Risk Sharing and Internal Migration

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Over the past two decades over half the population in rural Tanzania migrated within the country. Such mass internal migration has created geographically disperse networks, on which the authors collected detailed panel data. By quantifying how shocks and consumption co-vary across linked households, we show how migrants unilaterally insure their extended family members at home. This finding contradicts risk sharing models based on reciprocity, but is consistent with assistance driven by social norms. Migrants sacrifice 2 to 7 per cent of their very substantial consumption growth to provide this insurance, which seems too trivial to have any stifling effect on their growth through migration.

JEL codes: O12, O15, O17, R23

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1. Introduction

If, in the next decades, Africa catches up with the rest of the world, then that will almost certainly coincide with intergenerational mobility out of rural into urban areas and out of agriculture into non-agricultural activities (Lewis, 1954; Harris and Todaro, 1970). Historically, in both rich developed countries and fast-growing developing countries, this type of migration has moved in lockstep with development and poverty reduction (Collier and Dercon, 2009). Recently, China's urban population officially surpassed its rural one: of China's 1.35 billion people, 51.3 per cent lived in urban areas at the end of 2011, rising from under 20 per cent in 1980 (UN, 2012). Furthermore, UNDP (2009) reports that out of the one billion migrants worldwide, three quarters are internal migrants. With international migration open to only very few Africans, we should expect massive internal migration to form a core part of the development process.

The scale of this demographic process is captured in the data that form the basis of this paper, further motivating our focus on internal migration. These data are part of an exceptional panel data set from the Kagera region in Tanzania, spanning nearly two decades of migration and development. The 2010 follow-up survey attempted to trace all 6,353 individuals listed on the baseline 1991/94 household rosters and re-interview them irrespective of their location. Once we exclude the 1,275 individuals who had died by 2010, we are left with 4,996 baseline individuals whose 2010 locations are known.[1] Of those, 45 per cent were found residing in the baseline village, 53 per cent had migrated within the country, 2 per cent to another East African country (primarily Uganda) and 0.3 per cent had moved outside of East Africa. This region – not atypical of remote rural Africa – is clearly on the move, with internal migration dwarfing international migration.

This internal mobility is associated with large income gains. Our data show that despite only minor welfare differences during the 1991-94 baseline survey, those who moved out of the region to other parts of Tanzania have grown roughly twice as rich as those who did not by the time we interview them again nearly two decades later.

[1] We miss location information on 82 individuals. Because this is after multiple attempts through various sources it is unlikely that these individuals have moved outside of East Africa. Information on such an important, low-occurrence event is unlikely to be hidden.

As we are measuring consumption and not income, it is clear that the main beneficiaries of this migration-led growth were the migrants themselves and certainly not their relatives who remained at home.

But did these migrants simply leave and never look back, or did they maintain links with the home community? We investigate this question by exploiting the fact that the 3,314 households interviewed in 2010 are grouped in 817 geographically disperse extended family networks. Using techniques from the risk sharing literature we quantify how migrants' consumption responds to shocks experienced by others in their extended family network. Much of the migration literature has a very strong focus on dealing with the selectivity of the migration decision. Interestingly, in this paper the endogeneity of migration turns out to be irrelevant for our most important contribution: the documentation of the long-run dynamics of risk-sharing arrangements among extended family members in a context with large amounts of internal migration. Whether or not migration is causally responsible for any of our findings is an interesting, but secondary question, which we will not attempt to answer. Section 4 explains this point in more detail.

Our analysis departs from a number of other studies in the migration literature by focusing on consumption instead of transfers. This choice of the outcome variable is motivated by the fact that risk sharing and other economic exchange could happen through a multitude of different mechanisms, of which transfers is just one. Other mechanisms could include looking for a job for someone, employing them directly, providing them with tips, advice or a network link, or providing migration opportunities (Munshi, 2003). By analysing consumption we focus on the joint and final effect of all such mechanisms.

The observed divergence between migrants and non-migrants in these data also persists within extended family networks (Beegle, De Weerdt and Dercon, 2011), which violates the full risk sharing hypothesis (Townsend, 1994), and does not support the notion that migration is the result of a household level maximisation strategy (Stark and Bloom, 1985; Rosenzweig and Stark, 1989; Grimard, 1997). It could, however, be consistent with other reciprocity-based models (e.g. limited commitment, moral hazard, or hidden income) that permit the co-existence of

divergent consumption growth and risk sharing. In our empirical analysis, we find that migrants are affected by shocks to others in the network whereas non-migrants are not. Such unilateral insurance leads us to reject the reciprocity-based risk sharing models.

One explanation to this observed lack of reciprocity could be that migrants insure non-migrants in exchange for other benefits (Lucas and Stark, 1985; Hoddinott, 1994). These benefits could accrue to the migrant later in life and outside the purview of our survey data. We consider, but reject a number of such longer-run transactional motives for the observed unilateral insurance. The results are, however, very much line with findings from the diverse literature on social norms (Platteau, 2000; Cox and Fafchamps, 2008; Burke and Young, 2009), where those who move ahead remain obligated to their extended family in the home community.

These results speak to an emerging literature that worries about home communities imposing a stifling 'kin-tax' on the upwardly mobile. Baland, Guirkinger and Mali (2011) show how people take out costly loans in order to conceal their income, while Platteau (2010) sees migration as a means to escape the implied prying eyes and incessant demands of the kinship group. The kinship poverty trap model of Hoff and Sen (2006) predicts possible resistance from the home communities as they feel threatened by productive forces leaving and severing links with home to escape taxing demands for assistance. Anticipating this, the home community may set up subtle exit-barriers, which could lead to below-optimal levels of migration. Jakiela and Ozier (2012) report laboratory evidence from Kenya that women feel obliged to share four to eight per cent of the income gains realised in the experiment. In our sample, Tanzanian migrants sacrifice 2.5 to 8.1 percentage points out of a total growth of 119 per cent to insure their relatives. This estimate is equivalent to a 'tax' of 2 to 7 per cent. We regard this tax rate as too trivial to exert any constraining effect on migrants.

After describing the model, the data and the estimation strategy in the next three sections, we discuss the results in Section 5. Section 6 tackles the issue of longer-run transactional motives. In section 7 we calculate the cost of this insurance provision.

Section 8 discusses robustness of our results, while Section 9 provides the concluding discussion.

2. Risk sharing in theory

The full risk sharing hypothesis is based on the idea that the network acts as if it was a single household that maximised utility subject to a joint budget constraint. The model predicts that incomes are completely pooled (according to predetermined weights) and all idiosyncratic income shocks are smoothed through the network (e.g. Altonji et al., 1992; Townsend, 1994).

In a simple two-household extended family network, both households derive utility from consumption: $v(c)$. Insurance and credit markets are missing and income (y_s) is uncertain and depends on the state of the world (s). [2] We assume that households live infinitely. [3]

Assuming that households maximize a well behaving utility function [4], the standard utility maximization problem yields a following first order condition:

$$\frac{u'[c_1(y)]}{u'[c_2(y)]} = \lambda = \frac{\omega_2}{\omega_1}, \quad (1)$$

where λ, the Lagrange multiplier, is the marginal utility of income. According to equation (1), households equate their marginal utilities of consumption in all states of the world. The allocation depends on the Pareto weights ω_1 and ω_2 that are determined by the extended family.

[2] To simplify notation, we abstract away savings. This does not affect the main predictions of the model (see, for example, Ligon (1998) for a characterisation of the full risk sharing model with savings). However, the ability to save may exacerbate the efficiency problems if the key assumptions listed below do not hold (see Ligon, 1998; Chandrasekhar, Kinnan and Larreguy, 2012).

[3] If the time frame is finite, in the absence of altruism, households would not have any incentive for risk sharing in the final period, and as result in T-1, T-2, etc. The assumption of infinite time frame holds if the new household head inherits from the previous head and maintains the risk sharing contract with same households. See Fafchamps (1992) for an alternative justification for this assumption.

[4] The utility function is inter-temporally separable, strictly increasing but concave ($v' > 0$ & $v'' < 0$).

If utility functions follow a constant relative risk aversion function: $u(c) = \frac{c^{1-\psi}}{1-\psi}$, where ψ is a measure of risk aversion, the first order conditions for household i at time t become: $\omega_i c_{it}(y)^{-\psi} - \lambda = 0$. Equating these conditions for the two households, taking logarithms and re-arranging yields:

$$\Delta \ln c_1(y) = \Delta \ln c_2(y). \tag{2}$$

Equation (2) implies that if full risk sharing takes place, we should not expect to see households within the same extended family growing at different rates. Furthermore, assuming that there are no frictions between the households in the extended family, the model predicts that all idiosyncratic shocks experienced by households are completely smoothed through the extended family. These two predictions form the basis for our rejection of the full risk sharing model. First, the descriptive data confirms a highly unequal consumption growth between migrants and non-migrants. Second, in section 5 we show that after controlling for extended family fixed effects, household consumption growth remains responsive to idiosyncratic income shocks.

The rejection of full risk sharing is neither novel nor surprising and emerged as an empirically established stylized fact early on within this strand of literature being valid across a variety of different contexts (Altonji, Hayashi and Kotlikoff 1992; Townsend, 1994; Grimard, 1997). Most studies, however, find that at least some degree of insurance takes place and explain this theoretically by adding additional constraints (relating to the failure of assumptions regarding perfect information and full commitment) to the full risk sharing model. We will discuss each of these constraints in turn, but point out that an important common feature across all these augmented models is that, if the risk sharing contract survives, the ratio of marginal utilities become state contingent and are therefore no longer constant over time. This could allow the share of some members (migrants in our case) to increase over time.

In the presence of enforcement problems, the better-off households have an incentive to leave the arrangement and live in autarky. The limited commitment model (e.g. Coate and Ravallion, 1993; Attanasio and Ríos-Rull, 2000; Ligon, Thomas and

Worrall, 2002; Kinnan, 2012) appends the full risk sharing model with participation constraints (one for each household):

$$\sum_{t=1}^{\infty} \beta^t \sum_{s=1}^{S} \pi(y_s)\{v_1[c_{1t}(y_s)]\} \geq \mathrm{u}_A, \tag{3}$$

where u_A is the expected utility received in autarky. Solving the augmented maximization problem yields a following first-order condition:

$$\frac{u'[c_1(y)]}{u'[c_2(y)]} = \frac{\omega_2 + \sum_{s=1}^{S} \mu_2(y_s)}{\omega_1 + \sum_{s=1}^{S} \mu_1(y_s)} \ . \tag{4}$$

where μ_1 and μ_2 are the Lagrange multipliers attached to the participation constraints. Now, as can be seen from equation (4), if the participation constraints bind, the ratio of marginal utilities becomes state contingent. In the context of migration, a growth premium has to be granted to the migrant whose autarky options have improved. As a result, risk sharing is no longer efficient: the impact of idiosyncratic income shocks is not equally shared within the extended family network.

The other frictions have similar analytical consequences. If households cannot monitor other network members, the problem of free riding emerges. In moral hazard models (Lim and Townsend, 1998; Kinnan, 2012), the full-risk sharing model is augmented with incentive-compatibility constraints. The *ex ante* information asymmetry leaves the extended family to balance effort and insurance; migrants are motivated to exert effort by rewarding them with higher consumption. This comes with an efficiency cost: idiosyncratic shocks are not completely smoothed within the network. Finally, if there is imperfect information about the realised incomes, households may have an incentive to misreport their incomes to avoid payments or even claim transfers from other households. In hidden income models (Townsend, 1982; Fafchamps, 1992; Kinnan, 2012), the maximization problem is augmented with truth-telling constraints that require that households will not gain from misreporting. To encourage truthful reporting, migrants are allowed to enjoy a larger share of the consumption cake. As a consequence, Pareto-efficient risk sharing is again sacrificed.

These frictions can have important implications for the degree of risk sharing. [5] Distinguishing which of the three models of constrained insurance explain our data best is beyond the scope of this paper. [6] One common feature, however, is that despite friction, reciprocity remains intact: households engage in reciprocal risk sharing but the degree of its efficiency varies. In section 5, we study the existence of such reciprocal but partial risk sharing arrangements by testing whether households are responsive to income shocks faced by other households in the same extended family network.

In this paper, we contrast these reciprocity-based models with models that take into account social norms. Redistributive values may have been instilled since childhood and carefully nurtured through oral transmission, rituals and ceremonies in which the importance of the kinship group is strongly emphasised (Levi-Strauss, 1969). Remittances and other forms of assistance may also buy social prestige, political power or serve to perpetuate subordination (Platteau, 2010; Platteau and Sekeris, 2010). In risk sharing literature, social norms have been seen as the glue that keeps the risk sharing contract from breaking apart by alleviating enforcement and information problems (Stark and Lucas, 1988; Fafchamps, 1999; Foster and Rosenzweig, 2001). Theoretically this can be modelled as subjective satisfaction that individuals receive from participation. [7] The satisfaction can stem from the fulfilment of obligations and the avoidance of social sanctions, such as guilt, shame or ridicule, or fear of witchcraft. It can also include altruism, which we do not attempt to distinguish from social norms.

A recent empirical literature relying on experimental design highlights the importance of these forces. Chandrasekhar, Kinnan and Larreguy (2011, 2012) find that in the presence of hidden income and limited commitment, social proximity between the risk sharing partners increases the amounts transferred. The field experiments of Leider et al. (2009) and Ligon and Schechter (2012) show that altruism is more

[5] For example, Chandrasekhar, Kinnan and Larreguy (2011, 2012), using field experiments from Southern India find that limited commitment reduces transfers by 10 per cent and hidden income by 40 per cent.
[6] See Kinnan (2012) for such exercise with data from rural Thailand.
[7] In the context of limited commitment, we can re-write the right-hand side of equation (6) as $u_A - A$, where A captures such satisfaction (Fafchamps, 1999; Foster and Rosenzweig, 2001; De Weerdt and Fafchamps, 2011).

important than repeated interaction in determining the size of the transfer. Furthermore, social norms could weaken the constraints to risk sharing to the extent that they never bind and allow for the existence of sustained, unreciprocated transfers. [8] Below we will find evidence of such unilateral relations and argue that this is consistent with risk sharing motivated by social norms.

3. Data and descriptive analysis

Kagera is a region in the north-western part of Tanzania. A large part of Lake Victoria is contained within this region and it shares a border with Burundi, Rwanda, and Uganda. The region is overwhelmingly rural and agricultural production is the most important source of income, with more than 80 per cent of the region's economically active population engaged in it (URT, 2006). Bananas, beans, maize, and cassava comprise the main food crops while coffee, tea, and cotton are important cash crops. Recent years have seen a rise in improved banana varieties and sugar for use as cash crops. At the time of the last national census in 2002, Kagera had a population of roughly two million people.

The Kagera Health and Development Survey (KHDS) was originally designed and implemented by the World Bank and the Muhimbili University College of Health Sciences. It consisted of 915 households that were interviewed up to four times from autumn 1991 to January 1994.[9] The KHDS-2004 survey aimed to re-interview all individuals that were ever interviewed in the baseline survey and were alive in 2004. This effectively meant that the original household panel survey turned into a panel of individuals. A full household questionnaire was administered in a household where a panel respondent was found residing. Due to household dynamics, the sample size increased to more than 2,700 households.[10] The second KHDS follow-up was administered in 2010 with this time more than 3,300 households interviewed.[11]

Although KHDS is a panel of individuals and the definition of a household loses meaning after 10-19 years, it is common in panel surveys to consider re-contact rates

[8] Schechter and Yuskavage (2011) empirically document unreciprocated relations in Paraguay.
[9] See World Bank (2004).
[10] See Beegle, De Weerdt and Dercon (2006).
[11] See De Weerdt et al (2012).

in terms of households. Excluding households for which all previous members were deceased (17 households and 27 respondents), the KHDS 2004 field team managed to re-contact 93 per cent of the baseline households. In 2010, 92 per cent of the initial households were re-contacted. Taking into account the long, 10 or 16 year periods between surveys, the attrition rates in KHDS-2004 and KHDS-2010 are extremely low by standards of such panels (Alderman et al., 2001).

This paper exploits the fact that the survey includes all tracked split-offs from the original household and contains particularly rich information on the current links between them. The 2010 sample contains 3,314 households, originating from 816 initial households. The average baseline household spawned 4.1 households by 2010, out of which 2.4 were non-migrant and 1.7 were migrant households. Approximately three per cent of the initial households (99 households) did not have any split-offs. In what follows we will refer to these networks as extended family networks.

Figure 1 provides an overview of migration patterns. By 2010 nearly 45 per cent of the households were still residing in their original baseline community and 10 per cent were found in a neighbouring village. Covariate shocks, rainfall in particularly, are likely to be correlated between neighbouring villages. This correlation should decrease when the distance between the households increases. All else equal, risk sharing with households in the same community is less effective than risk sharing with households residing further away. Since we are mainly interested in seeing whether households exploit potential gains from spatial diversification, we define migrants as households that in 2010 are not located in the original village or in a nearby village but are found elsewhere within Kagera or outside Kagera. [12]

[Figure 1 here]

Remittances offer one medium for risk sharing between households. Table 1 provides a summary of the average remittance flows over the past 12 months in 2010 between the migrant households and households living in or near their baseline villages. While non-migrant households were net-receivers of remittances, Table 1 shows that transfers flow both ways. This could lead one to think – mistakenly as the analysis

[12] Our results are robust to alternative migrant definitions, such as also defining households that moved to a nearby village as migrant households.

below reveals – that these are relationships of reciprocal risk sharing. The data in Table 1 are self-reported and it is interesting to note that migrants claim to send more home than non-migrants acknowledge. A similar discrepancy does not exist in migrant-migrant or stayer-stayer dyads [13].

[Table 1 here]

Table 2 provides an overview of the reasons for leaving the baseline village. More than one third of the female respondents, but none of the male respondents cited marriage as the reason for migrating, which is what one would expect in a culture with patrilocal marriages. Less than 15 per cent of the female respondents reported that they left because of work. In contrast, almost 45 per cent of the male migrants reported to have moved because they had found work or went looking for one.

[Table 2 here]

The consumption data originate from extensive food and non-food consumption modules in the survey, carefully designed to maintain comparability across survey rounds and controlling for seasonality. The aggregates are temporally and spatially deflated using data from a price questionnaire included in the survey. Consumption is expressed in annual per capita terms using 2010 Tanzanian shillings. [14]

Table 3 provides the summary of the consumption and poverty developments of the panel respondents with respect to their 2010 location. On average, consumption levels in the sample almost doubled over 19 years. Individuals who stayed in their community saw their consumption increasing by more than 40 per cent. Consumption growth for migrants was much higher: those who left Kagera saw their consumption nearly triple over the same two decades. The poverty statistics tell the same story: nearly all respondents who left the region managed to escape poverty, while poverty reduction among non-migrants was more modest. These descriptive statistics reinforce the results reported in Beegle, De Weerdt and Dercon (2011): individuals who moved did considerably better than those who decided to stay.

[13] By dyad we refer to a pair of households.
[14] Using adult equivalent units as the denominator instead of household size produces almost identical results across all specifications.

[Table 3 here]

After moving, migrants remain linked to extended family members at home: 87 per cent of the migrants report that they communicated with a non-migrant network member in the 12 months preceding the survey. Migrants who maintained some form of communication experienced an average consumption growth of 123 per cent, while those who did not grew by 87 per cent. [15] This difference is statistically significant at the 1 per cent level. The severing of the most basic links does not seem to be associated with higher consumption growth; if anything, the reverse is true.

We use data from shock modules administered in 2004 and 2010. During both of these rounds, the panel respondents were asked to consider each year between the survey rounds and indicate whether a particular year was, in economic terms, 'Very good', 'Good', 'Normal', 'Bad', 'Very bad'. For each 'Very bad' response, the respondents were asked to provide the main reason for the hardship. We consider each 'Very bad' response as an economic shock. More than 60 per cent of the panel respondents reported experiencing at least one such shock between 1994 and 2009.

Table 4 provides an overview of the shocks experienced. Most frequently reported economic shocks were death of a family member, serious illness and poor harvest due to bad weather.

[Table 4 here]

The shock data were collected at individual level – in particular for each person on the 2010 roster who also appears on the original 1991/94 rosters. Since our focus is to examine the role of shocks on household consumption, the data had to be reformatted from individual to household level. [16] If at least one individual in the household reported to have experienced a shock, we interpret it as a household level shock. We should also exclude shocks that occurred before the households split. Fortunately, we

[15] The mean consumption growth among those who maintained contact was 441,212 TZS and among those who severed links 293,188 TZS.
[16] We repeated the complete analysis of the following sections using individual level data and find it does not change the conclusions.

know the year in which the respondents moved to their 2010 location, allowing us to include only shocks that occurred one year after this move. [17]

Furthermore, some of the shock categories are problematic to our network analysis. Mortality shocks may trigger inheritance flows within extended families. As such, a negative shock in one household may actually be a positive income shock in another household. A similar problem arises with the loss of remittance shocks, if these capture the loss of transfers from a household within the same extended family. We therefore exclude these two shock categories from our final shock variable.

Finally, there are 542 households that belong to a network that contains only non-migrants or only migrants. As our interest lies in the role of migration in risk sharing, we cannot use these households for empirical identification for risk sharing between migrant and non-migrant households. These households are therefore dropped from the final sample. Table 5 presents the summary statistics for the final sample of 2,246 households by 2010 migration status.

[Table 5 here]

4. Econometric strategy

We begin the econometric analysis by testing the full risk sharing hypothesis for those extended family networks that contain both migrants and non-migrants. The difference in logged per capita consumption between 2010 and the baseline ($\Delta \ln c_{ij}$) for household i in extended family j is formally modelled as:

$$\Delta \ln c_{ij} = \beta s_{ij} + x'_{ij}\gamma + \alpha_j + \varepsilon_i \quad (5)$$

where s_{ij} has a value 1 if the household experienced a shock in 1994-2009, or if a migrant household, after migrating to its current location. The term x_{ij} is a vector of household characteristics in 2010 capturing the characteristics of the previous

[17] This means that for households that remained in the baseline village we consider shocks that took place between 1994 and 2009. An alternative strategy would be to only use shocks that occurred after these household lived with *any* other network household member. Applying this strategy does not, however, change the conclusions.

household members [18] such as the number of previous household members in the 2010 household, the age of the oldest and the education (in years) of the most educated previous household members in the household. We also include dummies capturing their relation to the 2010 household head and their marital status. [19] The term α_j represents the network fixed effect and ε_i is the error term. The inclusion of the network fixed effects means that we compare the impact of shocks between the households originating from the same initial household. As such, the full risk sharing model presented earlier requires that $\beta=0$.

The rejection of the full risk sharing model using equation (5) implies either that the risk sharing arrangement is not efficient – or that the network does not engage in risk sharing at all. To explore for the existence of partial risk sharing, we assess whether household per capita consumption growth is responsive to shocks experienced by other households in the same extended family. This test builds on equation (5). We drop the network fixed effects and replace them with baseline village fixed effects (θ_v) and network characteristics (w_j) comprising the number of migrant and non-migrant households in the network and variables capturing characteristics of the initial household, such as its demographic composition, the household head's characteristics, including education, gender, age and the quadratic of age. We also include (logged) per capita consumption at the baseline ($\ln c_{j,1991}$). The network shock variable, z_{ij}, measures the number of households affected by an income shock. The household's own shocks are excluded from this variable. The partial risk sharing specification is formulated as:

$$\Delta \ln c_{ij} = \beta s_{ij} + \delta z_{ij} + x'_{ij}\gamma + w'_j\vartheta + \gamma \ln c_{j,1991} + \theta_v + \varepsilon_i \quad (6)$$

A negative and statistically significant δ would imply that some risk sharing takes place within the extended families.

[18] Previous household member refers to a person interviewed at the baseline in 1991/94.
[19] To address concerns about some of these 2010 household characteristics variables being potentially endogenous, we run all main regressions again, but drop each of these control variable in turn. We find the shock and network shock coefficients remain stable across all specifications.

We will assess the impact of these network shocks separately for migrant and non-migrant households and fully acknowledge that selection into migration is unlikely to be random. The differences in the observed level of risk sharing may be caused by migration or by some unobserved characteristics[20] that differ between migrant and non-migrant households, or by some combination of both. As a result, these regressions do not allow us to say whether migration is causally responsible for the migrant taking on the role of insuring sedentary extended family network members, or whether the effect is driven by unobservables. In particular, we cannot make any statements about what would have happened if migrants had stayed home or the home-stayers had migrated. It is possible that in this parallel universe roles would have switched (migration is causally responsible) or not (it is driven by the unobserved differences between migrants and non-migrants). For the purpose of this paper it is of no consequence whether the observed unilateral insurance provision from migrants to those who stayed at home is driven by migration itself, by some unobserved characteristic of the migrants or by some interaction of the two. Our interest lies in documenting the fact that migrants provide unilateral insurance to non-migrants, while at the same time shooting ahead of them in consumption terms.

Finally, the baseline per capita consumption variable in equation (6) raises a concern about endogeneity. The error term ε_i could be correlated, for example due to measurement error, with the lagged consumption variable. This would then bias the estimate measuring the impact of the lagged consumption but it may also affect other coefficients. Fortunately, we can think of a credible instrument that allows us to assess this possibility. Rainfall is one of the main inputs in agricultural production in Kagera and poor rainfall (i.e. droughts) can have serious consequences for incomes. Excess rains are less of a problem due to the focus of the production on tree crops and also because the terrain is relatively undulating. The region has two rainy seasons, a long rainy season usually between March and May and a short rainy season usually between October and December. The agricultural production takes place during these seasons. Therefore, we employ average monthly z-score deviations of rainfall during the two rainy seasons preceding the interview and truncate the positive rainfall

[20] For example, migrants may have different attitudes towards risk (de la Briere et al., 2002) or access to other risk-coping mechanisms, such as credit and savings (Yang and Choi, 2009).

deviations to zero. [21] Rainfall during the agricultural production is expected to influence consumption through income fluctuations but is unlikely to be correlated with the potential measurement error in the per capita consumption variable. The baseline village fixed effects (θ_v) in equation (6) wipe out the level effects of rainfall in the first stage regression. Therefore, exploiting the fact that rainfall shocks will affect different types of households in different ways, we interact the rainfall variable with head's gender, age and education yielding a total of three instruments.

5. Results

We begin by testing the full risk sharing model described above. Column 1 in Table 6 provides the results for the base specification of equation (5) with network fixed effects (NFE). [22] The control variables capture the characteristics of the previous household members, including their position within the 2010 household. The signs of the control variables are *a priori* correct. For example, education has a positive impact on consumption growth, while households with widowed or divorced previous household members experience lower consumption growth than others within the same extended family network.

The statistical significance of the shock coefficient, despite the inclusion of NFE, reveals that shocks are not insured within extended families. Households that experienced a shock had 15 percentage points lower consumption growth, on average and *ceteris paribus,* than households from the same extended family who did not experience a shock. The emergence of this wedge in the face of a shock implies a clear rejection of the full risk sharing model in the extended family networks in this study.

[Table 6 here]

In column 2 we drop the NFE and replace them with network characteristics, such as the number of migrant and non-migrant network members (which together control for network size and composition) and the wealth and demographics of the baseline

[21] Beegle, De Weerdt and Dercon (2008) employ a similar instrumental variable approach for their lagged consumption variable in assessing the long-term impact of adult deaths on consumption growth in Kagera.

household from which the network is formed. We also include baseline village fixed effects. The size of the shock coefficient is nearly identical to the one obtained with NFE, which gives confidence in the network level controls we use later in the analysis on reciprocal risk sharing.

Finally, column 3 provides the Two-Stage Least Squares results that address the potential endogeneity problem arising from the inclusion of the initial logged per capita consumption variable. The first stage regression results and the standard IV-diagnostic tests are presented in Table A1 in the Appendix. The included instruments show how households headed by older and more educated males enjoy higher baseline consumption. The excluded instruments are zero-truncated negative z-score deviations of rainfall interacted with the household head's age, education and gender. They show that the positive level effects of each of these three household head characteristics are attenuated with the inclusion of negative rainfall shocks. The Cragg-Donald (1993) test yields 15.84 indicating that our instruments are relevant. Comparison with the critical values provided in Stock and Yogo (2005) implies that the bias of our IV-estimate is less than five per cent of the OLS estimate. The Hansen (1982) J-test provides a p-value of 0.394. Thus, the null hypothesis of zero correlation between the instrument and the error term is upheld at conventional levels. The shock coefficient and the standard error from the 2SLS estimates are almost identical to those from OLS, indicating that the potential endogeneity of the logged per capita baseline consumption has a negligible influence on the shock variable. In the light of this, we use the more efficient OLS method to make inferences in the remainder of the text.

Next we test whether any risk sharing takes place in these networks. As discussed earlier, we replace the NFE with network characteristics and baseline village fixed effects and augment the specification with the network shock variable. The first column in Table 7 reports the results for the migrant households and the second column for the non-migrant households. For migrants, the network shock coefficient is negative and highly significant. These network shocks have a sizeable impact on migrant households' consumption: on average, a shock in one household in the network resulted in a drop of six percentage points in consumption growth. As shocks are not correlated within the extended family networks (the intra-class correlation

coefficient equals 0.012 with a standard error of 0.016), this finding reveals that migrants insure other households in their extended families. Non-migrant households, on the other hand, do not appear to be affected by the network shocks. The point estimate is nearly zero and insignificant. These results suggest that the risk sharing arrangement is not reciprocal.

[Table 7 here]

In order to investigate this further, we decompose the network shock variable into shocks in non-migrant and migrant households. The first variable measures the number of non-migrant households that experienced a shock in the extended family. The second network shock variable measures the number of migrant households affected by shocks. As before, household's own shocks have been excluded from these variables. Table 8 presents the regression results. Migrants are susceptible to shocks affecting other migrant and non-migrant households within their extended family network, while non-migrants are sensitive to neither. On average, a shock in one non-migrant household in the network leads to a drop of five percentage points in migrant household's consumption growth. Shocks in other migrant households have a larger negative effect on migrant's consumption than shocks experienced in stayer households. However, this difference is not statistically significant from zero at a conventional level.

[Table 8 here]

We conclude that migrant households are partially and unilaterally insuring households that stay behind. This lack of reciprocity violates the predictions of the reciprocity-based models (without a social norms term). Because, on average, migrants are twice as rich as those who remained at home, these findings are consistent with reciprocity-based models augmented with a social norms term, which attenuates the participation, truth-telling or incentive compatibility constraints.

6. Other transactional insurance motives

An alternative explanation to the observed lack of reciprocity could be that migrants insure non-migrants in exchange for other benefits. By concentrating on consumption differences we have considered only current pay-offs from any risk sharing

arrangement. It is quite possible that the benefits are still to accrue to the migrant in the more distant future. Lucas and Stark (1985) mention that there could be exchange motives for insurance provision relating to the desire for non-migrants to look after local assets, the intention to return home and the aspiration to inherit. In a context that lacks technology to allow future income to be consumed now, we could confuse unilateral insurance with postponed reciprocity. Fortunately, the KHDS questionnaire is particularly rich and we are thus able to explore some of these issues.

The questionnaire asks each migrant about asset holdings in the baseline village. As our outcome variable is consumption growth we cannot use these asset holdings as explanatory variables: current wealth is surely endogenous to growth in wealth. We attempt to circumvent this problem by looking at the share of assets in the current portfolio that are located in the village. While it remains possible that portfolio composition is endogenous to consumption growth, we believe the results are informative enough to report.

About 12 per cent of migrants have assets in the baseline village and 10 per cent of migrants own land in the baseline village. For land we have exact area measurements, but not monetary values. If migrants engage in risk sharing with those who remain at home for the purpose of maintaining land and ensuring their continued entitlement to the land (which is important in a country with few formal land deeds), then we would expect more responsiveness to network shocks from people with a larger share of their land holdings in the baseline village. The first column of Table 9 explores this. As before, the dependent variable is logged per capita consumption growth. We interact the non-migrant network shocks with a variable measuring the share of the land in the baseline village. The coefficient on this interacted variable turns out insignificant implying that the share of land in the baseline village neither increases nor decreases the insurance provision.

In the second column in Table 9 we interact the non-migrant network shock variable with the length of the migration spell. Following Dustmann and Mestres (2010), we argue this to be a measure of the permanence and success of the move and an inverse measure of the return likelihood. We find that the duration of the migration spell does

not have any impact on migrant's insurance provision. This also holds when we use non-linear versions of the migration duration in the form of a piecewise linear spline.

The third column in Table 9 investigates whether the expectation to inherit is a plausible motive for unilateral insurance. Nearly 40 per cent of the migrant households have parental clan land holdings waiting for them in the baseline village. By interacting the non-migrant network shocks with a parental clan land holdings dummy, we find that that these households are no more (or less) engaged in insurance provision than households that do not expect to inherit land.

[Table 9 here]

A final transactional motive that could be consistent with the regression results is that non-migrants pay insurance premiums to migrants in return for their continued insurance provision. This does not seem consistent with the findings of Table 1, where we noted that migrants are net senders of transfers.

7. Is there a kin tax?

Is the price to the migrant for providing this unilateral insurance significant? From Table 8 we observe that for each shock in the extended family network at home there is a drop of 4.7 percentage points in the migrant's consumption, which appears to be a permanent deviation from the growth curve. The average migrant has 0.54 network shocks of non-migrants, resulting in an implied overall consumption growth penalty of 2.5 percentage points, on average, over the 19-year period. Over this same period, the average consumption growth among migrants was 119 per cent, implying that insurance constituted an average annual growth penalty of around 0.063 of one percentage point (reducing average annual growth roughly, from 4.206 per cent to 4.143 per cent). [23] Put another way, migrants share about 2.1 per cent of their very substantial growth by insuring family members at their original location. [24]

This is a lower-bound estimate because we cannot exclude the possibility that we are only measuring a subset of relevant shocks: if shocks are self-reported then

[23] We use geometric (rather than arithmetic) means to calculate the average annual growth rates.
[24] The 95%-confidence interval for the annual growth penalty is [0.008, 0.118] and for the 'kin-tax' [0.25, 3.99].

respondents may fail to mention those that were effectively insured. Fortunately, the survey provides an alternative shock measure, which is not self-reported. We have historical rainfall data from the Tanzanian Meteorological Agency for gauges in 9 weather stations across the Kagera Region. Each baseline village can be linked to its closest rainfall station as both databases contain record GPS location. The mean distance to a rainfall station is 19 km and the median is 15 km. For each village we can calculate average monthly z-score deviations of rainfall during the two rainy seasons, in relation to the 30 year average (1980-2010) for that village. Rainfall shocks are then constructed by truncating the positive yearly average rainfall deviations to zero. We calculate a stayer's own shock as the most negative shock in the 1994-2009 period. The first two columns in Table 10 show that rainfall shocks are important in determining consumption growth, with every standard deviation decrease in (negative) rainfall deviation causing consumption growth to decline by 25 to 30 percentage points, depending on the specification.

Knowing that rainfall shocks drive the incomes of the stayer households, we can use them as an alternative shock indicator. To do this we run a slightly modified version of Equation (6). First, we replace the cluster fixed effects with district fixed effects as the former are perfectly collinear with baseline village rainfall shocks for stayers. Secondly, we drop the own shock variable as we do not have rainfall data available for migrants. Column 3 repeats the results from Table 8 for comparison purposes, column 4 omits the own shock variable, while column 5 further replaces the cluster fixed effects with district fixed effects. As we drop these controls the coefficient become more negative. The last column uses a variable indicating the worst rainfall shock among stayer network members and looks at its effect on migrants. As expected, once we use rainfall shocks the kin tax goes up to -6.74%, with the 95 per cent confidence interval ranging from 2.83 to 14.31. We did run the regression with a control for self-reported own shock and found the coefficient goes down slightly consistent with the results in columns (3) to (5), where the controls serve to dampen the effect. We therefore consider the parsimonious rainfall specification from Table 10 the upper bound effect, binding the estimate between 2.1 and 6.7 per cent.

It is interesting to point out that our implicit tax rate of 2 to 7 per cent is of the same order of magnitude as that found in two other studies. Jakiela and Ozier (2012)

estimate that women in a laboratory setting in Kenya acted as if they were expecting to be pressured to share four per cent of their experiment winnings with relatives that were not present at the experiment. Ambler (2012) reports that El Salvadorian migrants living around Washington DC remit 5 per cent more of a windfall income if they know the potential recipients at home will be informed about it. Obviously in neither case is a one-to-one comparison of the results is appropriate, most importantly because the above studies focus on the effect of providing full information about the windfall, while our figures are based on the actual belief sets.

8. Discussion on the robustness

We conducted an array of robustness checks to verify our findings. [25] First, we find that the results are robust to an alternative migrant definition where also households that moved to a nearby village are defined as migrants.

Second, the results are not driven by the configuration of the data. The shock data were initially defined at individual level while our outcome variable is measured at household level. Conducting the empirical analysis at individual level does not affect our main findings.

Third, defining household consumption per adult equivalent instead of per household member yields close to identical results in all specifications.

Fourth, changing the way we isolate the shocks that occurred before the households split does not change our results either.

Fifth, we also checked whether the potential endogeneity of some our control variables is driving our results. Instrumenting the lagged consumption variable does not affect the shock coefficient. In addition, when the 2010 household level control variables are omitted one-by-one, the estimated shock coefficients remain stable across all specifications.

[25] The results of these robustness checks are available upon request.

9. Conclusions

Starting from the household rosters of a representative household survey conducted nearly two decades ago in Kagera, we find that over half of the original household members had moved internally, while very few moved internationally. Internal migrants provide unilateral insurance to those who remain at home, which seems to be driven by social norms rather than exchange motives. The total, final, long-run effect of this insurance provision on the migrant's growth amounts to a 2 to 7 per cent sacrifice in consumption (2.5 to 8.1 percentage points off the 119 % total growth realised by the migrant).

While our study cannot conclusively say where migrants would be without their extended family networks back home, a tax rate of 2 to 7 per cent seems too low to be an important brake on a migrant's growth. We do know that migrants who have severed links with home perform worse than other migrants and one should not overlook the fact that, while starting out from similar baseline welfare levels, migrants realise a total consumption growth which is three times higher than that of non-migrants.

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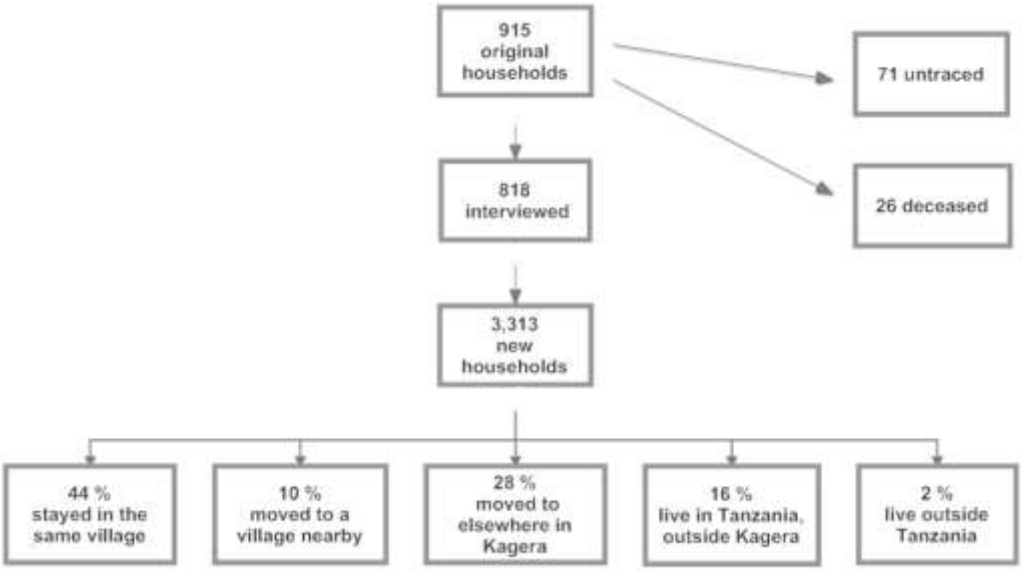
Figures

Figure 1: KHDS-2010 – Re-contacting after 16+ years



Tables

Table 1: Reported remittances in and out between migrants and non-migrants, as reported by the first half of the dyad with respect to transfers to (first column) or from (second column) the second half of the dyad. [26]

dyad	gifts out	gifts in	net in
stayer-migrant	9,914	14,440	4,526
migrant-stayer	20,439	10,392	-10,047
stayer-stayer	8,744	10,022	1,278
migrant-migrant	18,738	17,281	-1,458
Total	**13,208**	**12,328**	**-880**

Table 2: Reasons for leaving the baseline village

Reason	males (%)	females (%)
To look for work	29.8	7.5
Own schooling	16.0	10.3
Found work	15.1	6.7
To live in a healthier environment	10.4	11.7
Marriage	0.0	38.9
Other reason	28.8	24.9
Total	**100.0**	**100.0**

[26] This table is based on self-reported remittance flows in households in the past 12 months in 2010.

Table 3: Consumption and poverty movements of the panel respondents in 1991-2010 by 2010 location[27]

	mean 91	mean 2010	difference in means	N
Consumption per capita (TZS) by 2010 location				
Within community	343,718	492,398	148,680***	2,224
Nearby community	364,099	569,438	205,339***	382
Elsewhere in Kagera	357,930	695,951	338,021***	1,007
Out of Kagera	389,379	1,110,827	721,449***	658
Full Sample	355,926	642,558	286,632***	4,271
Consumption Poverty Head Count (%) by 2010 location				
Within community	31	19	-13***	2,224
Nearby community	30	20	-10***	382
Elsewhere in Kagera	31	16	-15***	1,007
Out of Kagera	23	3	-21***	658
Full Sample	30	16	-14***	4,271

*Note: Significance of the difference in means using a paired t-test, *** p<0.01, ** p<0.05, * p<0.1.*

Table 4: Shocks reported by the panel respondents 1994-2009

Type of shock	Freq.	Percentage
Death of family member	797	26%
Poor harvest due to adverse weather	638	21%
Serious illness	577	19%
Loss in wage employment	219	7%
Loss of assets	205	7%
Eviction/resettlement	99	3%
Poor harvest due to pests or crop diseases	98	3%
Low crop prices	85	3%
Loss in off-farm employment	78	3%
Low income due to lower remittances	43	1%
Loss of livestock	6	0.2%
Loss of Gifts and support by organisations	4	0.13%
Other reasons	172	6%
Total	**3,021**	**100%**

[27] All consumption values are in annual per capita terms and expressed in 2010 Tanzanian shillings.

Table 5: Descriptive statistics

	Migrant households		Non-migrant households	
	mean	**std. dev.**	**mean**	**std. dev.**
1991 household per capita consumption	355,539	184,744	354,120	200,839
2010 household per capita consumption	777,748	642,810	511,348	410,482
Per capita consumption growth between 1991-2010	422,210	653,545	157,228	394,930
Natural log of per cap. consumption growth between 1991-2010	0.6249	0.798	0.2908	0.631
Own shock	0.2009	0.401	0.4816	0.500
# of hhs that reported a shock in the network	0.7349	1.063	1.3621	1.336
2010 household characteristics:				
Age of oldest PHHM in the 2010 hh	31.6	11.378	42.3	17.705
A PHHM is head of this 2010 hh	0.4484	0.498	0.7276	0.445
A PHHM is spouse of this 2010 hh's head	0.4121	0.492	0.3390	0.474
A PHHM is child of this 2010 hh's head	0.0512	0.220	0.1913	0.393
Divorced PHHM in 2010 hh	0.0391	0.194	0.0598	0.237
A widowed PHHM in 2010 hh	0.0456	0.209	0.1725	0.378
A married PHHM in 2010 hh	0.6316	0.483	0.6892	0.463
Max yrs edu of PHHM in this 2010 hh	6.8335	3.190	6.2203	2.960
Number of PHHMs in this 2010 hh	1.1107	0.438	1.4987	0.979
HH size in 2010 hh	4.3591	2.472	4.9035	2.365
HH size in adult eq. units in 2010 hh	3.4625	1.965	3.8656	1.893
Initial household characteristics:				
HH total per capita consumption	355,539	184,744	354,120	200,839
Natural log value of assets in 1991	13.6922	1.095	13.7340	1.073
Education of hh head	4.3135	3.168	4.2383	3.071
Head was male	0.7553	0.430	0.7720	0.420
Age of hh head	49.1	15.9	49.0	15.5
Males 0-5 years	0.7684	0.908	0.7438	0.874
Males 6-15 years	1.3888	1.223	1.3373	1.138
Males 16-60 years	1.3516	1.023	1.3988	1.051
Males 61+ years	0.1991	0.399	0.2058	0.404
Females 0-5 years	0.8474	0.970	0.7839	0.881
Females 6-15 years	1.4493	1.347	1.4261	1.240
Females 16-60 years	1.9088	1.351	1.8215	1.200
Females 61+ years	0.2344	0.462	0.1998	0.417
HH had a non-earth floor in 1991	0.1926	0.394	0.1605	0.367
Observations	**1075**		**1171**	

Note: 1) PHHM refers to previous household member (i.e. person interviewed at the baseline).
2) All consumption values are in annual per capita terms and expressed in 2010 Tanzanian shillings.

Table 6: The effect of shocks on consumption growth

Dependent variable: (logged) per capita consumption growth	1	2	3
	OLS, NFE	**OLS**	**2SLS**
Own shock	-0.154***	-0.157***	-0.156***
	(0.027)	(0.025)	(0.025)
2010 household characteristics:			
Age of oldest PHHM in the household	-0.001	0.000	-0.000
	(0.001)	(0.001)	(0.001)
A PHHM is head of the household	0.175***	0.166***	0.166***
	(0.048)	(0.035)	(0.042)
A PHHM is spouse of the household head	0.111**	0.087**	0.083*
	(0.049)	(0.043)	(0.045)
A PHHM is child of the household head	-0.225***	-0.181***	-0.181***
	(0.052)	(0.043)	(0.047)
A divorced PHHM in the household	-0.329***	-0.299***	-0.293***
	(0.069)	(0.074)	(0.063)
A widowed PHHM in the household	-0.340***	-0.341***	-0.337***
	(0.058)	(0.052)	(0.052)
A married PHHM in the household	-0.488***	-0.463***	-0.459***
	(0.040)	(0.040)	(0.038)
Max years of education of PHHM in the household	0.061***	0.065***	0.064***
	(0.006)	(0.005)	(0.006)
Number of PHHMs in the household	-0.004	-0.024	-0.021
	(0.023)	(0.020)	(0.024)
Network characteristics:			
Number of split-off households stayed		-0.043***	-0.045***
		(0.010)	(0.010)
Number of split-off households moved		-0.006	-0.007
		(0.012)	(0.011)
Household characteristics at the baseline:			
Natural log value of assets in 1991		0.015	0.007
		(0.018)	(0.024)
Education of 1991 household head		0.002	0.000
		(0.006)	(0.006)
Head was male in 1991		-0.070**	-0.079**
		(0.035)	(0.044)
Age of household head in 1991		0.006	0.006
		(0.006)	(0.004)

Table 6: The effect of shocks on consumption growth

Dependent variable: (logged) per capita consumption growth	1	2	3
	OLS, NFE	**OLS**	**2SLS**
Age of head squared		-0.000	-0.000*
		(0.000)	(0.000)
Number of males 0-5 years in the household		-0.004	0.004
		(0.017)	(0.022)
Number of males 6-15 years in the household		0.047***	0.051***
		(0.012)	(0.015)
Number of males 16-60 years in the household		-0.002	-0.004
		(0.016)	(0.014)
Number of males 61+ years in the household		0.139***	0.156***
		(0.048)	(0.065)
Number of females 0-5 years in the household		0.005	0.009
		(0.017)	(0.018)
Number of females 6-15 years in the household		0.030**	0.032***
		(0.013)	(0.013)
Number of females 16-60 years in the household		0.002	0.003
		(0.014)	(0.013)
Number of females 61+ years in the household		0.015	0.019
		(0.032)	(0.033)
Household had a non-earth floor in 1991		-0.005	-0.032
		(0.046)	(0.068)
(logged) household per capita consumption in 1991 (in 2010 TZS)		-0.902***	-0.797***
		(0.041)	(0.218)
Number of observations	2,246	2,246	2,246
R^2	0.215	0.417	0.414
Adjusted R^2	0.212	0.410	0.393

*note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors are in parenthesis. Regressions in column 1 includes NFE, regressions in columns 2 and 3 include baseline village fixed effects. PHHM refers to previous household member (i.e. person interviewed at the baseline).*

Table 7: The effect of network shocks on consumption growth

Dependent variable: (logged) per capita consumption growth	**Migrant households** **1** **OLS**	**Non-migrant households** **2** **OLS**
Number of households that experienced a shock in the network	-0.061***	-0.018
	(0.014)	(0.017)
Own shock	-0.090**	-0.085***
	(0.044)	(0.027)
Number of split-off hhs stayed	-0.019	-0.023**
	(0.018)	(0.010)
Number of split-off hhs moved	0.002	-0.031*
	(0.011)	(0.017)
Age of oldest PHHM in the 2010 hh	0.003	0.000
	(0.002)	(0.001)
A PHHM is head of this 2010 hh	0.168***	0.180***
	(0.065)	(0.060)
A PHHM is spouse of this 2010 hh's head	0.015	0.107*
	(0.073)	(0.057)
A PHHM is child of this 2010 hh's head	-0.410***	0.008
	(0.119)	(0.052)
A Divorced PHHM in 2010 hh	-0.380***	-0.168**
	(0.122)	(0.076)
A widowed PHHM in 2010 hh	-0.313**	-0.244***
	(0.124)	(0.057)
A married PHHM in 2010 hh	-0.442***	-0.303***
	(0.056)	(0.048)
Max years of education of PHHM in this 2010 hh	0.080***	0.034***
	(0.007)	(0.007)
Number of PHHMs in this 2010 hh	0.033	-0.057**
	(0.045)	(0.023)
(logged) hh per capita consumption in 1991 (in 2010 TZS)	-1.029***	-0.801***
	(0.049)	(0.051)
Number of observations	1,075	1,171
R^2	0.487	0.346
Adjusted R^2	0.473	0.330

*note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors are in parenthesis.*
Regressions include baseline village fixed effects and variables controlling for household characteristics at the baseline. PHHM refers to previous household member (i.e. person interviewed at the baseline).

Table 8: Network shocks in migrant and non-migrant households

Dependent variable: (logged) per capita consumption growth	**Migrant households**	**Non-migrant households**
	1	**2**
	OLS	**OLS**
Number of non-migrant hhs that experienced a shock in the network	-0.047**	-0.031
	(0.021)	(0.022)
Number of migrant hhs that experienced a shock in the network	-0.095***	0.002
	(0.034)	(0.026)
Own shock	-0.090**	-0.086***
	(0.044)	(0.028)
Number of split-off hhs stayed	-0.021	-0.020
	(0.018)	(0.013)
Number of split-off hhs moved	0.005	-0.035*
	(0.012)	(0.019)
Number of observations	1,075	1,171
R^2	0.487	0.347
Adjusted R^2	0.473	0.331

*note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors are in parenthesis.*
Regressions include baseline village fixed effects, 2010 household level variables capturing characteristics of the previous household members and variables controlling for household characteristics at the baseline.

Table 9: Other transactional insurance motives

	Migrant households		
Dependent variable: (logged) per capita consumption growth	**1**	**2**	**3**
Number of non-migrant hhs that experienced a shock in the network	-0.039*	-0.019	-0.063*
	(0.023)	(0.068)	(0.036)
--- Interacted with:			
* Share of land in BLV in total land portfolio	-0.038		
	(0.062)		
* Number of years since the last PHHM migrated into this hh		-0.003	
		(0.005)	
* Hh has inheritable land in the baseline village			0.031
			(0.047)
* Hh member's parent lives in BLV			-0.011
			(0.052)
Own shock	-0.081**	-0.108**	-0.090**
	(0.040)	(0.044)	(0.044)
Share of land in BLV in total land portfolio	0.275***		
	(0.074)		
Household does not own land	0.236***		
	(0.039)		
Number of years since the last PHHM migrated into this hh		0.002	
		(0.004)	
Hh has inheritable land in the baseline village			0.011
			(0.073)
Hh member's parent lives in BLV			0.034
			(0.062)
Hh owns land			-0.147***
			(0.038)
Number of split-off hhs stayed	-0.020	-0.020	-0.019
	(0.016)	(0.018)	(0.018)
Number of split-off hhs moved	-0.009	-0.007	-0.007
	(0.012)	(0.012)	(0.012)
Number of observations	1,075	1,075	1,075
R^2	0.507	0.485	0.492
Adjusted R^2	0.492	0.470	0.476

*note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors are in parenthesis.*
Regressions include baseline village fixed effects, 2010 household level variables capturing characteristics of the previous household members and variables controlling for the household characteristics at the baseline. BLV refers to baseline village, NW to network, PHHM to previous household member (i.e. person interviewed at the baseline).

Table 10: Re-calculating the kin-tax through rainfall data

	stayers		Migrants			
	1	**2**	**3**	**4**	**5**	**6**
max rain shock 1994-2009	0.304***	0.246*				
	(0.067)	(0.126)				
number of stayer hhs in NW affected by shock			-0.047**	-0.059***	-0.062**	
			(0.020)	(0.021)	(0.024)	
max rainfall shock in stayer households in the network						0.188***
						(0.049)
Number of observations	1,171	1,171	1,075	1,075	1,075	1,075
R2	0.316	0.336	0.487	0.486	0.487	0.487
Adjusted R2	0.300	0.320	0.473	0.472	0.474	0.474
own shocks	n/a	n/a	yes	no	no	no
baseline village fixed effects?	no	no	yes	yes	no	no
baseline district fixed effects?	no	yes	no	no	yes	yes

*note: *** p<0.01, ** p<0.05, * p<0.1*

robust standard errors are in parenthesis, regressions include 2010 household level variables capturing characteristics of the previous household members and variables controlling for household characteristics at the baseline.

Table A1: First-stage regression results of Column 3 in Table 6

Dependent variable: (logged) hh per capita consumption in 1991	
Included instruments:	
Own shock	-0.015
	(0.017)
Number of split-off hhs stayed	0.022***
	(0.005)
Number of split-off hhs moved	0.009
	(0.006)
Age of oldest PHHM in the 2010 hh	0.002**
	(0.001)
A PHHM is head of this 2010 hh	0.000
	(0.025)
A PHHM is spouse of this 2010 hh's head	0.034
	(0.028)
A PHHM is child of this 2010 hh's head	0.003
	(0.033)
A divorced PHHM in 2010 hh	-0.065
	(0.041)
A widowed PHHM in 2010 hh	-0.028
	(0.034)
A married PHHM in 2010 hh	-0.043*
	(0.023)
Max years of education of PHHM in this 2010 hh	0.013***
	(0.003)
Number of PHHMs in this 2010 hh	-0.029*
	(0.015)
Natural log value of assets in 1991	0.077***
	(0.011)
Education of hh head in 1991	0.024***
	(0.006)
Head was male in 1991	0.173***
	(0.042)
Age of hh head in 1991	0.005*
	(0.003)
Age of head squared	0.000
	(0.000)
Males 0-5 years in 1991	-0.075***
	(0.009)

Table A1: First-stage regression results of Column 3 in Table 6

Males 6-15 years in 1991	-0.041***
	(0.007)
Males 16-60 years in 1991	0.018**
	(0.008)
Males 61+ years in 1991	-0.145***
	(0.035)
Females 0-5 years in 1991	-0.039***
	(0.010)
Females 6-15 years in 1991	-0.025***
	(0.007)
Females 16-60 years in 1991	-0.015**
	(0.007)
Females 61+ years in 1991	-0.032
	(0.021)
Hh had a non-earth floor in 1991	0.259***
	(0.024)
Excluded instruments:	
(Negative rainfall deviation) * (Age of hh head in 1991)	0.005**
	(0.002)
(Negative rainfall deviation) * (Education of hh head in 1991)	0.035**
	(0.015)
(Negative rainfall deviation) * (Head was male in 1991)	0.294***
	(0.108)
Number of observations	2,246
R^2	0.229
Adjusted R^2	0.201
Under-identification test:	
Kleibergen-Paap rk LM statistic	49.02
p-value	0.000
Weak identification tests:	
Cragg-Donald Wald F Statistic	15.84
Kleibergen-Paap rk Wald F statistic	19.36
Over-identification test:	
Hansen-J statistic	1.862
p-value	0.394

*note: *** p<0.01, ** p<0.05, * p<0.1. Robust standard errors are in parenthesis. Regression includes baseline village fixed effects. PHHM refers to previous household member (i.e. person interviewed at the baseline).*