households also have a larger number of male earners (on average) and this increases the probability that any male will migrate. We thus include the household's own income, as an independent determinant of migration, when testing Proposition 1. Once this variable is included, the relative income effect is captured by including average caste income as an additional regressor. Conditional

³¹The advantage of including Karnataka and Tamil Nadu is that they increase the number of observations, particularly in the structural estimation where households are aggregated by income class in each caste. The reduced-form results reported below are largely unchanged and statistically significant when the sample is restricted to the non-ICRISAT states (Karnataka and Tamil Nadu).

³²Although we follow standard practice and ignore savings in our model of mutual insurance, an alternative explanation for the observed pattern of consumption is that wealthy households save a greater share of their income. This is unlikely to be a viable explanation, however, because private savings are extremely low in rural India (as documented, for example, by Breza and Chandrasekhar 2015). Consistent with a negligible role for private savings, the structural model, which ignores savings, will generate estimates of relative consumption that are almost identical to the consumption ratios reported in Table 5.

on the household's own income, an increase in average caste income implies that it is relatively less wealthy within its caste. To test Proposition 1 we thus estimate a regression of the form

$$(9) \quad M_i = \pi_0 + \pi_1 y_i + \pi_2 \bar{y}_i + \epsilon_i,$$

where M_i indicates whether any male member of household i had moved permanently from the village, y_i is the household's average income over time, and \bar{y} is the corresponding average statistic for its caste, which is constructed by averaging incomes over all households in the caste. As discussed above, this information is available from the 2006 REDS census.

Conditional on average caste income, an increase in a household's income implies that it is relatively wealthy and, therefore, should be more likely to have migrant members. However, household income could directly determine migration, as discussed above, and so the π_1 coefficient cannot be used to test the theory. The key test of Proposition 1 is $\pi_2 < 0$; conditional on the household's own income, an increase in caste income implies it is relatively less wealthy and, therefore, should be less likely to have migrant members.

Table 6, column 1 reports the estimates of equation (9). Coefficient standard errors are bootstrapped to account for the use of imputed incomes. As predicted by Proposition 1, the estimated coefficient on caste income, $\hat{\pi}_2$, is negative and significant. This result provides support for the theory in which the migration decision is made in the context of a caste network, and networks redistribute income in favor of the poor. The positive and significant coefficient on own household income, $\hat{\pi}_1$ in column 1, is also consistent with the theory but, as noted, there are other interpretations.³³

Proposition 2 indicates that households who face greater rural income risk should be less likely to have migrant members. We test this prediction by including the rural income risk faced by the household as an additional regressor in Table 6, column 2. Income risk in our theory is measured by the coefficient of variation of the household's income, squared. We construct the variance of the household's income over time using the same method that was used to impute average income.³⁴ Using the constructed variance to compute income risk, we see in Table 6, column 2 that households facing higher rural income risk are indeed less likely to have migrant members. While this result is consistent with our theory in which migration results in the loss of risk-reducing network services, it is inconsistent with standard models of individual migration in which adverse origin characteristics lead to higher out-migration rates.

³³Our analysis with 2006 REDS data restricts the sample to castes with at least 30 households in the census. This ensures that there will be a sufficient number of households in each income class in the structural estimation, where castes are divided into four to six income classes. The reduced-form results in Table 6 are robust to restricting the sample to castes with at least 10 households in the 2006 REDS census.

³⁴The specification in the first step, using ICRISAT data, is the same, except that the household characteristics are interacted with the variance of village rainfall in the ICRISAT villages. The set of household and village level regressors once again have sufficient power, with an R^2 around 0.3. The estimated coefficients from the first step are subsequently used to predict the variance of income for each of the REDS households using their characteristics and village-level rainfall variances. We estimate the relationship between log variance and the household and village characteristics in the first step. Predicted log variance for the REDS households can then be transformed to the variance of income, ensuring that no negative values are obtained.

TABLE 6—REDUCED-FORM MIGRATION ESTIMATES

	Migration					
	(1)	(2)	(3)	(4)	(5)	(6)
Household income	0.0059 (0.0024)	0.0051 (0.0024)	0.0026 (0.0033)	0.0025 (0.0033)	0.0021 (0.0030)	0.0021 (0.0033)
Caste income	-0.016 (0.0043)	-0.018 (0.0055)	-0.022 (0.008)	-0.024 (0.0107)	-0.025 (0.0107)	-0.017 (0.014)
Income risk	—	-0.00038 (0.00015)	-0.00037 (0.00016)	-0.00053 (0.00017)	-0.00053 (0.00017)	-0.00053 (0.00011)
Village income			0.007 (0.011)	0.006 (0.013)	— —	— —
Village/caste income					0.0073 (0.013)	0.0088 (0.027)
Village fixed effects	No	No	No	No	No	Yes
Infrastructure variables	No	No	No	Yes	Yes	No
<i>Joint sig. of infrastructure variables</i>						
χ^2	— —	— —	— —	16.14 [0.0011]	16.59 [0.00090]	— —
Observations	19,362	19,362	19,362	19,362	19,362	19,362

Notes: Bootstrapped standard errors in parentheses are clustered at the caste level in columns 1, 2, and 6 and two-way clustered at the caste and village level in columns 3–5. Income measured in lakhs of rupees, (1 lakh = 100,000). Infrastructure variables: whether there is a bank, secondary school, health center, or bus station in the village, as well as distance to the nearest town. χ^2 *p*-value reported in square brackets. Sample-size restricted to castes with at least 30 households.

Source: REDS census 2006

Recall that the relationship between relative income and migration in Proposition 1 was derived conditional on rural income risk. The relationship between income risk and migration in Proposition 2 was derived conditional on household income (and the household's position in the caste income distribution). The specification in column 2 allows us to (simultaneously) estimate these conditional effects, as required by the theory. The point estimates indicate that the magnitude of these effects are large. A one standard deviation decrease in the risk measure doubles the migration rate (from a baseline of 3.1 percent). A one standard deviation increase in own income increases the migration rate by 10 percent, while the same increase in caste income reduces the migration rate by 30 percent.

The theory does not specify what constitutes the domain of the network. Although the organization of Indian society, with individuals marrying strictly within their caste, leads us to posit that rural insurance networks are organized around the caste, they could potentially be organized at the level of the village, as assumed in previous studies on risk-sharing in India, e.g., Townsend 1994; Ligon 1998. To address this possibility, we include mean village income as an additional regressor in Table 6, column 3. Consistent with Mazzocco and Saini (2012) who report full risk-sharing at the caste level, but reject full risk-sharing in the village with ICRISAT data, we see that the coefficient on mean caste-income is stable and remains highly significant, whereas the corresponding coefficient on village income is small and imprecisely estimated. One remaining possibility is that the estimated village-income coefficient is biased because village income is correlated with village infrastructure, which directly determines migration. To address this possibility, we include

variables indicating whether a bank, secondary school, health center, or bus station is located in the village, as well as the distance to the nearest town, in Table 6, column 4. Although the infrastructure variables are jointly highly significant, the remaining coefficient values are largely unchanged.

We believe that the caste is exceptionally effective at consumption smoothing because of its large size and scope (extending over many villages). While the preceding results indicate that insurance networks in India are organized around the caste, they do not tell us whether the network extends beyond village boundaries. To answer this question, we replace village income with the mean income of caste households *within* the village in Table 6, column 5. Reassuringly, the complete caste-income coefficient maintains its size and significance, while the restrictive caste-income measure has little effect on migration and is statistically insignificant. The stability of the caste-income coefficient to the inclusion of a vector of village-level variables indicates that the results are not being driven by unobserved village-level effects.³⁵ Nevertheless, as a final robustness test, we include a full set of village fixed effects in column 6. Although the caste-income coefficient is now only statistically significant at the 12 percent level (one-tailed test), it remains as large (in absolute magnitude) as it was with the benchmark specification in column 2. Results from a Hausman test (available from the authors) indicate that the difference in the caste income coefficient between the benchmark specification and the specification with village fixed effects is not significant at the 5 percent level. The difference between the household income coefficients is also not statistically significant, at any level. The only statistically significant change is for the income-risk coefficient, but that coefficient is highly significant in both specifications (and gets more negative in the fixed effects specification). Overall, our results provide strong support for the hypothesized trade-off between the insurance provided by rural caste-based networks and the income gain from migration. Those households that benefit less (more) from the rural network are more (less) likely to have migrant members.

C. Structural Estimates

Having found evidence consistent with the theory, we now estimate the structural parameters of the model. The structural estimates are used to (i) externally validate the model; (ii) to provide independent support for the redistribution within castes that is predicted by the theory; and (iii) to carry out counterfactual simulations that compare the sensitivity of migration to the rural-urban wage gap and formal insurance. We also conduct counterfactual policy simulations that quantify the mobility-enhancing effects of a government safety net for poor households and a credit scheme benefiting wealthy households.

There are two exogenous variables in the model, measured at the level of the income class, k : mean-income, M_{Ak} , and normalized risk, $R_{Ak} \equiv V_{Ak}/M_{Ak}^2$. While there is a single caste (community) in the theoretical analysis, there are 100 castes in the 2006 REDS census, which we use for the structural estimation. To

³⁵ Standard errors in columns 3–5 are two-way clustered at the level of the caste and the village.

be consistent with the model, we thus proceed to average mean-income, M_{Ai} , and the variance of income, V_{Ai} , which were previously imputed for each household i , across all household in each income class, k , within each caste, j , to obtain M_{Akj} , R_{Akj} . The specifications that we report partition the households within each caste into income quintiles, but the results are robust to using four or six income classes.

To understand how the model is estimated, suppose, to begin with, that the β parameter and the F function are known. For a given income-sharing rule in caste j , described by the λ_{kj} vector, we can then solve for participation in each income class, N_{kj} , from the fixed-point conditions, equation (7), after substituting in the expressions for mean-consumption with insurance $M_{I_{kj}}$ and consumption risk with insurance $R_{I_{kj}}$ from equations (5) and (6). The total surplus generated by the insurance arrangement can then be computed in caste j from equation (8). Searching over λ_{kj} , the income-sharing rule that is ultimately selected in caste j will maximize the total surplus. If the model is correctly specified, predicted migration (which is one minus the participation rate) will match actual migration at that sharing rule.

Now suppose that β is unknown and must be estimated, but continue to assume that the F function is known. For an arbitrary β we can solve for the surplus-maximizing λ_{kj} , as described above. However, predicted migration will no longer match actual migration. To estimate β , we exploit the fact that migration in each income class in each caste will decline as this parameter increases. There thus exists a unique β for which overall predicted migration, across all income classes in all castes, matches actual migration. This will be our best estimate of β . We match on overall migration to estimate β because a unique solution is assured and because this moment will be the outcome of interest in the counterfactual simulation that follows.

The estimation procedure described above was based on the assumption that the F function, which characterizes the distribution of income gains from migration, ϵ , was known. We now proceed to describe how this function is derived. Recall that we made three assumptions about the ϵ distribution in the model: (i) the left support is equal to zero; (ii) the right support is unbounded; and (iii) the density of the distribution is declining in ϵ . The exponential distribution satisfies each of these assumptions and so we assume that ϵ is distributed exponentially. An additional advantage of the exponential distribution is that it is characterized by a single parameter, which we denote by ν ; $F(\epsilon) = 1 - e^{-\nu\epsilon}$, where $E(\epsilon) = 1/\nu$.

The distributional parameter, ν , is estimated in two steps. We first use REDS and NSS data to compute the average income gain from migration for households with permanent male migrant members in the 2006 REDS census. The household's land value, the number of working-age adults, and the education of the household head (which we assume applies to all working members) is available from the REDS census. Urban and rural wages, by education category, are available from the NSS. These data sources can be combined to compute the income gain from migration, $\tilde{\epsilon}$, and its utility equivalent, $\epsilon = \log(1 + \tilde{\epsilon})$.³⁶ We assume that this derived income

³⁶We compute average rural income as 5 percent of the household's total asset holdings at the beginning of the reference period (one year before the survey round) plus labor income, based on the assumption that the adults in the household work for 312 days in the year. Let V be the household's asset value, L the number of adults, and W_U , W_R the education-adjusted urban and rural wages from the NSS. The average number of working-age adults per household in the 2006 REDS census is 1.4, and so it is reasonable to assume that a single individual

gain, ϵ , is the representative (median) value for households with migrants. For example, if 4 percent of households have permanent migrant members, then ϵ applies to a household at the ninety-eighth percentile of the ϵ distribution. This assumption, together with the properties of the exponential distribution, can be used to derive the distributional parameter:

$$(10) \quad \nu = \frac{-\log(x/200)}{\epsilon},$$

where x equals four in the preceding example. Once ν is computed, the risk-parameter, β , can be estimated as described above.

As a basis for comparison with the estimates that follow, Table 7A, columns 1 and 2, list relative consumption and migration in each of the five relative income classes (averaged across all castes). Columns 3 and 4 report the β estimate, predicted relative consumption, λ_k , and predicted migration, $1 - N_k/P_k$, in each of those classes, k (once again averaged across all castes, j). Jackknifed standard errors, constructed by removing one caste at a time and reestimating the model, are reported in parentheses.³⁷ The point-estimate for the β parameter is 1.4. Similar results are obtained with four and six income classes in Appendix Table A3. $\beta < 1$ if migration reduces income risk, as commonly assumed. We will see momentarily that β does decline below one, and is much more precisely estimated, with a flexible specification of the model that does a better job of matching the data.

The β parameter is estimated by matching on overall migration, which is 4.3 percent. Notice, however, that migration rates predicted by the model are lower (higher) than actual migration rates in low (high) income classes. In contrast, relative consumption levels (λ s) predicted by the model match closely with actual relative consumption. We cannot reject that predicted and actual relative consumption are statistically equal, at conventional levels, in each income class and across all specifications in Table 7A. The match in terms of magnitudes is very close (less than 5 percent error for each income class).³⁸ This close match in predicted and actual λ s is effectively a test of external validity, given that consumption data are not used to estimate the model. It provides empirical support for both the (logarithmic) distributional assumption and our use of a utilitarian social welfare function, placing equal weight on all income classes in equation (8). The substantial redistribution documented in Table 5 appears to be driven entirely by an attempt to equate marginal utilities across income classes.

migrates. The income gain from migration in that case is $\frac{312(W_U - W_R)}{0.05V + 312W_R L}$. This statistic is averaged across all households with migrants.

³⁷The jackknifed standard error for any parameter θ is given by the expression $[n/n - 1 \sum_i (\theta_i - \bar{\theta})^2]^{1/2}$, where n is the number of times the model is reestimated, θ_i is the parameter estimate when it is reestimated for the i^{th} time, and $\bar{\theta}$ is the average across all θ_i . In our case, with 100 castes, $n = 100$.

³⁸An alternative approach to compare the closeness of the match, suggested by a referee, would be to compute the probability that the predicted and actual λ s are as close as they are by random chance. We implement this test for the lowest relative income class, since it will certainly not subsidize any other income class and so the range of feasible λ s is known. Given that the relative income for the lowest income class is 0.31 from Table 5, λ must lie in $[0.31, 1]$. If the predicted λ is distributed uniformly over this range, the probability that it will lie within 0.04 of the actual λ , as observed in Table 7A, is 12 percent. The standard error for the lowest income class is also small enough for us to reject that there is no redistribution (λ is equal to 0.31) with greater than 95 percent confidence.

TABLE 7A—STRUCTURAL ESTIMATES (*Single ν Parameter*)

	Actual		Predicted	
	Relative consumption (1)	Migration (2)	Relative consumption (3)	Migration (4)
<i>Relative income class</i>				
1	0.843	0.032	0.801 (0.071)	0.000 (0.00020)
2	0.854	0.034	0.817 (0.070)	0.014 (0.0073)
3	0.871	0.051	0.834 (0.063)	0.039 (0.0083)
4	0.887	0.046	0.868 (0.044)	0.060 (0.0089)
5	1.000	0.051	1.000	0.100 (0.014)
Overall		0.043		0.043
β			1.410 (0.91)	

Notes: Structural estimation is based on 100 castes with at least 30 households in the census. Relative income classes are defined by quintiles within each caste. Consumption is measured relative to the highest (fifth) income class in columns 1 and 3. Relative consumption and migration are computed as the average in each income class across all castes. Jackknifed standard errors in parentheses.

Source: REDS census 2006

One possible explanation for the mismatch between predicted and actual migration across income classes is that the income gain from migration varies with rural income, perhaps because members of wealthier households are better educated. Although the theoretical analysis assumes that the (proportional) income gain is independent of rural income, we relax this assumption when estimating the model in Table 7B, columns 3 and 4. The households in the 2006 REDS census are divided into five *absolute* income classes, without regard to their caste affiliation, and the ν parameter is then computed in each income class using the procedure described above.³⁹ Each *relative* income class k in caste j can be mapped into an *absolute* income class based on its mean income, $M_{A_{kj}}$. The ν parameter computed for that absolute income class is then assigned to class k in caste j . The results with this flexible specification in Table 7B, columns 3 and 4, are very similar to the results obtained with the benchmark specification, with a single ν , in Table 7A, columns 3 and 4. Allowing the income gain from migration to vary with rural income evidently does not reduce the mismatch between predicted and actual migration.

Although the theory focuses on the constraints faced by individuals migrating independently, the empirical analysis must take account of the small number of castes that have established urban networks in the city. While most castes report

³⁹ Appendix Table A3 shows, with five absolute income classes, that land value, the number of working-age adults, and the education of the household head are all generally increasing across income classes as expected. Although the NSS data indicate that the rural-urban wage gap is increasing with education (Hnatkovska and Lahiri 2014), wealthier households are starting from a higher rural income base and so the proportional gain in income will naturally be smaller for them. There is no obvious relationship between absolute income and the income gain from migration, which is reflected in the migration rate, in Table A3.

TABLE 7B—STRUCTURAL ESTIMATES (ν by Income Class and Caste)

	Actual		Predicted (ν varies by income class)		Predicted (ν varies by caste)			
	Relative consumption (1)	Migration (2)	Relative consumption (3)	Migration (4)	Relative consumption (5)	Migration (6)	Relative consumption (7)	Migration (8)
<i>Relative income class</i>								
1	0.843	0.032	0.794 (0.058)	0.000 (0.00018)	0.751 (0.097)	0.000 (0.000081)	0.730 (0.083)	0.032 (0.0095)
2	0.854	0.034	0.810 (0.059)	0.014 (0.0075)	0.767 (0.092)	0.011 (0.010)	0.744 (0.064)	0.032 (0.052)
3	0.871	0.051	0.827 (0.053)	0.039 (0.0088)	0.792 (0.070)	0.029 (0.025)	0.765 (0.055)	0.046 (0.027)
4	0.887	0.046	0.863 (0.038)	0.060 (0.0095)	0.842 (0.044)	0.055 (0.033)	0.825 (0.037)	0.044 (0.013)
5	1.000	0.051	1.000	0.101 (0.019)	1.000	0.119 (0.062)	1.000	0.051 (0.0074)
Overall		0.043		0.043		0.043		0.041
β			1.218 (0.78)		0.845 (0.92)		0.991 (0.18)	
α			—		—		0.012 (0.050)	
γ			—		—		4.45 (0.91)	

Notes: Structural estimation is based on 100 castes with at least 30 households in the census. Relative income classes are defined by quintiles within each caste. Consumption is measured relative to the highest (fifth) income class in columns 1, 3, 5, and 7. Relative consumption and migration are computed as the average in each income class across all castes. Jackknifed standard errors in parentheses. We match on two moments—migration in the lowest and the highest relative income class—in the flexible specification (columns 7 and 8). Standard errors for those two income classes in column 8 thus reflect sampling error due to the jackknife procedure. $\beta = \alpha + \gamma M_{ak}$ in columns 7 and 8. The estimated α and γ parameters are used to compute β for the representative household with mean-wealth in the REDS census.

Source: REDS census 2006

low migration rates in our data, a few do have substantial rates of permanent male migration. We account for these caste-level differences by estimating a flexible specification that allows for a caste-specific ν parameter. The income gain computed for households with migrants ϵ , using REDS and NSS data as described above, does not vary substantially across castes. The important difference is the point in the ϵ distribution where these households are located. For a caste with a strong network and many migrants, x in equation (10) is relatively large, which implies that $\nu = 1/E(\epsilon)$ will be small. A strong urban caste network, reflected in a high level of permanent migration in the data, thus effectively increases the income gain from migration. Despite the flexibility that is introduced with 100 ν parameters, the pattern of λ s and migration in Table 7B, columns 5 and 6, is very similar to the benchmark estimates obtained with a single ν parameter. Notice that the β coefficient is now smaller than one, consistent with the standard assumption that migration reduces income risk, although it continues to be imprecisely estimated.

We assumed, when deriving the theoretical results, that there was a single ν and a single β . The specifications reported in columns 3–6 relax the first assumption. We now relax the second assumption by allowing β to vary with rural household

income (there is no reason to allow β to vary by caste if households with migrants are excluded from their rural insurance networks). Recall that β reflects the income diversification that accompanies migration as well as access to nonnetwork consumption-smoothing mechanisms. It is not obvious how the reduction in income risk through migration varies with rural income. While we would expect wealthier households to have greater access to private credit, the effect on β is theoretically ambiguous. Consider, for example, two households trying to smooth their consumption with credit, where one household has twice as much income as the other in each state of the world. To smooth consumption completely, the wealthier household must receive twice as much credit in each state of the world. If land, the chief source of wealth in rural India, is not fully collateralized, then this condition may or may not be satisfied and the β -income relationship will be ambiguous.

The structural estimates in Table 7B, columns 7 and 8, allow β to be a linear function of average income in each relative income class within each caste; $\beta = \alpha + \gamma M_{Akj}$, without any theoretical guidance about the sign of the γ parameter. Empirically, predicted migration rates were too high (low) for higher (lower) relative income classes with a single β parameter. With the flexible specification, we thus expect the gamma coefficient to be positive to bring predicted migration in line with actual migration. The estimated coefficients in columns 7 and 8 are consistent with this prediction. Since there are two parameters to estimate— α and γ —we match on two moments: overall migration in the lowest and the highest relative income class.⁴⁰ It is reassuring to observe that predicted and actual migration now match closely in the remaining three income classes. We are even able to generate the non-monotonicity from relative income classes three to five, despite the fact that β is specified to be increasing linearly in mean-income M_{Akj} , without regard to the relative position of group k in caste j .

To compare the parameter estimates with the flexible specification with what we obtained earlier in Table 7A and Table 7B, columns 3–6, we use the estimated α and γ parameters to compute β for the representative household with mean-wealth in the REDS census. Jackknifed standard errors for the β parameter are constructed using the procedure described above. The β parameter with the preferred flexible specification now lies below one and is very precisely estimated. If we take a different approach that exploits variation in household income in the sample, then the β parameter would be less than one (based on the estimated α and γ parameters) for 86 percent of the households in the REDS census. This result is consistent with evidence from the NSS indicating that employment-risk is lower in urban than in rural areas, which implies (everything else equal) that income risk will be lower in urban areas. The fact that households facing greater rural income risk were nevertheless less likely to have migrant members in Table 6, indicates that the loss in network insurance must be large enough to dominate the substantial income gain and the (possible) income risk-reduction associated with migration.

Given that the structural model appears to fit the data reasonably well, we use the estimated parameter values to perform counterfactual simulations that quantify the magnitude of the mobility restrictions we have uncovered. There are two structural

⁴⁰The standard errors for these two income classes in column 10 simply reflect the sampling error that is generated when we reestimate the model, removing one caste at a time, in the jackknife procedure.

parameters in the model: the risk parameter, β , and the income-gain parameter, ν . We reduce the β parameter to assess the effect of an improvement in formal insurance on migration, taking as given the rural insurance arrangements that are in place. We assess the sensitivity of migration to changes in the income gain from migration by increasing the ν parameter.

Figure 6 reports overall migration rates over a range of counterfactual β values, using both the benchmark specification (single ν , single β) and the flexible specification (caste-specific ν , β increasing in household income). A 50 percent decline in β , which is still quite far from full insurance (β equal to zero) more than doubles the migration rate from 4 to 9 percent with the more precisely estimated flexible specification, highlighting the importance of risk in restricting mobility. How responsive is migration to an exogenous change in the income gain from migration? Given historically low migration rates, despite the persistently large wage gap in India, we would expect predicted migration to be insensitive to changes in the income gain if the model is correct. Counterfactual simulations in Figure 6 that vary the value of the ν parameter, but retain the assumption that formal insurance is unavailable, verify that this is indeed the case. An 80 percent increase in the ν parameter, which corresponds to an average decline in the income gain from 18 percent to 10 percent lowers migration by just 1 percentage point. This result further emphasizes the central message of this paper, which is that inadequate access to formal insurance, rather than wage differentials as commonly assumed in models of migration, may explain much of the low mobility in India.⁴¹

D. Testing the Mechanism

The key assumption underlying our theory is that permanent male migration is associated with a loss in network services. We test this assumption by examining how a household's relative wealth affects three variables: out-migration, network participation, and out-marriage. Recall that marriage within the caste is a prerequisite for participation in the network.

Each REDS round consists of a census of households in the representative sample of villages, followed by a detailed survey of a sample of households in those villages. The survey collects information on permanent migration by adult males, as in the census. In addition, it collects information on financial transactions within the caste, which directly measures participation in the insurance network, as well as marriage within the caste by household members.⁴²

⁴¹ Counterfactual simulations reported in Appendix Figure A1 shift the ν parameter in the opposite direction. A 50 percent decline in the ν parameter, which corresponds to a doubling of the wage gap from 18 percent to 36 percent, increases the migration rate by just 1.5 percentage points from a baseline of 4.3 percent. However, migration is much more responsive to the wage gap, in absolute and relative terms, when the β parameter is halved from a baseline value of 1.4; migration increases from 6 percent to 9 percent.

⁴² Participation in the caste network is measured by a binary variable that takes the value one if the household sent or received caste-transfers (gifts or loans from one or more members of the same subcaste) in the year preceding each survey round. The measures of out-marriage and out-migration are constructed from the 1999 retrospective histories on the marriages and migration of all of the siblings and children of each household head in the sample. From these histories we created a variable indicating whether any child of the household head married outside the caste in the 10-year period prior to each survey date. The measure of out-migration is whether any male aged 20–30 at the time of each survey and residing in the household prior to the survey had left the village permanently by the survey date.

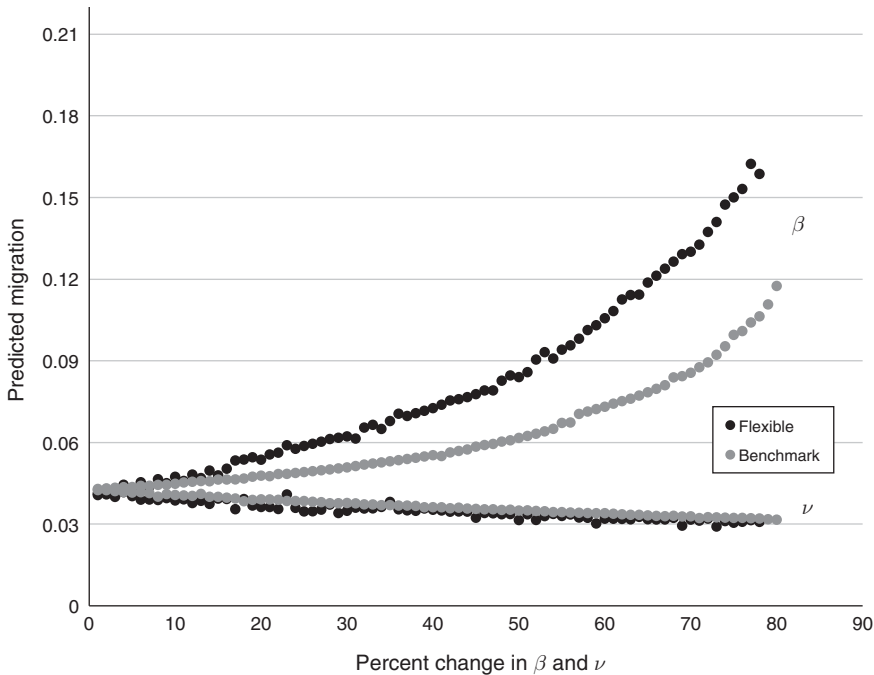


FIGURE 6. COUNTERFACTUAL SIMULATION

Although the number of households in the detailed survey is relatively small, a major advantage over the census is that they can be linked over successive REDS rounds. We thus construct a panel, using the surveyed households in the 1982 and 1999 rounds of the REDS, for the joint test of the key theoretical assumption.⁴³ We eliminate all castes with less than ten households in the REDS survey, and then proceed to estimate the following equation:

$$(11) \quad X_{it} = \pi_1 y_{it} + \pi_2 \bar{y}_{it} + f_i + \epsilon_{it},$$

where y_{it} is household i 's average income in survey round t (1982, 1999), \bar{y}_{it} is the corresponding caste average, and f_i is a household fixed effect. This is the same specification as equation (9), except that household fixed effects and time subscripts are included. Equation (11) is separately estimated with out-migration, out-marriage, and network participation as dependent variables.

Average household income, y_{it} , could be determined by unobserved household attributes that independently determine the outcomes of interest. These attributes could also be correlated with caste income, \bar{y}_{it} , to the extent that they are correlated across households within the caste. Differencing over the two years allows us to purge these fixed attributes,

$$(12) \quad \Delta X_{it} = \pi_1 \Delta y_{it} + \pi_2 \Delta \bar{y}_{it} + \Delta \epsilon_{it}.$$

⁴³To construct the panel, we start with the sample of households in the 1982 round. Because of household partitioning, many 1982 household members were distributed across multiple households in 1999. We thus aggregate all 1999 households to be consistent with 1982 household boundaries, resulting in a balanced panel over the two years.

However, shocks to income could still be correlated with unobserved changes in the determinants of out-migration, out-marriage, and network participation $\Delta\epsilon_{it}$; for example, if schools, banks, or other infrastructure that independently changed incomes *and* the outcomes of interest were introduced in the household's village between 1982 and 1999. To address this concern, we construct instruments for Δy_{it} , $\Delta \bar{y}_{it}$ in equation (12) above.

We make use of two technological features of the Indian Green Revolution to construct these instruments: (i) only certain parts of the country could profit from the new HYV seeds at the onset of the Green Revolution in the late 1960s, and (ii) the returns to investing in the HYV technology are much greater on irrigated land. The instrumental variable strategy allows for the possibility that initial availability of HYV technology and access to irrigation were independently correlated with unobserved village and household characteristics that had long-term effects on the outcomes of interest by including them in the second stage. Exploiting the technological complementarity between HYV and irrigation, and the fact that simultaneous access to these inputs was quasi-random, only the *interaction* of these variables, measured in 1971, is used as an instrument for changes in income from 1982 to 1999. The instrument is constructed by interacting a binary variable indicating whether anyone in household i 's village used HYV in 1971 with the share of irrigated land in the village in that year. These variables are obtained from the 1971 REDS round, which is at the onset of the Green Revolution but still close enough in time to predict changes in income over the 1982–1999 period.

All areas of the country did not benefit equally from the HYV technology at the onset of the Green Revolution (see Munshi 2004 for details). The early rice varieties, in particular, were unsuitable for cultivation in many areas, and it was only by crossbreeding with local varieties that the new technology could be adopted throughout the country. Although the new wheat varieties were more robust, marketing was restricted to specific districts in the early stages of the Green Revolution. Where the new HYV technology did become available, credit constraints would have prevented growers from making the complementary investments in irrigation that were required to exploit the enhanced potential of the new seeds. Access to irrigation at that time would have been largely determined by proximity to water bodies and preexisting irrigation technologies like tanks and canals, which was independent of the program priorities and technological constraints that determined access to HYV. The simultaneous access to HYV technology and irrigation, conditional on access to each of these inputs separately, can thus be treated as quasi-random (satisfying the exclusion restriction). At the same time, the combination of these inputs, net of their direct effects, would still have determined initial farm profits due to the fact that the new HYV technology performed much better on irrigated land. Given credit constraints, these initial profits would have determined the subsequent income trajectory and this is indeed what we observe in the first-stage regression.⁴⁴

The income variable that we construct depends on the household's assets and the number of working age adults (including permanent migrants, if any). The same

⁴⁴To add statistical power, we include two more instruments: the amount of land inherited by the household head, as reported in the 1982 REDS round, and the triple interaction of inherited wealth, HYV, and irrigation. The corresponding caste-level instruments are the caste averages of the three household-level instruments.

procedure was used in the structural estimation to construct the income gain from migration for households with migrants, except that we now use the village-level daily wage in the survey year rather than the NSS wage to measure rural labor income. Caste-level income is once again simply the caste average of household incomes. Note that we no longer need to impute incomes or to restrict the sample to southern states, and so the results that follow serve as a useful robustness test.

Table 8 reports the instrumental-variable estimates of equation (12). The Kleibergen-Paap F -statistics indicate that weak instruments are a concern (based on the Stock-Yogo critical values), particularly with participation in the insurance network as the dependent variable. This variable is measured in a single year prior to each survey round, which will severely underestimate the household's actual involvement, especially since the demand for major contingencies occurs relatively infrequently. This weakens the power of the instruments. The estimates in Table 8, particularly with network participation as the dependent variable, should thus be treated with caution. They nevertheless support the joint hypothesis that conditional on the household's own income an increase in caste income reduces the probability of out-migration and out-marriage, and simultaneously increases the probability of participating in the insurance network.⁴⁵ The point estimates with out-migration as the dependent variable are larger in magnitude than what we obtained in Table 6, possibly because the instruments purge measurement error in the income variables. However, the pattern of coefficients remains the same and the key caste-income coefficient is precisely estimated with all three outcomes. These results are difficult to reconcile with alternative explanations that do not make a connection between caste networks and migration.

IV. Conclusion

This paper provides an explanation for large spatial wage disparities and low male migration in India based on a combination of well-functioning rural insurance networks and the absence of formal insurance. When men migrate permanently to work, they (and their rural households) cannot credibly commit to honoring their future obligations at the same level as households without migrants. They also have an incentive to misreport their urban income. If the loss in network insurance due to these commitment and information problems is sufficiently severe, and alternative insurance is unavailable, then higher paying job opportunities will go unexploited. Imperfections in the insurance market thus give rise to a misallocation in the labor market.

We test this hypothesis by developing and estimating a model of *ex ante* mutual insurance in which participation in the network and the income-sharing rule are jointly determined. The main theoretical results are (i) that income is redistributed in favor of relatively poor households within the caste, which implies that relatively rich households (who benefit less from the insurance network) should be more

⁴⁵ Wild cluster bootstrapped standard errors in Table 8 are clustered at the state rather than the caste level because our instruments are correlated with agricultural extension and irrigation programs at the district level, and because the caste will often span a wide area covering multiple districts within but not across states. Appendix Table A5 reports the first-stage parameter estimates.

TABLE 8—FE-IV MIGRATION, OUT-MARRIAGE, AND PARTICIPATION ESTIMATES

Dependent variable:	Migration (1)	Out-marriage (2)	Network participation (3)
Household income	0.262 (0.147)	0.166 (0.102)	−0.520 (0.684)
Caste income	−0.110 (0.053)	−0.111 (0.054)	0.327 (0.173)
Time trend	0.059 (0.020)	0.026 (0.011)	0.014 (0.123)
Kleibergen-Paap <i>F</i> -statistic	10.52	8.05	2.91
Observations	1,049	998	1,049

Notes: Wild cluster bootstrapped standard errors are clustered at the state level. Income is constructed using wealth- and wage-based measure. Income measured in lakhs of rupees, (1 lakh = 100,000). Additional regressors: whether anyone in the village used HYV and share of village land irrigated in 1971 (household and caste average). Excluded variables: inherited land, interaction of any HYV and irrigation share, interaction of inherited land, any HYV and irrigation share (household and caste average). Sample restricted to castes with at least 10 households in the panel and households with heads at least age 18 in 1982. Stock-Yogo weak ID test: 5 percent critical value 15.72, 10 percent critical value 9.48.

Source: REDS panel 1982 and 1999

likely to have migrant members, and (ii) that households facing greater rural income risk (who benefit more from the network) should be less likely to have migrant members. We find, using a variety of data sources and estimation techniques, evidence that is consistent with these predictions. Structural estimates of the model allow us to quantify the magnitude of the misallocation due to the absence of formal insurance; a 50 percent improvement in risk-sharing for households with migrant members would more than double the migration rate, from 4 to 9 percent.

Why does India have migration rates that are so much lower than other comparable developing economies? In our framework, this could be because formal insurance, which includes private credit and government safety nets, is particularly weak in India or because informal insurance works particularly well there. There is no evidence suggesting that credit markets work better in other low-income countries or that superior public safety nets are available. Moreover, research on informal insurance has documented extremely high levels of risk-sharing throughout the developing world, not just in India. There is, however, more to consumption smoothing than risk-sharing. If the size and geographic scope of the network is small, as it often appears to be, e.g., Udry (1994); Fafchamps and Lund (2003); Angelucci, De Giorgi, and Rasul (2015), then consumption will not be smoothed appreciably even with full risk-sharing.⁴⁶ What is exceptional about India is the size and spatial-scope of the caste network, which appears to have given

⁴⁶To see why this is the case, consider a two-person network and a world with two income states: H and L , that occur with equal probability and are distributed independently across individuals and over time. With full risk-sharing, each individual consumes H with probability $1/4$, L with probability $1/4$, and $(H + L)/2$ with probability $1/2$, so there is still substantial variation in consumption. This variation will decline, however, as the number of individuals in the network increases. It will also decline if incomes are negatively correlated; i.e., if income risk can be diversified, which will be the case if the network is more dispersed.

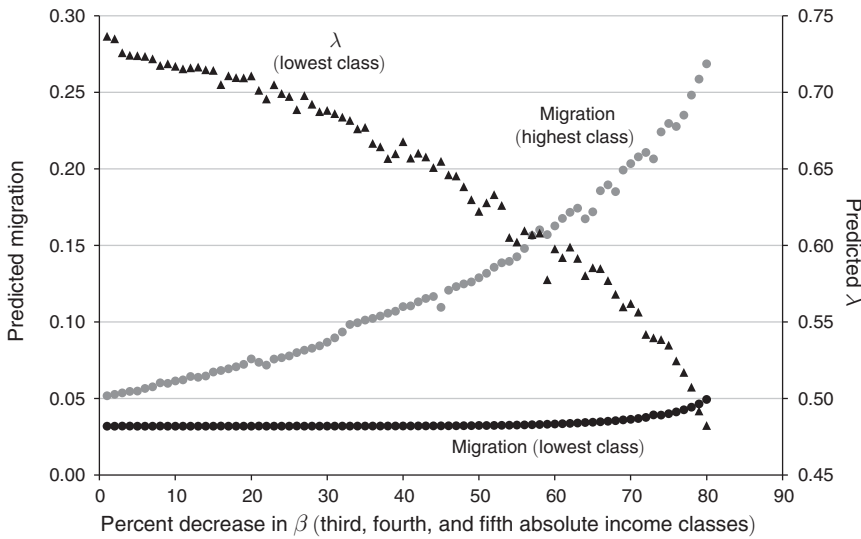


FIGURE 7. REDUCING RISK IN HIGHER INCOME CLASSES

rise to an equilibrium with strong rural insurance networks, weak formal insurance, and low mobility.

The model was developed to explain low mobility in India, but it is also useful in assessing the mobility and distributional impacts of interventions that provide formal insurance to rural households. One strategy would be to increase access to private credit, perhaps by allowing rural households to collateralize their assets. In our model, this would result in a decline in β for wealthier households. Figure 7 reports the results of a policy experiment in which we proportionately reduce β in the top three absolute income classes, using the flexible specification in Table 7, columns 9 and 10.⁴⁷ As can be seen, migration increases substantially for the highest relative income class as β declines. Although not reported, there is a substantial increase in migration for the next two income classes as well. The accompanying decline in the size of the network would adversely affect the ability of the households that remain to smooth their consumption. More interestingly, there is a substantial reduction in redistribution within the caste, as a way of getting the wealthier households to stay. For the lowest income class, for example, λ declines from 0.75 to 0.48, which is not far from the sharing rule without redistribution ($\lambda = 0.31$). Thus, while a credit program may reduce the labor-market misallocation, it will have large and unintended negative distributional consequences for the lower income classes who continue to have low migration rates. An evaluation of this credit program that failed to account for the interaction of the treatment with

⁴⁷ Access to credit depends on absolute (not relative) wealth. Each relative income class, k , within each caste, j , is thus assigned to one of five absolute income classes on the basis of its mean income, $M_{A_{kj}}$. Recall that the absolute income classes were constructed by sorting households in the REDS census by average income, without regard to caste affiliation. We lower $\beta \equiv \alpha + \gamma M_{A_{kj}}$ for those relative income classes, kj , that are in the top three absolute income classes in Figure 7. The counterfactual simulation that follows in Figure 8 reduces β for relative income classes, kj , in the bottom two absolute income classes.

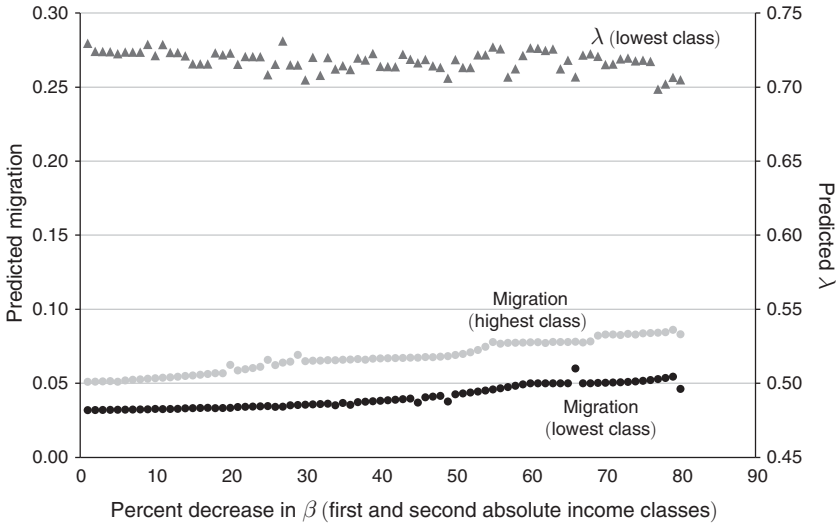


FIGURE 8. REDUCING RISK IN LOWER INCOME CLASSES

underlying informal institutions would be hard-pressed to explain why consumption *declined* in the untreated group.

We can also use the structural estimates to assess an alternative policy intervention that reduces β for lower absolute income classes. The policy in this case could be a government employment guarantee scheme for the rural poor, such as India's NREGA. If rural insurance networks were ignored, one would expect such a scheme to improve the welfare of the poor and reduce their migration from rural areas. What we observe instead in Figure 8, consistent with our theory, is that migration is increasing for the lowest (treated) income classes as β declines. Migration increases as much for the highest income class (which does not benefit directly from the scheme) as the lowest income class. The migration rate increases in the highest income class because the increased exit by the directly-targeted low income households reduces the ability of the network to smooth consumption. Once again, spillovers within the network have substantial impacts on household behavior.

APPENDIX

PROOF OF LEMMA 1:

$\Delta M = 0$ from equation (2). $\Delta R > 0$ by assumption. The term in parentheses in equation (4) is greater than zero. This implies, from Assumption 1, that $F(\Delta M + \Delta R)$ is greater than zero. $F(\Delta M + \Delta R) > N/P$ at $N = 0$. Moreover, $F(\Delta M + \Delta R)$ is less than one from Assumption 2. $F(\Delta M + \Delta R) < N/P$ at $N = P$. $F(\Delta M + \Delta R)$ is increasing in N because R_t is decreasing in N (hence, ΔR must be increasing in N). By a continuity argument, a fixed point, $N^* \in (0, P)$ at which equation (4) is satisfied must exist.

From Assumption 3, and given that $dR_I/dN < 0$, $d^2R_I/dN^2 > 0$ from equations (2) and (3), $F(\Delta M + \Delta R(N))$ is strictly concave:

$$(13) \quad F'(\Delta M + \Delta R) = -f \cdot \frac{1}{2} \frac{dR_I}{dN} > 0$$

$$(14) \quad F''(\Delta M + \Delta R) = f' \cdot \left[\frac{1}{2} \frac{dR_I}{dN} \right]^2 - f \cdot \frac{1}{2} \frac{d^2R_I}{dN^2} < 0.$$

This implies that the fixed point, which satisfies equation (4), is unique to complete the proof of Lemma 1. ■

PROOF OF PROPOSITION 1:

We first establish that a unique fixed point exists for the sharing rule without redistribution.

LEMMA 2: *Equilibrium participation in each income class $k \in \{L, H\}$ is characterized by a unique fixed point, $N_k^* \in (0, P_k)$.*

Without redistribution, $\Delta M_k = 0$ from equation (5). As with the case without income heterogeneity, we assume that $\Delta R_k > 0$. The term in parentheses in equation (7) is positive for both income classes. The right-hand side of the equation is strictly positive from Assumption 1 and less than one from Assumption 2. Following the same argument as in Lemma 1, this implies that the F function must cross the N_k/P_k line in equation (7) at least once.

For a unique fixed point to be obtained, we need in addition that the F function should be strictly concave. The conditions for strict concavity corresponding to inequalities (13) and (14) are

$$F'(\Delta M_k + \Delta R_k) = f \cdot \left(\frac{1}{M_{Ik}} \frac{dM_{Ik}}{dN_k} - \frac{1}{2} \frac{dR_I}{dN_k} \right) > 0$$

$$F''(\Delta M_k + \Delta R_k) = f' \cdot \left(\frac{1}{M_{Ik}} \frac{dM_{Ik}}{dN_k} - \frac{1}{2} \frac{dR_I}{dN_k} \right)^2 + f \cdot \left(\frac{d}{dN_k} \left[\frac{1}{M_{Ik}} \frac{dM_{Ik}}{dN_k} \right] - \frac{1}{2} \frac{d^2R_I}{dN_k^2} \right) < 0.$$

Without redistribution, $\lambda_k = M_{Ak}/M_{AK}$. It follows from equation (5) that $M_{Ik} = M_{AK}$ and, hence, that $dM_{Ik}/dN_k = 0$. Given Assumption 3, the preceding inequalities will evidently be satisfied if $dR_I/dN_k < 0$ and $d^2R_I/dN_k^2 > 0$.

From equation (6),

$$(15) \quad \frac{dR_I}{dN_k} = \frac{V_{Ak} \sum_k N_k M_{Ak} - 2 M_{Ak} \sum_k N_k V_{Ak}}{(\sum_k N_k M_{Ak})^3}.$$

Given that R_{Ak} is the same (denoted by R) in all income classes, and that N_k is the same (denoted by N) in all income classes in the absence of redistribution since they are of equal size,

$$\frac{dR_I}{dN_k} = \frac{R [M_{Ak}^2 \sum_k M_{Ak} - 2 M_{Ak} \sum_k M_{Ak}^2]}{N^2 (\sum_k M_{Ak})^3}.$$

Collecting terms, the required condition is

$$(16) \quad M_{Ak} < 2 \frac{\sum_k M_{Ak}^2}{\sum_k M_{Ak}}.$$

From equation (15),

$$(17) \quad \frac{d^2 R_I}{dN_k^2} = \frac{-4 V_{Ak} M_{Ak} \sum_k N_k M_{Ak} + 6 M_{Ak}^2 \sum_k N_k V_{Ak}}{(\sum_k N_k M_{Ak})^4}.$$

Simplifying as above,

$$\frac{d^2 R_I}{dN_k^2} = \frac{R [-4 M_{Ak}^3 \sum_k M_{Ak} + 6 M_{Ak}^2 \sum_k M_{Ak}^2]}{N^3 (\sum_k M_{Ak})^4}.$$

Collecting terms, the required condition is

$$(18) \quad M_{Ak} < \frac{3}{2} \frac{\sum_k M_{Ak}^2}{\sum_k M_{Ak}}.$$

The condition in equation (18) is binding. For the case with two income classes, assume without loss of generality that $M_{AL} = M(1 - \theta)$, $M_{AH} = M(1 + \theta)$. We showed in Lemma 1 that there is a unique fixed point when $\theta = 0$. We now show that the condition in equation (18) is satisfied for all $\theta \geq 0$. With two income classes, that condition can be rewritten as

$$M_{AH} < \frac{3}{2} \frac{(M_{AL}^2 + M_{AH}^2)}{(M_{AL} + M_{AH})},$$

which reduces to

$$3\theta^2 - 2\theta + 1 > 0.$$

The left-hand side of the preceding inequality is positive for $\theta \geq 0$ (reaching a minimum value of $2/3$ at $\theta = 1/3$) to complete the proof of Lemma 2.

The next step in proving Proposition 1 is to show that an increase in λ , evaluated at $\lambda = M_{AL}/M_{AH}$, must increase the surplus generated by the insurance network,

$$W = \sum_{k=L,H} N_k \epsilon_{Ik} - P_k \int_0^{\epsilon_{Ik}} \epsilon f(\epsilon) d\epsilon,$$

where $\epsilon_{Ik} = \Delta M_k + \Delta R_k$.

Differentiate W with respect to λ , applying Leibniz integral rule and noting that $\Delta M_k, R_I$ are functions of N_L and N_H ,

$$\frac{dW}{d\lambda} = \sum_{k=L,H} [N_k - P_k \epsilon_{Ik} f(\epsilon_{Ik})] \frac{d\epsilon_{Ik}}{d\lambda} + \left\{ \epsilon_{Ik} + \sum_{m=L,H} [N_m - P_m \epsilon_{Im} f(\epsilon_{Im})] \frac{d\epsilon_{Im}}{dN_k} \right\} \frac{dN_k}{d\lambda}.$$

$N_k = P_k F(\epsilon_{Ik})$ and so the terms in square brackets must be positive. Moreover, at $\lambda = M_{AL}/M_{AH}$, $\epsilon_{IL} = \epsilon_{IH}$ and $N_L - P_L \epsilon_{IL} f(\epsilon_{IL}) = N_H - P_H \epsilon_{IH} f(\epsilon_{IH})$. Thus, the sign of $dW/d\lambda$ depends on $d\epsilon_{Ik}/d\lambda$, $d\epsilon_{Im}/dN_k$, $dN_k/d\lambda$. We show below that $d\epsilon_{IL}/d\lambda > |d\epsilon_{IH}/d\lambda|$, $d\epsilon_{Im}/dN_L > d\epsilon_{Im}/dN_H$, $dN_L/d\lambda > |dN_H/d\lambda|$ to establish that $dW/d\lambda > 0$.

Since R_I is independent of λ ,

$$\frac{d\epsilon_{Ik}}{d\lambda} = \frac{1}{M_{Ik}} \frac{dM_{Ik}}{d\lambda}.$$

From equation (5), with two income classes,

$$\frac{1}{M_{IL}} \frac{dM_{IL}}{d\lambda} = \frac{N_H}{\lambda(\lambda N_L + N_H)}$$

$$\frac{1}{M_{IH}} \frac{dM_{IH}}{d\lambda} = \frac{-N_L}{(\lambda N_L + N_H)}.$$

At $\lambda = M_{AL}/M_{AH}$, $N_L = N_H$ since the two income classes are of equal size. Since $\lambda < 1$, $d\epsilon_{IL}/d\lambda > |d\epsilon_{IH}/d\lambda|$ and so the direct effect of an increase in λ on W is positive.

$$\frac{d\epsilon_{Im}}{dN_k} = \frac{1}{M_{Im}} \frac{dM_{Im}}{dN_k} - \frac{1}{2} \frac{dR_I}{dN_k}.$$

We know from Lemma 2 that $dM_{Im}/dN_k = 0$ when there is no redistribution. We also know from Lemma 2 that $dR_l/dN_k < 0$. We thus need to show that $|dR_l/dN_L| > |dR_l/dN_H|$.

From equation (15), the required condition is

$$\frac{V_{AL}}{V_{AH}} > \frac{(N_L + 2N_H) \frac{M_{AL}}{M_{AH}} - N_H}{(2N_L + N_H) - N_L \frac{M_{AL}}{M_{AH}}}.$$

$V_{AL}/V_{AH} = (M_{AL}/M_{AH})^2$ because $R_{AL} = R_{AH}$ by assumption. It follows that both the left-hand side (LHS) and the right-hand side (RHS) of the preceding inequality are increasing and convex functions of M_{AL}/M_{AH} . It is straightforward to verify that the LHS starts above the RHS at $M_{AL}/M_{AH} = 0$, cuts it from above at $M_{AL}/M_{AH} = N_H/N_L$, and then converges to the RHS from below at $M_{AL}/M_{AH} = 1$. The inequality, $LHS > RHS$, is thus satisfied for $M_{AL}/M_{AH} < N_H/N_L$. $N_L = N_H$ when $\lambda = M_{AL}/M_{AH}$. Since $M_{AL}/M_{AH} < 1$ by construction, the preceding condition is always satisfied, ensuring that $|dR_l/dN_L| > |dR_l/dN_H|$.

To show that $dN_L/d\lambda > |dN_H/d\lambda|$, apply the Implicit Function theorem to the fixed-point equation (7), which we know has a unique solution from Lemma 2, to obtain

$$\frac{dN_k}{d\lambda} = \frac{f(\epsilon_{Ik}) \frac{d\Delta M_k}{d\lambda}}{\frac{1}{P_k} - f(\epsilon_{Ik}) \left(\frac{d\Delta M_k}{dN_k} - \frac{1}{2} \frac{dR_l}{dN_k} \right)}.$$

Recall from Lemma 2 that the slope of the F function is shallower than the slope of the straight line, $1/P_k$, at the fixed-point since it cuts it from above. This implies that the denominator of the preceding equation must be positive for each income class, $k \in \{L, H\}$. When $\lambda = M_{AL}/M_{AH}$, $f(\epsilon_{IL}) = f(\epsilon_{IH})$, $d\Delta M_L/dN_L = d\Delta M_H/dN_H = 0$, and as shown above, $|dR_l/dN_L| > |dR_l/dN_H|$. $P_L = P_H$ because income classes are of equal size. This implies that the denominator must be smaller for the low income class. Turning to the numerator, we showed above that $\frac{d\Delta M_L}{d\lambda} > |\frac{d\Delta M_H}{d\lambda}|$ at $\lambda = M_{AL}/M_{AH}$. Since $f(\epsilon_{IL}) = f(\epsilon_{IH})$, the numerator of the preceding equation will be greater for the low income class (in absolute magnitude), reinforcing the difference in the denominator derived above, to establish that $dN_L/d\lambda > |dN_H/d\lambda|$.

The indirect participation effect reinforces the direct effect, implying that $dW/d\lambda$ is unambiguously positive at $\lambda = M_{AL}/M_{AH}$. Thus, there must be redistribution in equilibrium, $\lambda^* > M_{AL}/M_{AH}$, to complete the proof. ■

PROOF OF PROPOSITION 2:

To prove Proposition 2, replace income classes, $k \in \{L, H\}$, with risk classes, $k \in \{R, S\}$, and appropriately relabel key equations and inequalities that were used to prove Proposition 1:

$$W = \sum_{k=R,S} N_k \epsilon_{Ik} - P_k \int_0^{\epsilon_{Ik}} \epsilon f(\epsilon) d\epsilon,$$

$$\frac{dW}{d\lambda} = \sum_{k=R,S} [N_k - P_k \epsilon_{Ik} f(\epsilon_{Ik})] \frac{d\epsilon_{Ik}}{d\lambda}$$

$$+ \left\{ \epsilon_{Ik} + \sum_{m=R,S} [N_m - P_m \epsilon_{Im} f(\epsilon_{Im})] \frac{d\epsilon_{Im}}{dN_k} \right\} \frac{dN_k}{d\lambda}.$$

We assumed when proving Proposition 1 that $M_{AL} < M_{AH}$, $R_{AL} = R_{AH}$. Since the community is now divided by risk, we assume instead that $R_{AR} > R_{AS}$, $M_{AR} = M_{AS}$. Without redistribution, $\lambda = M_{AR}/M_{AS} = 1$. $\Delta M_R = \Delta M_S = 0$. $\Delta R_R > \Delta R_S$. Given that $\epsilon_{Ik} = \Delta M_k + \Delta R_k$, this implies that $\epsilon_{IR} > \epsilon_{IS}$. The ϵ distribution, characterized by the F function, is assumed to be the same in both risk classes. Risk classes are of equal size; $P_R = P_S$. Since $N_k = P_k F(\epsilon_{Ik})$, it follows that $N_R > N_S$. If the surplus increases with redistribution, we will see below that λ can decline only as far as a threshold $\underline{\lambda}$ at which $N_R = N_S$; $\lambda \in [\underline{\lambda}, 1]$. We evaluate $dW/d\lambda$ at $\lambda = 1$ and $\lambda = \underline{\lambda}$. We will see that the sign of the derivative is ambiguous at $\lambda = 1$ but strictly positive at $\lambda = \underline{\lambda}$.

As with the proof of Proposition 1, we examine $d\epsilon_{Ik}/d\lambda$, $d\epsilon_{Im}/dN_k$, and $dN_k/d\lambda$, in turn:

$$\frac{d\epsilon_{Ik}}{d\lambda} = \frac{1}{M_{Ik}} \frac{dM_{Ik}}{d\lambda}$$

$$\frac{1}{M_{IR}} \frac{dM_{IR}}{d\lambda} = \frac{N_S}{\lambda(\lambda N_R + N_S)}$$

$$\frac{1}{M_{IS}} \frac{dM_{IS}}{d\lambda} = \frac{-N_R}{(\lambda N_R + N_S)}.$$

Without redistribution, $\lambda = 1$ and $N_R > N_S$. This implies that $d\epsilon_{IR}/d\lambda < |d\epsilon_{IS}/d\lambda|$. Noting that $N_k = P_k F(\epsilon_{Ik})$, it follows that $N_R - P_R \epsilon_{IR} f(\epsilon_{IR}) > N_S - P_S \epsilon_{IS} f(\epsilon_{IS})$. The direct effect of an increase in λ on W is consequently ambiguous. If the sign of the derivative is positive, $\lambda = 1$ and $N_R > N_S$ in equilibrium. If the sign of the derivative is negative, the surplus can be increased by reducing λ , but only as long as $N_R > N_S$. As λ declines, the gap between N_R and N_S declines. At $\underline{\lambda} < 1$, $\epsilon_{IR} = \epsilon_{IS}$ and $N_R = N_S$. From the equations above, the direct effect is unambiguously positive at $\lambda = \underline{\lambda}$,

$$\frac{d\epsilon_{Im}}{dN_k} = \frac{1}{M_{Im}} \frac{dM_{Im}}{dN_k} - \frac{1}{2} \frac{dR_I}{dN_k}.$$

From equation (5),

$$M_{Ik} = \left(\frac{\lambda_k}{\sum_k \lambda_k N_k} \right) \sum_k N_k M_{Ak}.$$

$$M_{IR} = \left(\frac{\lambda}{\lambda N_R + N_S} \right) (N_R M_{AR} + N_S M_{AS}) \quad M_{IS} = \left(\frac{1}{\lambda N_R + N_S} \right) (N_R M_{AR} + N_S M_{AS}).$$

$$\frac{1}{M_{IR}} \frac{dM_{IR}}{dN_R} = \frac{(1-\lambda)}{(\lambda N_R + N_S)} \frac{N_S}{N_R + N_S} \quad \frac{1}{M_{IR}} \frac{dM_{IR}}{dN_S} = \frac{-(1-\lambda)}{(\lambda N_R + N_S)} \frac{N_R}{N_R + N_S}$$

$$\frac{1}{M_{IS}} \frac{dM_{IS}}{dN_R} = \frac{(1-\lambda)}{(\lambda N_R + N_S)} \frac{N_S}{N_R + N_S} \quad \frac{1}{M_{IS}} \frac{dM_{IS}}{dN_S} = \frac{-(1-\lambda)}{(\lambda N_R + N_S)} \frac{N_R}{N_R + N_S}.$$

Without redistribution ($\lambda = 1$) we have already noted that $d\Delta M_{Im}/dN_k = 0$. With redistribution, the preceding equations indicate that $d\Delta M_{IR}/dN_R > d\Delta M_{IR}/dN_S$, $d\Delta M_{IS}/dN_R > d\Delta M_{IS}/dN_S$:

$$\frac{dR_I}{dN_k} = \frac{V_{Ak} \sum_k N_k M_{Ak} - 2M_{Ak} \sum_k N_k V_{Ak}}{(\sum_k N_k M_{Ak})^3}.$$

Given that M_{Ak} is the same in all risk-classes, it is straightforward to show that $dR_I/dN_k < 0$. The required condition, from the preceding equation, is

$$V_{Ak} < 2 \frac{\sum_k N_k V_{Ak}}{\sum_k N_k}.$$

With two risk-classes, the binding condition is

$$V_{AR} < 2 \frac{(N_R V_{AR} + N_S V_{AS})}{(N_R + N_S)},$$

which reduces to

$$V_{AR}(N_R - N_S) + 2N_S V_{AS} > 0.$$

$N_R \geq N_S$ for $\lambda \in [\underline{\lambda}, 1]$, which implies that this condition is always satisfied.

Given that $dR_I/dN_k < 0$, we can show that $|dR_I/dN_R| > |dR_I/dN_S|$. The required condition is

$$\frac{V_{AR}}{V_{AS}} > \frac{(N_R + 2N_S) \frac{M_{AR}}{M_{AS}} - N_S}{(2N_R + N_S) - N_R \frac{M_{AR}}{M_{AS}}},$$

which is always satisfied since $M_{AR} = M_{AS}$ and $V_{AR} > V_{AS}$.

For $m \in \{R, S\}$, $d\Delta M_{Im}/dN_R = \Delta M_{Im}/dN_S = 0$ when $\lambda = 1$. $d\Delta M_{Im}/dN_R > d\Delta M_{Im}/dN_S$ when $\lambda = \underline{\lambda}$. $|dR_I/dN_R| > |dR_I/dN_S|$. Thus,

$d\epsilon_{Im}/dN_R > d\epsilon_{Im}/dN_S$ for $\lambda = 1$ and $\lambda = \underline{\lambda}$. When $\lambda = 1$, $\epsilon_{IR} > \epsilon_{IS}$ and $N_R - P_R \epsilon_{IR} f(\epsilon_{IR}) > N_S - P_S \epsilon_{IS} f(\epsilon_{IS})$. When $\lambda = \underline{\lambda}$, $\epsilon_{IR} = \epsilon_{IS}$ and $N_R - P_R \epsilon_{IR} f(\epsilon_{IR}) = N_S - P_S \epsilon_{IS} f(\epsilon_{IS})$. The term in curly brackets in the $dW/d\lambda$ equation is thus unambiguously larger for the risky class:

$$\frac{dN_k}{d\lambda} = \frac{f(\epsilon_{Ik}) \frac{d\Delta M_k}{d\lambda}}{\frac{1}{P_k} - f(\epsilon_{Ik}) \left(\frac{d\Delta M_k}{dN_k} - \frac{1}{2} \frac{dR_I}{dN_k} \right)}.$$

At $\lambda = 1$, $\epsilon_{IR} > \epsilon_{IS}$, $d\Delta M_k/dN_k = 0$, and $|dR_I/dN_R| > |dR_I/dN_S|$. At $\lambda = \underline{\lambda}$, $\epsilon_{IR} = \epsilon_{IS}$, $d\Delta M_R/dN_R > d\Delta M_S/dN_S$, and $|dR_I/dN_R| > |dR_I/dN_S|$. The denominator of the right-hand side of the preceding equation is unambiguously smaller for the risky class. However, without redistribution, $d\Delta M_R/d\lambda < |d\Delta M_S/d\lambda|$. The numerator is not necessarily larger for the risky class. At $\underline{\lambda}$, however, $\epsilon_{IR} = \epsilon_{IS}$, and we saw above that $d\Delta M_R/d\lambda > |d\Delta M_S/d\lambda|$. It follows that $dN_R/d\lambda > |dN_S/d\lambda|$.

Each term on the right-hand side of the $dW/d\lambda$ equation is positive at $\lambda = \underline{\lambda}$. This implies that $\lambda > \underline{\lambda}$ in equilibrium and, hence, that $N_R > N_S$ to complete the proof. ■

TABLE A1—RURAL-URBAN WAGE GAPS IN INDIA OVER TIME

Year	Nominal wage		Consumer price index		Real wage		Wage gap (%) (7)
	Rural	Urban	Rural	Urban	Rural	Urban	
	(1)	(2)	(3)	(4)	(5)	(6)	
1983	5.94	9.96	81.90	111.20	24.81	40.17	61.92
1993	17.66	28.79	177.11	258.00	34.11	50.05	46.75
1999	34.80	54.29	309.00	428.00	38.52	56.90	47.73
2004	42.54	62.66	342.00	520.00	42.54	54.05	27.04
2009	80.74	104.70	530.00	754.70	52.10	62.23	19.44

Notes: Consumer price index (CPI) is based on Government of India statistics. Conversion to real wage in 2004 is based on representative rural consumption bundle in that year. CPI is used to adjust wages in other years.

Source: Nominal wage per day is derived from NSS

TABLE A2—LOAN CHARACTERISTICS BY SOURCE

	Loan type	
	Without interest	Without collateral
	(1)	(2)
Bank	-0.243 (0.021)	-0.329 (0.037)
Friends	-0.058 (0.041)	0.127 (0.072)
Employer	-0.227 (0.030)	0.009 (0.052)
Moneylender	-0.247 (0.025)	-0.005 (0.044)
Other	0.075 (0.061)	0.206 (0.108)
Observations	1,045	1,045

Notes: The other categories are not reported. Observations are weighted by the value of the loan and sample weights. Caste is the omitted source in these regressions.

Source: REDS (1982)

TABLE A3—HOUSEHOLD CHARACTERISTICS, INCOME GAIN, MIGRATION RATE, AND ν PARAMETER BY ABSOLUTE INCOME CLASSES

	Household income	Land value	Number of working adults	Education of household head	Income gain from migration	Migration rate	ν parameter
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>Absolute income class</i>							
1	0.023	38,084.74	1.462	4.684	0.218	0.043	19.609
2	0.040	35,956.19	1.451	3.909	0.228	0.010	27.352
3	0.051	114,863.00	1.656	3.744	0.179	0.024	28.030
4	0.100	129,849.50	1.777	5.083	0.190	0.051	21.576
5	0.894	90,414.75	1.950	4.741	0.166	0.026	27.329

Notes: Absolute income classes are defined by quintiles across the entire income distribution, without regard to caste affiliation. Household income based on assets, number of working adults, education (from REDS), and rural-urban wages (from NSS). Income gain is computed assuming a single member of the household migrates. Migration rate is obtained from the REDS, and together with the income gain from migration, is used to derive the ν parameter.

Source: REDS census 2006

TABLE A4—STRUCTURAL ESTIMATES WITH FOUR AND SIX RELATIVE INCOME CLASSES

Relative income class	Four relative income classes				Six relative income classes			
	Actual		Predicted		Actual		Predicted	
	Relative consumption (1)	Migration (2)	Relative consumption (3)	Migration (4)	Relative consumption (5)	Migration (6)	Relative consumption (7)	Migration (8)
1	0.857	0.031	0.817 (0.059)	0.000 (0.0008)	0.831	0.033	0.785 (0.052)	0.000 (0.0000)
2	0.880	0.042	0.837 (0.058)	0.022 (0.0084)	0.837	0.027	0.797 (0.052)	0.013 (0.0073)
3	0.894	0.048	0.866 (0.044)	0.051 (0.0077)	0.854	0.050	0.812 (0.049)	0.025 (0.0106)
4	1.000	0.050	1.000	0.097 (0.0138)	0.854	0.047	0.829 (0.045)	0.043 (0.0110)
5	—	—	—	—	0.883	0.047	0.867 (0.037)	0.064 (0.0102)
6	—	—	—	—	1.000	0.053	1.000	0.113 (0.0178)
Overall		0.043		0.043 (0.0060)		0.043		0.043 (0.0063)
β			1.319 (0.813)				1.487 (0.814)	

Notes: Structural estimation is based on 100 castes with at least 30 households in the census. Four and six equal-sized income classes are constructed within each caste. Benchmark specification (single ν) in all estimations. Consumption is measured relative to the highest income class in column 1 and column 5. Relative consumption and migration are computed as the average in each relative income class across all castes. Jackknifed standard errors in parentheses.

Source: REDS census 2006

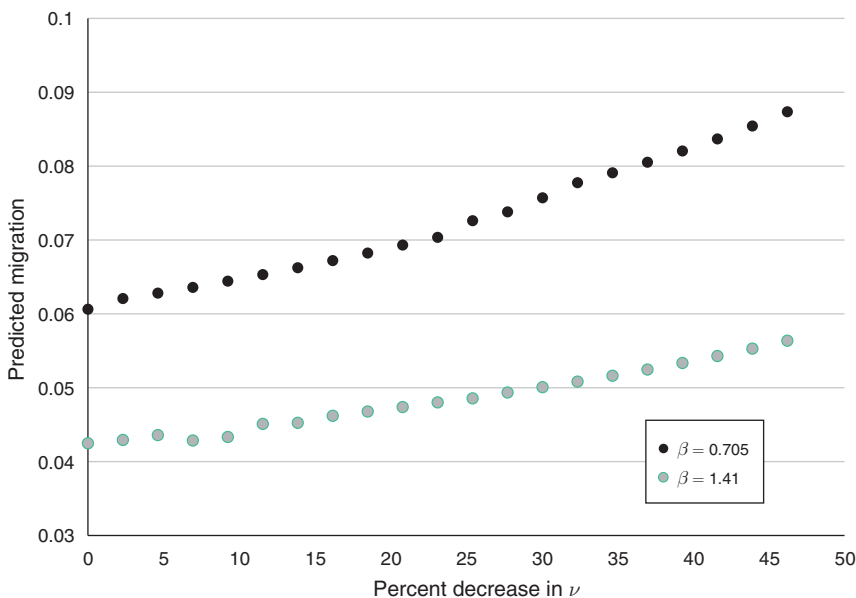


FIGURE A1. COUNTERFACTUAL SIMULATIONS

TABLE A5—FIRST-STAGE ESTIMATES

	Household income change (1)	Caste income change (2)
Anyone in the village used HYV (household)	-2,787.65 (3,719.06)	20,369.17 (17,642.30)
Anyone in the village used HYV (caste average)	-3,348.87 (4,990.63)	-42,467.91 (39,691.53)
Share of village land irrigated (household)	-56.24 (4,063.92)	5,578.18 (9,886.26)
Share of village land irrigated (caste average)	-4,094.89 (3,144.29)	-22,929.07 (32,659.12)
HYV × irrigation share (household)	1,844.46 (5,847.23)	-26,942.14 (22,377.58)
HYV × irrigation share (caste average)	-1,646.46 (6,428.40)	24,413.53 (45,272.26)
Inherited land (household)	3.48 (1.12)	-5.32 (4.55)
Inherited land (caste average)	6.48 (7.88)	24.11 (43.01)
HYV × irrigation share × inherited land (household)	3.44 (6.06)	6.58 (14.78)
HYV × irrigation share × inherited land (caste average)	41.09 (19.56)	140.68 (37.31)
Constant	8,084.81 (2,828.20)	25,683.28 (20,544.49)
<i>F</i> -statistic (excluded variables)	104.96	12.28
<i>p</i> -value	0.00	0.00
<i>R</i> ²	0.02	0.07
Observations	2,335	2,335

Notes: Standard errors in parentheses are robust to clustering at the state level. Dependent variables are computed as the change between 1982 and 1999. Income is constructed using a wealth-based measure. Excluded variables: HYV × irrigation, inherited land, HYV × irrigation × inherited land (household and caste average). Regressions restricted to castes with at least 10 households in sample and households with heads at least age 18 in 1982.

Source: REDS panel 1982 and 1999

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