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Abstract

This paper analyzes the impact of rural-to-urban migration on income inequality and gender wage gap in source regions using a newly constructed panel dataset for around 100 villages over a ten-year period from 1997 to 2006 in China. Since income inequality is time-persisting, we use a system GMM framework to control for the lagged income inequality, in which contemporary emigration is also validly instrumented. We found a Kuznets (inverse U-shaped) pattern between migration and income inequality in the sending communities. Specifically, contemporary emigration increases income inequality, while lagged emigration has strong income inequality-reducing effect in the sending villages. A 50-percent increase in the lagged emigration rate translates into one-sixth to one-seventh standard deviation reduction in inequality. These effects are robust to the different specifications and different measures of inequality. More interestingly, the estimated relationship between emigration and the gender wage gap also has an inverse U-shaped pattern. Emigration tends to increase the gender wage gap initially, and then tends to decrease it in the sending villages.

JEL classification: O15; J61; D31; C33

Keywords: Internal Migration; Inequality; System GMM.

1. Introduction

Inequality is closely and reciprocally intertwined with migration. On the one hand, income inequality between source and destination areas is widely believed to be one of the most important factors that drive economic migration. This is inherent in the Lewis model of dual economy and made more explicit in the Harris-Todaro model of rural–urban migration (J. R. Harris and M. P. Todaro, 1970; W. A. Lewis, 1954). Internal rural–urban migration is modeled as a response to the wage disparities between the urban and rural sectors (D. Ray, 1998).[1] By the same token, international migration can be viewed as an outcome of global inequality (R. Black et al., 2005). Large and increasing wage gaps across countries are cited as an irresistible force impelling greater labor mobility across national boundaries (L. Pritchett, 2006).

The growth of the literature on the New Economics of Labour Migration has brought about focus on the inequality within sending communities as drivers of migration: the household's relative position with respect to a specific reference group in addition to the household's absolute income serves as strong motivation for migration (O. Stark and J. E. Taylor, 1991, Oded Stark and David E. Bloom, 1985). Conversely, labor mobility generates a feedback effect on income inequality in both sending and receiving communities. In fact, one of the controversial arguments put forward to control international migration is that immigrants increase income inequality in the receiving countries because the influx of low-skilled immigrants suppresses the wage of the locals at the end of the skill distribution. However, more recent items of evidence have suggested that only 5-10 percent of wage inequalities can be attributed to immigration in the United States between 1980 and 2000, much smaller than is commonly presumed (David Card, 2009; C. D. Goldin and L. F. Katz, 2008).

This paper examines the impact of migration on inequality in the sending communities. The evidence presented is mixed. It depends on the structural factors that affect the distribution of the costs and benefits of migration and the associated selectivity of migration itself (R. Black, C. Natali and J. Skinner, 2005). If the costs of migration are sizeable and the poor face binding credit constraints, which is often the case in the context of developing countries, migration will

[1] Needless to say, the inequality among the other dimensions of life opportunities, such as education, health, civil liberty, and other dimensions are also drivers of migration.

be positively selected by the poor trapped in nonproductive activities at the source communities. This means that the rich will benefit the most from migration because migration will widen rather than narrow the income gaps in the sending communities. This is indeed what earlier research undertakings have found (R. H. Adams, 1993, 1998; M. Lipton, 1980; O. Stark et al., 1988).

Recent empirical works have challenged this traditional view. It is increasingly recognized that the impacts of migration on income inequality in the migrant sending communities crucially hinge on the selectivity of migration and how this selectivity changes over time. Although pioneer migrants may come from relatively wealthy households who can afford the cost of migration and have better information on urban employment, their migration is likely to induce more migration from people in the bottom of the income distribution for two reasons. First, the increase in relative deprivation among nonmigrants tends to boost their desire to migrate. Second, the establishment of migrant networks in the destination areas lowers the costs and risks of migration, which in turn facilitates more waves of migration of the poor. As a result, the initially negative effect of remittances on income equality might therefore be dampened or even reversed (H. de Haas, 2007; David McKenzie and Hillel Rapoport, 2007; Hillel Rapoport and Frédéric Docquier, 2005;O. Stark, J. E. Taylor and S. Yitzhaki, 1988).[2]

The difference in the results may also come from the differences in the methodology. Earlier studies treated remittance income as an exogenous transfer, and compared Gini coefficients with and without the inclusion of remittance income, whereas more recently, remittances have been treated as a potential substitute for home earnings. In addition, the observed income distribution with remittances are compared to a counterfactual scenario in

[2] Docquier et al. (2006), propose a dynamic theoretical framework that goes part of the way towards reconciling the conflicting results from empirical studies and complements the "networks" view in showing that the same predictions may be obtained with exogenous (i.e., constant) migration costs. They investigate the impact of migration on income and inequality both via the direct effect of migrants' households increase in their income via higher wages abroad and also the indirect effects of the outbound flow of individuals on the local labor market. They do so in a way that demonstrates the importance of the pre-migration distribution of wealth in determining the impact of migration on the dynamic path and long-run levels of income and wealth inequality. They show that migration and remittances always lower wealth inequality. In contrast, although income inequality is also reduced in the long run, it may either increase or decrease in the short run, depending on the initial distribution of endowments. That is to say, the short- and the long-run effects on the income distribution may be of opposite signs and display an inverse U-shaped relationship.

which no migration takes place, but includes an imputed level of home earnings (R. Black, C. Natali and J. Skinner, 2005; Nong Zhu and Xuebei Luo, 2008). Although the earlier approach is unrealistic in assuming that remittance-earning migrants are separate entities from their households in the rural areas and excludes their income in destination areas, the improvement that the counterfactual model provides is limited because the selection into migration is difficult, if not impossible, to model.[3] This led some researchers, after reviewing the literature, to conclude that any overarching generalization about impacts on inequality is unlikely to be robust (R. Black, C. Natali and J. Skinner, 2005).

In addition to these methodology challenges, the usefulness of earlier literature is also tempered by the cross-section nature and the small sample size. The lack of panel data at the community level seriously limits the researchers' ability to quantify the temporal dimension of migration and inequality. The alleged inequality-reducing effect of migration over time remains elusive. Moreover, it is primarily based on anecdotes rather than evidence. Without a large sample of communities, previous literature focuses mainly on the examination of the effect of migration on inequality in only a couple of communities (see McKenzie and Rapoport 2007 for a detailed review). The external validity of these studies is questionable, at best. McKenzie and Rapoport (2007) contribute to the literature by constructing cross-section data of 57 rural communities from the Mexican Migration Project (MMP) survey and short panel data of 214 rural communities in Mexico from the national demographic dynamics survey (ENADID) in 1992 and 1997. This identification strategy essentially estimates the effect of the development on emigration on the change in inequality. Community-level emigration rates are instrumented by the historic state-level migration rates and U.S. labor market conditions to deal with the endogeneity of migration.

McKenzie and Rapoport (2007) find that further migration reduces inequality among communities with reasonably high initial levels of migration experience. Furthermore, migration

[3] In earlier studies of this strand for example, an econometric model does not control for the selection problem involved in the original migration decision. The migrant and non-migrant were treated as drawn randomly from the population (Barham and Boucher, 1998). In Barham and Boucher (1998), they estimate individual earnings equations in a double-selection model involving migration choices and non-migrants' labor force participation decisions. However, these kinds of selection models require at least one additional exogenous instrument for each selection. Otherwise the identification will only come from non-linearity.

has positive but insignificant effects on inequality in communities with smaller migration networks. Employing the panel data for a sample of communities observed in 1992 and 1997, they find suggestive evidence for an inverse U-shaped Kuznets relationship between migration and inequality, with migration increasing inequality at first before subsequent migrations lower it. However, since they have observed the same communities only twice in time, they essentially use the contemporary variation across communities in migration to proximate the effect of changes in migration on inequality.

This paper analyzes the impact of rural-to-urban migration on inequality using the panel data derived from around 100 villages in rural China as observed four times (1997, 2000, 2004, and 2006) over a ten-year period. To our best knowledge, this is the first paper that examines the dynamic aspects of migration and income inequality using a dynamic panel data analysis. First, we are able to construct a relatively long panel of variables of many communities with a range of different migration experiences from individual and community level panel surveys on both incomes and migration, which are ideal in studying the dynamics of migration and income inequality (David McKenzie and Hillel Rapoport, 2007). In our study, the large sample size (N) is critical because we exploit both cross-section and time-series variations in the panel data. The large N provides the precondition for the asymptotic property of linear regressions to hold. Furthermore, the relatively long panel allows us to examine the dynamic aspects of migration and income inequality using a linear dynamic panel analysis.[4] Second, unlike earlier studies focusing exclusively on remittances, our data include the total labor earnings of migrants in the destination areas, which allow us to capture the general equilibrium effects. Third, we also look at the impact of migration on gender wage inequality within the sending communities, which is calculated from the key informant interview in the community panel survey. Last but not least, the massive wave of rural urban migrants in China, since its reform in 1980s, provides a unique context to test the relationship between migration and inequality at the community level. The structural barriers of integration into the urban society and the economic and psychological security offered by the home villages cause temporary migrants to maintain strong linkage with the source communities through remittances and return (R. Murphy, 2002). Therefore, the

[4] Recall that the panel data used by McKenzie and Rapoport (2007) only observed the same villages twice, in 1992 and 1997.

impacts of migration on the sending communities are more palpable than in the other contexts. Moreover, there are evidences to prove that selectivity for temporary migrants in particular has declined using the 1990 and 2000 Census data (Cindy Fan and Mingjie Sun, forthcoming). This decline in selectivity for temporary migrants provides suggestive evidence that migration, among other factors, has the potential to reduce inequalities within the sending communities in the long term.

Since income inequality is time-persisting, we used the system GMM to control the lagged income inequality in estimating the effect of emigration on income inequality in the sending villages. At the same time, contemporary emigration is instrumented in the GMM framework because of the unobserved time-varying community shocks that correlate with emigration and income inequality, and the potential reverse causality from income inequality to emigration. We found a Kuznets (inverse U-shaped) pattern between migration and income inequality in the sending communities. Specifically, contemporary emigration increases income inequality, while lagged emigration has strong income inequality-reducing effect in the sending villages. A 50- percent increase in the lagged emigration rate translates into one-sixth to one-seventh standard deviation reduction in inequality. Contemporary emigration has slightly smaller effects in raising the income inequality within villages. These effects are robust to different specifications and different measures of inequality. More interestingly, the estimated relationship between emigration and gender wage gap also has an inverse U-shape. Emigration tends to increase gender wage gap initially, and then decrease it in the sending villages.

The remainder of the paper is structured as follows: Section 2 gives the background of the rural–urban migration in China and briefly reviews the literature on migration and inequality in China. Section 3 presents the empirical strategy, a linear dynamic panel data analysis. Sections 4 and 5 describe the data and report the empirical results. Section 6 briefly discusses the implications and concludes the paper.

2. Background: Rural–urban migration in China

2.1 The *Hukou* system and its reform

Modeled after the *propiska* system in the Soviet Union and with roots that date back to ancient China, the Residence Registration System (*hukou*) was established in 1958; it ties citizens to a specific location within China through residency permits (K. W. Chan and W. Buckingham, 2008). The *hukou* also outlines an individual's rights to entitlements: in an agricultural area, the *hukou* entitles the holder to farmland, while a *hukou* in an urban area grants the holder access to jobs, housing, food, and other public services. The 1984 reform liberalized the movement of the rural poor, but without changing the *hukou* system; and without a local *hukou* (i.e., permanent change in residency) they are not fully entitled to social benefits (e.g., government housing) or public services (e.g., urban education system) or access to jobs in the destination areas. As in other areas of reform, the Chinese government has chosen a gradual and partial approach: providing labor rights but falling short of full abolishment of the *hukou* system. Analogous to the point-based system in host countries for international migrants, the *hukou* system engineered a two-tier migration scheme, whereby changes in permanent r*esiden*cy may be permitted for the highly skilled and college-educated migrant urbanites, but only temporary *resid*ency is usually granted for the less-skilled and less *e*ducated rural-to-urban migrants (Cindy Fan, 2008).

From the late 1980s to the mid-1990s, many city governments offered the "blue-stamp" *hukou* to well-off migrants who were able to make sizable investments. Training of rural–urban migrants is one of the foci in the early 1990s. The Migration Work Registration Card at the migrants' place of *hukou* origin and the Employment License at the place of destination were created to facilitate job searching and give migrants access to employment service from government agencies. In 1997, the State Council approved a pilot scheme to grant urban *hukou* to rural migrants who held stable jobs and had resided in selected towns and small cities for more than two years (Cindy Fan, 2008). In recent years, governments have undertaken reforms to establish a unified *hukou* regime to effectively eliminate the distinction between agricultural and non-agricultural *hukou*. Experiments in 12 provinces have been underway since 2007, although in general, small cities and towns have liberalized faster than the big metropolitan areas where barriers remain high. In 2007, the new Labor Contract Law gave migrant workers along with other ordinary workers better rights and greater protection in terms of entitlement to written labor contracts and long-term job security. In 2008, the Ministry of Human Resources and Social Security announced that the measures on the portable pension for migrant workers will be implemented by the end of 2008 (Fang Cai et al., 2009).

2.2. Trends and Basic Characteristics

The upsurge in the movement since 1990s was driven by the rapid growth in manufacturing jobs and higher pay in the coastal areas, in addition to the above-mentioned liberalization in the rule and regulations. The massive wave of migration of rural laborers to urban centers is estimated to result in 278 million increase in the permanent urban population from 1979 to 2003 (Xianghu Lu and Yonggang Wang, 2006). In 1990, the intercounty floating population was less than two percent of the total population, rising to 6.3 percent in 2000. Combining intercounty and intracounty in 2000 suggests that about one in nine Chinese are movers, amounting to a stock of migrants of 144 million (Cindy Fan, 2008). The five-year interprovincial migration flows from 1995 to 2000 also trebled from 12 million to 32 million, most of whom are temporary migrants without *hukou*.[5]

Figure 1 indicates that the rural-to-urban migration rate has increased dramatically from 6.8 percent in 1997 to 22.4 percent in 2006. The characteristics of migrants are distinct from the general population indicating they are a selective group. Permanent migrants tend to be positively selected in terms of educational attainment and the selectivity has increased over time. Permanent migrants, most of whom are college educated, have higher occupational attainment and tend to be employed as professionals, managers, and government officials. In contrast, there is evidence that selectivity for the temporary migrants in particular has declined based on the 1990 and 2000 Census data (Cindy Fan and Mingjie Sun, forthcoming). Two-thirds of both *hukou* and non-*hukou* interprovincial migrants fall into the age group between 20 and 39, and statistics show that the mean age of migrants has declined over time. Both permanent and

[5] Fan (2008) shows that temporary migrants account for almost three quarters of all inter-county migrants from 1995 to 2000.

temporary migrants are sex-selective, especially for temporary migrants with a sex ratio of 1.56 (males/females) in 1990. However, the sex ratio of migrants has declined sharply to 1.1, indicating increased female participation in migration. Over 80 percent of temporary migrants only have junior secondary education or below, although there is sign that they are becoming better educated as well. Owing to their low educational attainment, it is not surprising that temporary migrants concentrate on industrial and commercial service occupations. This decline in selectivity for temporary migrants provides suggestive evidence that migration in the long term may indeed reduce inequalities within the sending communities.

The structural barriers to integration into urban society and the economic and psychological security offered by the home villages cause temporary migrants to maintain strong linkage with the source communities through remittances and return (R. Murphy, 2002). Rural–urban migrants remit around 200 billion to 250 billion RMB (around US$25–30 billion) back to their families in the countryside, which is more than half of the central government's budget on agricultural development (Qiang Li et al., 2008). A national representative survey in 2004 by the National Statistical Bureau shows that seasonal migrants account for 20 percent of the total rural–urban migrants. Small-scale household surveys show that migration tends to have a cyclical character with more than one-third of migrants spending at least three months a year at origin counties (Hongyuan Song and Nansheng Bai, 2002). Therefore, the impacts of migration on the sending communities are more palpable in China than in other contexts.[6]

2.3 Impact of Rural Urban Migration on Migrants[7]

The overall impact of internal migration on the migrants themselves and their families in the rural areas are generally believed to be positive. Migration is shown to improve consumption and income levels and reduce poverty among migrant households. Most studies find positive impacts on education, with some negative impact related to rigid policies. Health outcomes are also generally promising, although segmentation and discrimination at the destinations limit the gains.

[6] China is one of the few remaining countries that still operates the household registration system and limits the integration of migrants into the urban economy. It is in this sense that migrants tend to see the sending communities as their home.

[7] For an overview of the rural and urban *hukou* related inequality, see *Human Development Report: China 2007/08* (http://hdr.undp.org/en/reports/nationalreports/asiathepacific/china/China_2008_en.pdf) (UNDP, 2008).

Migration often has economic payoffs, but the road is not as rosy as one would like to believe. The following aspects of migration are worth our attention.

- **Income gain.** Income gain is the first and foremost motive for migrants. In 2004, rural–urban migrants on the average, earn 780 RMB per month, which is more than three times the average income per capita of a typical rural farmer (Research team on the Issue of Chinese Farmer-turned Workers, 2006). However, due to labor market segmentation in China created by the *hukou* system, temporary migrants typically move to areas with large numbers of low-skilled jobs and become either self-employed or employed in dangerous, difficult, and low-paying temporary jobs (e.g., in manufacturing, service, construction, and similar jobs), that are not desired by the local residents (X. Meng and J. Zhang, 2001). The poverty incidence of migrants is double that of urban residents with *hukou*.

- **Working condition and employment benefits.** Low-skilled migrants tend to work in the informal sectors that have inadequate labor protection and benefit package. According to one survey covering three provinces, migrants' work hours are 50 percent longer than locals while their pay is only 60 percent of their counterparts. Migrants are often hired without any written contracts. According to the China Urban Labor Survey, in all work units with migrant workers, less than 10 percent of migrant workers are provided with old-age social security and medical insurance, while more than two-thirds of their urban counterparts enjoy these benefits (Fang Cai, Yang Du and Meiyan Wang, 2009). Occupational hazards are high in the mining and construction industries where migrants tend to concentrate. Migrants account for 75 percent of all the 11,000 fatalities in 2005 in these industries (Quanquan Huang, 2006).

- **Access to services and outcomes.** Children who move with their parents tend not to do as well as the locals, largely due to their parents' temporary status. They pay additional fees yet lack access to the elite schools. The total number of migrant children who lack access to education in China is estimated to be 14 to 20 million. Official surveys have found drop-out rates at the primary and secondary schools to exceed 9 percent (compared to almost universal for locals). The proportion of those who had never been in school has also increased from 0.8 percent to 15.4 percent for ages 8 to 14 and over 60 percent of drop-outs aged 12 to 14 had already ventured into child labor (Research team on the Issue of Chinese Farmer-turned Workers, 2006). Access to basic health service is limited. Even in Shanghai, one of the model cities in terms of providing social service to migrants, only two-thirds of migrant children received vaccination in 2004 compared to the universal immunization for local children. The occupational health hazard in urban areas is high. When they become ill, they may have to move back to the rural areas to receive medical treatment owing to the exorbitant health expense in urban hospitals and their lack of health insurance.

- **Participation and integration.** Many of them have gradually adopted the value system of urban society in the way they consume and socialize with people but remain marginalized in the destination places owing to institutional barriers epitomized by the *hukou*. Migrant workers in China have few channels to express their interests and to protect their rights in the workplace. Seventy-eight percent of the migrants say their work units have no trade union, workers' representative conference, labor supervisory committees, or other labor organizations compared to the state of the 22 percent of urban workers (Fang Cai and Dewen Wang, 2008). Long-distance migration also hinders the expression of their voting rights in their villages. In a survey among migrants in Wuhan City, only 20 percent of the migrants confirmed that they have voted in the last village election (Zengyang Xu and Huixiang Huang, 2002).

2.4 Impact of Migration on Sending Communities

There are also many studies on the impact of migration on those left-behind. However, these studies tend to take a micro perspective focusing on the effect of migration at the household level. The impact on how migration transforms the rural areas, which were isolated for more than two decades, has been largely overlooked by the existing literature (R. Murphy, 2002). This imbalance is what this paper intends to address. In general, rural–urban migration has been viewed as positive in China. However, if migration increases inequality in the sending communities among the poor who were left-behind in rural areas, the benefits of migration need to be reevaluated. Nonetheless, existing studies find that remittances significantly contribute to the sending household's per capita income and consumption, which in turn reduces poverty. Remittances also allow migrants to subsidize the expenditure on schooling and medical care.

- **Consumption and income**. Per capita consumption and income are 8.5–17 percent higher for households with a rural–urban migrant (Yang Du et al., 2005; Nong Zhu and Xuebei Luo, 2008).

- **Poverty Reduction**. Among households with a migrant, poverty incidence falls from 28 percent to 14 percent (Nong Zhu and Xuebei Luo, 2008). However, the effect on aggregate poverty (e.g., 1 percent point drop in aggregate poverty) is limited because migrants do not come from the poorest households (Yang Du, Albert Park, and Sangui Wang, 2005).

- **Education**. Many migrants explicitly state that their primary motive is to pay for their sibling or children's education. However, the separation from one or both of their parents can put children at risk due to lack of supervision and interaction with parents. Furthermore, high school fees and poor career prospects may deter investments in higher levels of school education (A. de Brauw and J. Giles, 2008b).

- **Health**. Jalan and Ravallion (2001) find that rural households in China are more likely to send members to migrate in order to pay for health expenditures. Studies report that children left behind are marginally less healthy psychologically than other children because of the separation from parents (Biao Xiang, 2005).

- **Inequality**. The conventional wisdom is that migration contributes to greater inequality in the places of origin because of positive selection (see Zhu and Luo (2008) for a review of this literature). However, more recent findings suggest the opposite. Following Barham and Boucher (1998), Zhu and Luo (2008) simulated the counterfactual income distribution in the absence of migration and remittances, and found that migration reduces the Gini coefficient by 16.7 percent. Moreover, de Brauw and Gils (2008a) suggest that out-migration from the village leads to growth in per capita income and consumption, and that migrant opportunity is contributing to more rapid economic growth among poorer households within villages. This result is consistent with ethnographic studies where migration is shown to help promote equality in the natal communities because in obtaining resources which exist outside the power-based distributional mechanisms of the villages, migrants make the boundaries of stratification more fluid (D. Benjamin et al., 2005; R. Murphy, 2000).

3. Empirical strategy

This section specifies the system GMM model to test the effect of rural-to-urban migration on the income inequality in the sending communities in China by using the panel data at the village level. Specifically, the following regression equation will be estimated:

$$Ineq_{i,t} = \beta_0 + \beta_1 Ineq_{i,t-1} + \beta_2 Emig_{i,t} + \beta_3 Emig_{i,t-1} + X_{i,t}\beta_4 + \tau_i + \eta_t + v_{it} \qquad (1)$$

where i and t index village and time period ($i = 1,...,I$ and $t = 0,...T$), respectively; $Ineq$ measures the inequality such as the Gini coefficient, Theil index, or gender wage ratio; $Emig$ is the share of emigrants out of the total labors whose *hukou* is in the village while working outside; X is a vector of other control variables; τ and η are community and time fixed effects, respectively; v is an error term; and $\beta_j\,(j = 0,...,4)$ are coefficients to be estimated.

Note that we have included the lagged inequality in the regression equation because aggregate variables such as GDP and inequality are time-persisting which means that they are serially correlated over time. Therefore, we are estimating a linear dynamic panel data model. In addition, we have included both current emigration and lagged emigration in Equation (1) because current emigration and lagged emigration may have different effects on inequality. The current migration reflects an immediate effect, while the lagged emigration reflects a more accumulated effect of migration with a build-up of migration specific to human capital and networks.

There are two potential pitfalls in Equation (1), which will bias the ordinary least square (OLS) estimates. First, the unobserved community heterogeneities (τ) may be correlated with the other independent variables in the right hand side of this equation, which creates an omitted variable bias problem. Second, the "relative deprivation" model (O. Stark and J. E. Taylor, 1991, Oded Stark and David E. Bloom, 1985) states that the household's relative position with respect to a specific reference group in addition to the household's absolute income serves as strong motivation for emigration. Therefore, causality may also go from inequality to emigration ($Emig_{i,t}$).

In discussing the problem of community heterogeneities, we assume that $Emig_{i,t}$ is exogenous for the moment. Unlike the static model, the fixed effects method could not eliminate the inconsistency induced by the community heterogeneities in the dynamic model of Equation (1), because $v_{i,t}$ will correlate with the future value of the regressors due to the presence of the lagged dependent variable in the right hand of the regression equation. In other words,

$$E[(Ineq_{i,t-1} - \overline{Ineq_{i,t-1}})(v_{i,t} - \overline{v_{i,t}})] \neq 0$$

where $\overline{Ineq_{i,t-1}}$ and $\overline{v_{i,t}}$ are the within the group mean values of $Ineq_{i,t-1}$ and $v_{i,t}$, respectively.

Arellano and Bond (1991) developed a difference GMM to deal with this kind of community heterogeneities. To illustrate this method clearly, we make a first difference with respect to Equation (1),

$$\Delta Ineq_{i,t} = \beta_1 \Delta Ineq_{i,t-1} + \beta_2 \Delta Emig_{i,t} + \beta_3 \Delta Emig_{i,t-1} + \beta_4 \Delta X_{i,t} + \Delta v_{it} \qquad (2)$$

where Δ is the operator of the first difference.[8] It could be clearly seen that $\Delta Ineq_{i,t-1}$ is endogenous because

$$\begin{aligned} & E(\Delta Ineq_{i,t}\Delta v_{i,t}) \\ &= E[(Ineq_{i,t} - Ineq_{i,t-1})(v_{i,t} - v_{i,t-1})] \\ &= E(Ineq_{i,t}v_{i,t}) + (1-\beta_1)E(Ineq_{i,t-1}v_{i,t-1}) \\ &\neq 0 \end{aligned}$$

It is assumed that (1) error terms ($v_{i,t}$) are serial uncorrelated, $\operatorname{cov}(v_{i,t}, v_{i,s}) = 0$ if $s \neq t$; (2) initial condition, $E(Ineq_{i,0}v_{i,t}) = 0$ for $t \geq 1$, and (3) $E(\eta_i v_{i,t}) = 0$.[9] Under these three assumptions, we can derive the moment conditions for the difference GMM method as follows:

$$E[Ineq_{i,s}\Delta v_{i,t}] = E[Ineq_{i,s}(v_{i,t} - v_{i,t-1})] = 0, \qquad \text{when } s = 0,1,...,T-2 \tag{3}$$

Thus, all $Ineq_{i,s}$ ($s = 0,1,...,t-2$) are valid instruments for $\Delta Ineq_{i,t-1}$ in Equation (2).

We then use a similar method to deal with the second problem, that is, the endogeneity of $Emig_{i,t}$ in Equation (1) above. Assuming that there is simultaneity and feedback between emigration and inequality in the sending communities, that is, $E(Emig_{i,t}v_{i,s}) \neq 0$ for $s \leq t$ and $E(Emig_{i,t}v_{i,s}) = 0$ for $s > t$.[10] Under this assumption, the moment condition is

$$E[Emig_{i,s}\Delta v_{i,t}] = E[Emig_{i,s}(v_{i,t} - v_{i,t-1})] = 0, \qquad \text{when } s = 0,1,...,T-2 \tag{4}$$

8 η_t in Equation (1) is ignored because we assume that all variables in Equation (2) are the deviations from the period mean.

9 For further discussion, see Arellano and Bond (1991).

10 It implies that $E(Emig_{i,t}v_{i,t}) \neq 0$ and $E(Emig_{i,t-1}v_{i,t}) = 0$ in Equation (1).

Thus, all $Emig_{i,s}$ ($s = 0,1,...,t-2$) are valid instruments for both $\Delta Ineq_{i,t}$ and $\Delta Ineq_{i,t-1}$ in Equation (2).[11] Therefore, the difference GMM exploits the two sets of moment conditions (3) and (4) to estimate Equation (2).

Blundell and Bond (1998) show that when the time period is short (T is small) or the dependent variable is highly time-persisting ($|\beta_1|$ is close to 1), the standard difference GMM suffers from the problem of weak instrumental variables. They label moment conditions such as Equations (3) and (4) as moment conditions in differences. In addition to these moment conditions in differences, they also exploit another set of moment conditions, which are called moment conditions in levels.

Under the same assumptions discussed above, Blundell and Bond (1998) derive the following moment conditions for Equation (1)

$$E[\Delta Ineq_{i,t-1}(\eta_i + v_{i,t})] = 0, \qquad \text{for } t = 2,3,...T \tag{5}$$

and

$$E[\Delta Emig_{i,t-1}(\eta_i + v_{i,t})] = 0, \qquad \text{for } t = 2,3,...T \tag{6}$$

In other words, $\Delta Ineq_{i,t-1}$ and $\Delta Emig_{i,t-1}$ are used to instrument $Ineq_{i,t-1}$ and $Emig_{i,t-1}$ in Equation (1) for $t \geq 2$, respectively.

Blundell and Bond (1998) call this the system GMM that estimates Equations (1)-(2) simultaneously by exploiting the moment conditions (3)-(6) and demonstrate that the system GMM estimators are very robust even in a finite sample.

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[11] Under the assumption of $E(Emig_{i,t}v_{i,s}) \neq 0$ for $s \leq t$ and $E(Emig_{i,t}v_{i,s}) = 0$ for $s > t$, $Emig_{i,t-1}$ is an exogenous variable in Equation (1) because $E(Emig_{i,t-1}v_{i,t}) = 0$. However, $\Delta Ineq_{i,t-1}$ is not an exogenous variable in Equation (2) because $E(\Delta Emig_{i,t-1}\Delta v_{i,t}) = E[(Emig_{i,t-1} - Emig_{i,t-2})(v_{i,t} - v_{i,t-1})]$, which is $-E(Emig_{i,t-1}v_{i,t-1})$ by manipulation and it is not equal to zero.

In summary, there are several benefits in using the system GMM to estimate the effect of emigration on income inequality in the sending communities. First, the unobservable community heterogeneities that may affect emigration and income inequality simultaneously are safely swept out. Second, the lagged inequality is controlled for in the regression equation. To our best knowledge, there is no other study that has treated the lagged income inequality properly when estimating the effect of emigration on inequality. Since aggregate time series such as inequality often exhibit strong time persistency, it is necessary to control for it when estimating the effect of emigration on inequality. Lastly and most importantly, contemporary emigration is validly instrumented and the concern of a potential reverse causality running from income inequality to contemporary emigration is cleared out in our system GMM framework. Thus, the system GMM estimate of the effect of emigration on inequality suggests a causal relationship.[12]

4. Data

We use the Chinese Health and Nutrition Survey (CHNS), which is a panel dataset with seven survey waves (1989, 1991, 1993, 1997, 2000, 2004, and 2006).[13] The CHNS is conducted by the Carolina Population Center (CPC) at the University of North Carolina, Chapel Hill, The Institute of Nutrition and Food Hygiene, and the Chinese Academy of Preventive Medicine. The CHNS surveys were conducted by an international team of researchers whose backgrounds include nutrition, public health, sociology, Chinese studies, demography, and economics. The CPC expended considerable effort in staff training and quality control to ensure that the data are of high quality.

The survey was conducted at both the community and household levels. A community refers to a village in a rural area or a neighborhood in an urban area. It is the basic level of China's administrative hierarchy. The survey sampled communities, which were randomly drawn in nine provinces of Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong. These provinces vary substantially in geography, economic development, public

[12] These estimates capture the general equilibrium effects of migration on inequality; including the effects through direct remittances, multiplier effects of remittances from the spending of remittances on local non-tradable products and services, as well as the network effects (McKenzie and Rapoport, 2007).

[13] We are grateful to the UNC Carolina Population Center for providing the data.

resources, and health indicators.[14] A multistage, random cluster process was used to draw the sample surveyed in each province. In 1989–1993, there were 190 primary sampling units: 32 urban neighborhoods, 30 suburban neighborhoods, 32 towns, and 96 rural villages. Since 2000, the primary sampling units have increased to 216: 36 urban neighborhoods, 36 suburban neighborhoods, 36 towns, and 108 villages.[15] Currently, there are about 4,400 households in the survey, with a total of 19,000 individuals.[16]

Between 20 and 35 households were randomly drawn from each community, and the CHNS survey covers all the members formally registered in a household or those with permanent residence or *hukou*. The CHNS survey is designed to examine the effects of health, nutrition, and family planning policies in China, and collects detailed information on economic, demographic, and social characteristics of individuals, households, and communities.

This dataset provides us a precious and unique opportunity to examine the dynamic relationship between migration and inequality. In fact, the CHNS has just issued both constructed household and individual income data in longitudinal form on November 11, 2008 for all the seven waves. Thus, this paper should be one of the first studies that have used the CHNS longitudinal income file to conduct economic research.

We restrict our analysis to the last four waves of the CHNS survey (1997, 2000, 2004, and 2006) because it is only in these waves that specific questions on labor migration were asked at the household level. A community panel dataset is constructed by aggregating household information to the village level, and we eventually have 395 observations from four different waves, in which there are 95 unique communities sequentially appearing at least three times in the four waves.[17]

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[14] Jiangsu, Liaoning, and Shandong are among the richest provinces; Henan and Hunan among the middle, and Guangxi and Guizhou are among the poorest. Geographically, Jiangsu, Liaoning, Shandong, and Guangxi are coastal regions, while the other provinces are inland regions (Chen and Zhou, 2007).

[15] For further information, see http://www.cpc.unc.edu/projects/china/design/survey.html

[16] Probably, attrition problems may exist in the CHNS data because aside from households, communities are also replaced. However, we find that more than 90 percent of the communities surveyed in 1997 were also surveyed in 2006.

[17] The system GMM estimation needs at least three waves.

Table 1 presents the summary statistics for the main variables.

- **Inequality**. The Gini coefficient and Theil index are calculated using both household total income and per capita income from the income module of the household survey. We also retain the information on the gender wage ratio from the key informant in the community survey as a measure of gender wage disparity in the villages.[18] It is shown that the mean value of the Gini coefficient by using the household total income in our study period is 0.39. It is highly consistent with the calculated Gini coefficient in the rural areas in Gustafsson, Li, and Sicular (2008) study, which implies that we are using a representative sample. When we calculated the Gini coefficient by using the household per capita income, the Gini coefficient is slightly smaller.

- **Emigration.** In this paper, we define an emigrant as one whose *hukou* was in the village, while he/she was working or seeking employment outside the village when the survey was carried out. In this sense, we are analyzing the effect of *temporary emigration* on the inequality of the sending village because the *hukou* of these emigrants are still in the village.[19] The emigration rate is defined as the ratio of emigrants to the total number of labors in the village whose ages are older than 17 and younger than 60. Table 1 shows that during our study period, there was about 15.15 percent of the total labor in the village working outside the village on the average. Figure 1 shows that emigration rates have increased dramatically during the study period from 1997 to 2006.

- **Additional controls.** Table 1 also presents the summary statistics for other control variables, which are time-varying at the community level. They include demographic structural variables such as shares of elder people, young people (younger than 15

[18] Although it is difficult to measure the gender wage gap in the rural areas because a lot of females only work within the household, the community survey in the CHNS provide the information for both the male and female daily wages at the village reported by the village leader.

[19] Since a community in the CHNS refers to a village in a rural area or a neighborhood in an urban area and since our study only focuses on rural-to-urban migration, the word community level is used interchangeably with village level in this paper.

years old), the ethic indicator of Han, and the number of people in relation to the total population who were born in other provinces, while their *hukou* was in the village; educational compositional variables such as the shares of people with different educational attainments. We also include the household per capita income (measured in the 2006 constant price) in the regression as a robustness test.

5. Empirical results

This section presents our main empirical results. We first report the estimated effects of emigration on the income disparity across households in the sending villages. Secondly, we present the estimation results of the effect of emigration on the gender wage gap in the sending villages.

5.1 The effect of emigration on income inequality in the sending villages

Table 2 presents the system GMM estimates of the effects of emigration on the Gini coefficient calculated based on household income. As we have discussed in the Empirical Strategy section, we used all lagged inequality and all lagged emigration to instrument the difference in the lagged inequality, and the differences in emigration and the lagged emigration in Equation (2). At the same time, the difference in the lagged inequality and the difference in emigration were used to instrument the lagged inequality and emigration in Equation (1). We then estimated Equations (1) and (2) simultaneously.[20]

Column (1) includes only the lagged Gini coefficient, contemporary emigration, and lagged emigration besides the survey year fixed effects. First of all, it is noted that the estimated coefficient on the Gini coefficient is positive and statistically significant at a high level of 1 percent. It justifies the use of the system GMM model because the dependent variable of the Gini coefficient is serially correlated. Thus, omitting the lagged dependent variable of inequality in

[20] Note that the number of observations in Tables 2-5 is smaller than that of Table 1. Equation (1), being a dynamic equation accounts for this. Thus, one wave of observations is not counted. In addition, GMM estimation requires that a village appears at least three sequential times in the waves (1997, 2000, 2004, and 2006).

Equation (1) will necessarily result in the problem of omitted variables, and the usual fixed effects estimates will be biased. The test result of the Arellano-Bound test for the first-order serial correlation further substantiates the necessity of using the system GMM model because the null hypothesis of no first-order serial correlation is statistically rejected at a high level of 1 percent.[21] In addition, the estimated coefficient on the Gini coefficient is 0.39, or less than 1. It means that the economic pattern of inequality follows a standard conditional convergence.

Second, we found that the estimated coefficient on the current emigration was positive. Although it is statistically insignificant, the t-statistic of the estimated coefficient on current emigration is larger than 1. On the contrary, the estimated coefficient on the lagged emigration is negative and it is statistically at a high level of 5 percent. It is interesting to find that the positive effect of current emigration and the negative effect of the lagged emigration on inequality is just consistent with the inverted U-shaped effect of emigration on inequality (David McKenzie and Hillel Rapoport, 2007).

One may wonder whether the insignificance of the current migration variable and the opposite signs of the current and lagged migrations result from the high correlation (0.67 with a p-value less than 0.01) between the current and lagged migrations. In Tables A1 and A2 in the Appendix, we report the GMM estimates for specifications with current and lagged migrations entering the regressions separately. The coefficients of the current (lagged) migration remain positive (negative) and statistically insignificant (significant). Thus, the high correlation is not the main reason for the results. As discussed before, the current migration reflects the immediate effect of migration, which is likely to be small. On the other hand, the lagged migration reflects a more accumulated effect of migration with a build-up of migrant specific human capital and networks. Thus, it is not surprising that the lagged migration has a much larger effect than the current migration.

[21] Since we only have four waves of observations, the Arellano-Bond test for second-order serial correlation could not be carried out.

Finally, we further conduct a Hansen overidentification test to examine the validity of the additional instruments. The p-value of the Hansen tests suggests that these instruments are statistically valid. [22]

Columns (2)-(5) of Table 2 sequentially add other variables to control for the demographic and educational differences across villages. It is shown that the estimated positive effect of the current migration and the estimated negative effect of the lagged migration on income inequality are strongly robust to the inclusion of these control variables. Neither the magnitudes nor the standard errors of the estimated coefficients on both contemporary emigration and lagged emigration have changed substantially after controlling for these demographic and educational variables.

In terms of the control variables, the share of old people tends to increase the income inequality, while the share of young people tends to decrease the income inequality. Compared with the omitted baseline group, the share of illiterate or semi-illiterate people, the shares of primary school-educated, middle school-educated, and other higher level-educated people tend to decrease income inequality. However, all these demographic and educational variables are insignificant in the inequality equation. The reason may be that there is insufficient change across the time within a village during our study period, although emigration behavior has experienced a substantial change, which is illustrated in Figure 1.

Column (6) of Table 2 includes the per capita income. It is shown that per capita income increases the income inequality at the village level although the estimated coefficient is statistically insignificant. However, caution should be exercised in interpreting the effect of per capita income on income inequality because inequality can simultaneously affect economic growth and per capita income (A. V. Banerjee and E. Duflo, 2003; K. J. Forbes, 2000). Since this paper focuses on the relationship between emigration and inequality and there is no good

[22] The Hansen test is a test of overidentification restrictions, which allow for heteroscedasticity robust standard errors. The joint null hypothesis is that the excluded instruments are correctly excluded from the structural growth equation, and the structural equation is correctly specified. Under the null, the test statistic is asymptotically distributed as chi-squared with the degree of freedom equal to the number of overidentification restrictions. For further discussion, see **Hayashi, F.** *Econometrics*. Princeton: Princeton University Press, 2000.

instrument for per capita income, we included it in the inequality equation only as a robustness check.

In Table 3, we calculate the Gini coefficient based on the per capita income and use it as the dependent variable. It is found that the estimation results and test results in Table 3 are very similar to those in Table 2. In addition, the estimated coefficients on the contemporary emigration are marginally significant in Columns (1)-(4), and become statistically significant at the level of 5 percent in the last two columns. This finding further confirms the inverse U-shaped relationship between emigration and inequality.

Although the Gini coefficient is perhaps the most popular measure of inequality in the past years, it possesses many undesirable properties (A. Deaton, 1997).[23] Thus, we replace the Gini coefficient with the Theil index, another popular measure of inequality in practice, and repeat the exercises in Tables 2-3. The results of the Theil index, as the dependent variable, are reported in Tables 4-5. We find that the estimated inverted U-shaped relationship between emigration and income inequality is robust to different measures of inequality. Specifically, the estimated coefficients on the contemporary emigration are consistently positive in all the specifications of Tables 4-5, and they are statistically significant at the level of 10 percent in Table 5. In contrast, the estimated coefficients on the lagged emigration are consistently negative and statistically significant, at least at the level of 10 percent in all the specifications of Tables 4-5.

5.2 The effect of emigration on the gender wage gap in the sending villages

Tables 6-7 report the fixed effects estimates of the effect of emigration on the gender wage gap in the sending villages. Before discussing the empirical results, it should be noted that we have also performed the GMM estimation for the gender wage gap equation. In contrast to the regression results of the income inequality equation (Tables 2-5), it is found that the lagged gender wage gap is insignificant in the regression equation. In addition, the Arellano-Bond test

[23] For example, economies with similar incomes and the Gini coefficients can still have very different income distributions. The ability of the Lorenz curves to have different shapes and yet still yield the same Gini coefficient is the reason.

results for the first-order serial correlation show that the null of the no first-order serial correlation is not rejected.[24] Thus, there is no need to use the dynamic model of GMM, and a static panel model such as fixed effects estimation is sufficient.

Unlike in the GMM estimation, we have no valid instrument for contemporary emigration. Since contemporary emigration is a possible endogenous variable, Table 6 only includes the lagged emigration as an interested independent variable aside from the other control variables. It is interesting to find that the estimated coefficients on the lagged emigration are consistently negative and statistically significant at least at the 10 percent level in all the specifications of Table 6. It implies that emigration decreases the gender wage gap in the long run.

We have also included the gender ratio of emigrants (measured by the share of males out of the total emigrants) in the gender wage gap equation to detect whether the gender composition of emigrants has an effect on the gender wage gap. It is found that the coefficients on gender ratio become statistically significant. The sign of the coefficients on the gender ratio is positive, which is consistent with the demand–supply story. When the gender ratio of emigrants is high, more men migrate, thus few men remain in the source regions. Consequently, the high wage for men results in a large gender wage gap. However, even while controlling for gender ratio, the negative effect of migration on the gender wage gap still holds.

Although we do not have a convincing instrument for the contemporary emigration, it is interesting to see what happens when we include it into the regression equation to test for robustness. In addition, unlike the income inequality equation in which income inequality is one of the driving forces for emigration, the reversal effect of the gender wage gap on contemporary emigration does not seem to be so obvious. Table 7 reports the results when we include both the current and lagged emigration in the gender wage gap equation. It is found that the estimated coefficients on the contemporary emigration are consistently positive although they are statistically insignificant. At the same time, the estimated coefficients on the lagged emigration are consistently negative and they are statistically significant.

[24] These results are available upon requests.

Thus, similar to the relationship between emigration and income inequality, the relationship between emigration and gender wage gap also seems to have an inverse U-shaped. This inverse U-shaped relationship between emigration and the gender wage gap is consistent with the Fan and Sun's (2009) finding. Fan and Sun (2009) find that earlier emigrants in the 1990 census are male-biased while the later emigrants in the 2000 census are more gender-neutral.

6. Conclusion and discussion

Migration and inequality have a complex and intimate relationship. It is widely believed that income inequality between the source and destination areas is one of the most important factors driving economic migration within and across borders. In contrast to this consensus in the literature, existing studies on the impact of migration on income inequality are scarce and the results are less unequivocal.

Although researchers have long contemplated the Kuznets pattern between migration and inequality in the sending communities – that is, inequality rises in the beginning of the migration process and drops after migration becomes more established – the literature has little to offer in terms of solid empirical evidence. Earlier studies treated remittance income as an exogenous transfer, and compared income inequality with and without the inclusion of remittance income. More recently, remittances are treated as a potential substitute for home earnings and the observed income distribution with remittances are compared to a counterfactual scenario in which no migration takes place but includes an imputed level of home earnings. Although the earlier approach is unrealistic in assuming that remittance-earning migrants are separate entities from their households in rural areas, the improvement that the counterfactual model provides is limited because the selection into migration is difficult, if not impossible, to model. In addition to these methodological challenges, the usefulness of earlier literature is also tempered by the cross-section nature and small sample sizes of the sources of their data. The lack of panel data at the community level seriously limits the researchers' ability to quantify the temporal dimension of migration and inequality.

The massive wave of rural urban migrants in China since its reform in 1980s provides a unique context to test the relationship between migration and inequality at the community level. The massive wave of migration of rural laborers to urban centers is estimated to result in 278 million increase in the urban population from 1979 to 2003 (Xianghu Lu and Yonggang Wang, 2006). However, the majority remains as temporary migrants. The structural barriers for integration into the urban society and the economic and psychological security offered by the home villages cause temporary migrants to maintain strong linkage with the source communities through remittances and return (R. Murphy, 2002). Therefore, the impact of migration on the sending communities is more palpable than in the other contexts. Moreover, there are pieces of evidence that selectivity for temporary migrants in particular has declined based on the 1990 and 2000 Census data (Cindy Fan and Mingjie Sun, forthcoming). This decline in selectivity for temporary migrants provides suggestive evidence that migration in the long term has the potential to reduce inequalities within the sending communities.

This paper analyzes the impact of rural-to-urban migration on inequality using a newly constructed panel for around 100 villages over a ten-year period from 1997 to 2006 in China. To our best knowledge, this is the first paper that examines the dynamic aspects of migration and income inequality using a panel data analysis. Unlike earlier studies focusing exclusively on remittances, our data take into account the total labor earnings of migrants in destination areas. Furthermore, we also look at the gender dimension of the impact of migration on wage inequality within the sending communities.

Since income inequality is time-persisting, we used a system GMM framework to control for the lagged income inequality in estimating the effect of emigration on income inequality in the sending villages. At the same time, contemporary emigration is instrumented in the GMM framework because of the unobserved time-varying community shocks that correlate with emigration and income inequality and potential reverse causality from income inequality to emigration. We found a Kuznets (inverse U-shaped) pattern between migration and income inequality in the sending communities. Specifically, contemporary emigration increases income inequality while lagged emigration has strong income inequality-reducing effect in the sending villages. A 50-percent increase in the lagged emigration rate translates into one-sixth to one-seventh standard deviation reduction in income inequality. Contemporary emigration has slightly smaller effects in resolving the income inequality within the villages. These effects are robust to the different specifications and different measures of inequality. More interestingly, the estimated relationship between emigration and the gender wage gap also has an inverse U-shape. Emigration tends to increase the gender wage gap initially, and then decrease it in the sending villages.

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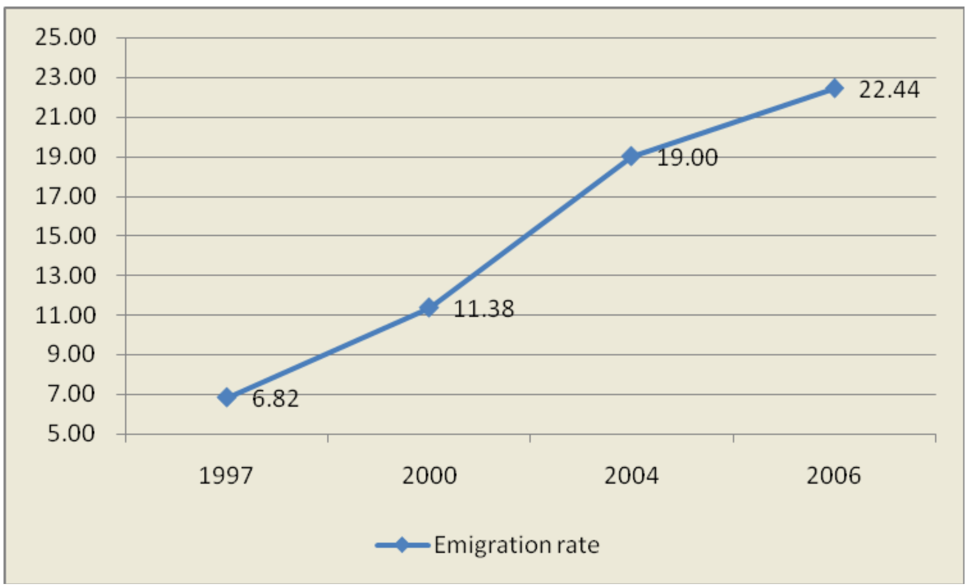

Figure 1: Emigration rates in rural China (1997–2006)

Data source: CHNS data (1997, 2002, 2004, 2006)

Note: The emigration rate is defined as the share of population (older than 17 and younger than 60) to the total labor force in the village who work outside the village where their *hukou* is registered.

Table 1: Summary statistics for the main variables

Variable	Obs	Mean	Std. Dev.
Dependent variables			
Gini coefficient (household total income)	395	0.3890	0.0975
Gini coefficient (household per capita income)	395	0.3717	0.0985
Theil index (household total income)	395	0.2921	0.1776
Theil index (household per capita income)	395	0.2640	0.1707
Gender wage gap (ln(male wage)-ln(female wage))	352	0.2713	0.2817
Interested independent variable			
Emigration: Share of working outside the village in total labor (%)	395	15.1536	13.1922
Control variables			
Share of elder people (age>64) (%)	395	8.9466	5.9216
Share of young people (age<15) (%)	395	15.6382	6.7367
Share of Han (%)	395	89.7811	25.3118
Share of people born in other provinces (%)	395	2.9352	6.1644
Share of illiterate or semi-illiterate people (%)	395	25.1617	14.9757
Share of primary-educated people only (%)	395	23.2928	11.0782
Share of middle school-educated people only (%)	395	34.3526	11.4260
Share of high school and higher-educated people (%)	395	17.0885	16.6549
Household per capita income (RMB/2006)	395	4638.0340	2872.3160

Data source: CHNS data (1997, 2002, 2004, 2006)

Notes: Sample is restricted to rural communities

Table 2: System GMM estimates of the effects of emigration on income inequality in rural China

	Dependent variable: Gini coefficient (Household total income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Gini (Household total	0.388***	0.389***	0.386***	0.387***	0.389***	0.389***
income)(t-1)	(0.116)	(0.114)	(0.115)	(0.112)	(0.110)	(0.107)
ln (% Emigration)	0.018	0.019	0.020	0.020	0.017	0.018
	(0.016)	(0.016)	(0.016)	(0.016)	(0.016)	(0.016)
ln (% Emigration)(t-1)	-0.025**	-0.026**	-0.026**	-0.026**	-0.028**	-0.027**
	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)	(0.013)
ln [% Old (age>64)]		0.013	0.013	0.013	0.012	0.014
		(0.008)	(0.008)	(0.008)	(0.009)	(0.008)
ln [% Young (age<15)]		-0.018	-0.019	-0.019	-0.019	-0.016
		(0.015)	(0.015)	(0.016)	(0.015)	(0.015)
ln (% Han)			-0.009	-0.009	-0.007	-0.008
			(0.008)	(0.008)	(0.007)	(0.007)
ln (% Born in other				0.000	0.001	-0.000
provinces)				(0.009)	(0.009)	(0.009)
ln (% Primary)					-0.005	-0.005
					(0.016)	(0.017)
ln (% Middle school)					-0.001	-0.001
					(0.015)	(0.015)
ln (% Higher education)					-0.006	-0.008
					(0.009)	(0.009)
ln (per capita income)						0.014
(RMB at year 2006)						(0.016)
Wave 2000	-0.024	-0.012	-0.011	-0.011	-0.015	-0.010
	(0.024)	(0.028)	(0.028)	(0.028)	(0.030)	(0.029)
Wave 2004	-0.040**	-0.037**	-0.036**	-0.036**	-0.037**	-0.035**
	(0.015)	(0.016)	(0.016)	(0.016)	(0.017)	(0.017)
Hansen overidentification test	15.55	16.86	16.92	16.91	15.87	15.76
P-values	0.556	0.464	0.460	0.461	0.533	0.541
Arellano-Bond test for First-order serial correlation	-2.919	-3.104	-3.109	-3.120	-3.041	-3.277
P-values	0.004	0.002	0.002	0.002	0.002	0.001
Observations	238	238	238	238	238	238
Number of villages	95	95	95	95	95	95

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table 3: System GMM estimates of the effects of emigration on income inequality in rural China

	Dependent variable: Gini coefficient (Per capita income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Gini (Per capita income income)(t-1)	0.192*	0.199*	0.197*	0.207*	0.204*	0.205*
	(0.110)	(0.109)	(0.110)	(0.107)	(0.108)	(0.107)
ln (% Emigration)	0.022	0.022	0.021	0.024	0.024*	0.025*
	(0.015)	(0.016)	(0.015)	(0.015)	(0.015)	(0.015)
ln (% Emigration)(t-1)	-0.024*	-0.024*	-0.026**	-0.027**	-0.027**	-0.027**
	(0.012)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
ln [% Old (age>64)]		0.003	0.002	0.001	0.002	0.003
		(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
ln [% Young (age<15)]		-0.007	-0.007	-0.009	-0.009	-0.008
		(0.016)	(0.016)	(0.016)	(0.015)	(0.015)
ln (% Han)			-0.011	-0.011	-0.012	-0.012
			(0.008)	(0.009)	(0.009)	(0.009)
ln (% Born in other provinces)				0.007	0.007	0.007
				(0.009)	(0.008)	(0.008)
ln (% Primary)					0.006	0.006
					(0.019)	(0.019)
ln (% Middle school)					0.011	0.011
					(0.016)	(0.016)
ln (% Higher education)					0.002	0.002
					(0.008)	(0.009)
ln (per capita income) (RMB at year 2006)						0.004
						(0.018)
Wave 2000	-0.040*	-0.035	-0.037	-0.035	-0.037	-0.036
	(0.022)	(0.027)	(0.027)	(0.027)	(0.028)	(0.029)
Wave 2004	-0.038**	-0.037**	-0.038**	-0.040**	-0.042**	-0.041**
	(0.016)	(0.017)	(0.017)	(0.017)	(0.018)	(0.018)
Hansen overidentification test	15.37	16.30	16.01	16.10	16.38	16.42
P-values	0.569	0.503	0.523	0.517	0.497	0.494
Arellano-Bond test for First-order serial correlation	-2.986	-3.057	-3.076	-3.040	-3.101	-3.123
P-values	0.003	0.002	0.002	0.002	0.002	0.002
Observations	238	238	238	238	238	238
Number of villages	95	95	95	95	95	95

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table 4: System GMM estimates of the effects of emigration on income inequality in rural China

	Dependent variable: Theil index (Household total income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Theil index (Household total income)(t-1)	0.509***	0.499***	0.497***	0.502***	0.504***	0.507***
	(0.128)	(0.126)	(0.126)	(0.123)	(0.120)	(0.116)
ln (% Emigration)	0.037	0.038	0.039	0.041	0.030	0.034
	(0.034)	(0.035)	(0.033)	(0.034)	(0.031)	(0.031)
ln (% Emigration)(t-1)	-0.054*	-0.054*	-0.055*	-0.055*	-0.062*	-0.059*
	(0.031)	(0.031)	(0.032)	(0.032)	(0.035)	(0.033)
ln [% Old (age>64)]		0.010	0.010	0.009	0.005	0.010
		(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
ln [% Young (age<15)]		-0.035	-0.036	-0.038	-0.036	-0.027
		(0.028)	(0.029)	(0.029)	(0.027)	(0.028)
ln (% Han)			-0.011	-0.011	-0.006	-0.007
			(0.015)	(0.016)	(0.014)	(0.014)
ln (% Born in other Provinces)				0.005	0.007	0.003
				(0.015)	(0.015)	(0.015)
ln (% Primary)					-0.022	-0.021
					(0.033)	(0.035)
ln (% Middle school)					-0.012	-0.011
					(0.029)	(0.031)
ln (% Higher education)					-0.022	-0.025
					(0.018)	(0.020)
ln (per capita income) (RMB at year 2006)						0.043
						(0.037)
Wave 2000	-0.048	-0.028	-0.028	-0.025	-0.038	-0.022
	(0.050)	(0.058)	(0.060)	(0.059)	(0.065)	(0.057)
Wave 2004	-0.079**	-0.073**	-0.073**	-0.074**	-0.076*	-0.070*
	(0.034)	(0.035)	(0.036)	(0.037)	(0.040)	(0.037)
Hansen overidentification test	12.75	13.58	13.55	13.48	13.25	12.99
P-values	0.753	0.697	0.698	0.704	0.719	0.737
Arellano-Bond test for First-order serial correlation	-2.098	-2.227	-2.237	-2.222	-2.070	-2.505
P-values	0.036	0.026	0.025	0.026	0.038	0.012
Observations	238	238	238	238	238	238
Number of villages	95	95	95	95	95	95

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table 5: System GMM estimates of the effects of emigration on income inequality in rural China

	Dependent variable: Theil index (Per capita income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Theil index (Per capita	0.199	0.196	0.197	0.203	0.189	0.191
income)(t-1)	(0.140)	(0.140)	(0.140)	(0.135)	(0.137)	(0.134)
ln (% Emigration)	0.048*	0.049*	0.046*	0.049*	0.043*	0.045*
	(0.026)	(0.027)	(0.026)	(0.027)	(0.024)	(0.024)
ln (% Emigration)(t-1)	-0.048*	-0.048*	-0.051*	-0.052*	-0.057*	-0.055*
	(0.028)	(0.029)	(0.030)	(0.029)	(0.032)	(0.031)
ln [% Old (age>64)]		0.005	0.004	0.003	0.003	0.006
		(0.015)	(0.014)	(0.014)	(0.014)	(0.014)
ln [% Young (age<15)]		-0.021	-0.019	-0.022	-0.020	-0.014
		(0.026)	(0.027)	(0.028)	(0.025)	(0.026)
ln (% Han)			-0.010	-0.009	-0.009	-0.010
			(0.015)	(0.015)	(0.015)	(0.015)
ln (% Born in other				0.008	0.009	0.007
provinces)				(0.015)	(0.015)	(0.014)
ln (% Primary)					-0.001	-0.001
					(0.035)	(0.036)
ln (% Middle school)					0.011	0.012
					(0.028)	(0.029)
ln (% Higher education)					-0.011	-0.013
					(0.018)	(0.020)
ln (per capita income)						0.028
(RMB at year 2006)						(0.036)
Wave 2000	-0.068	-0.056	-0.061	-0.058	-0.072	-0.060
	(0.046)	(0.053)	(0.054)	(0.054)	(0.060)	(0.056)
Wave 2004	-0.063*	-0.059*	-0.061*	-0.063*	-0.069*	-0.064
	(0.035)	(0.036)	(0.037)	(0.038)	(0.041)	(0.040)
Hansen overidentification test	13.95	14.83	14.53	14.73	14.18	14.14
P-values	0.671	0.608	0.629	0.615	0.654	0.657
Arellano-Bond test for First-order serial correlation	-1.622	-1.639	-1.649	-1.638	-1.638	-1.697
P-values	0.105	0.101	0.099	0.101	0.101	0.090
Observations	238	238	238	238	238	238
Number of villages	95	95	95	95	95	95

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table 6: Fixed effects estimates of the effects of emigration on the gender wage gap in rural China

	Dependent variable: ln (male wage) - ln (female wage)					
	(1)	(2)	(3)	(4)	(5)	(6)
ln (% Emigration)(t-1)	-0.055*	-0.054*	-0.055*	-0.047	-0.055*	-0.060**
	(0.028)	(0.029)	(0.029)	(0.025)	(0.024)	(0.025)
Male migrants / total migrants (t-1)	0.196**	0.197**	0.200**	0.200**	0.196***	0.190**
	(0.066)	(0.070)	(0.070)	(0.065)	(0.055)	(0.057)
ln [% Old(age>64)]		-0.002	-0.002	-0.009	-0.002	-0.012
		(0.041)	(0.041)	(0.043)	(0.041)	(0.035)
ln [% Young (age<15)]		0.010	0.015	0.028	0.026	0.004
		(0.060)	(0.060)	(0.065)	(0.059)	(0.056)
ln (% Han)			0.088	0.073	0.070	0.071
			(0.125)	(0.118)	(0.108)	(0.125)
ln (% Born in other rovinces)				-0.046	-0.044	-0.051
				(0.034)	(0.030)	(0.028)
ln (% Primary)					-0.045	-0.065
					(0.094)	(0.093)
ln (% Middle school)					-0.188	-0.180
					(0.161)	(0.154)
ln (% Higher education)					-0.086	-0.078
					(0.116)	(0.111)
ln (per capita income) (RMB at year 2006)						-0.110
						(0.063)
Observations	219	219	219	219	219	218
Number of villages	94	94	94	94	94	93
R-squared	0.05	0.05	0.05	0.06	0.08	0.10

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table 7: Fixed effects estimates of the effects of emigration on the gender wage gap in rural China

	Dependent variable: ln (male wage) - ln (female wage)					
	(1)	(2)	(3)	(4)	(5)	(6)
ln (% Emigration)	-0.023	-0.024	-0.024	-0.023	-0.024	-0.011
	(0.053)	(0.053)	(0.054)	(0.051)	(0.055)	(0.061)
Male migrants / female migrants	0.031	0.032	0.032	0.077	0.066	0.056
	(0.084)	(0.084)	(0.084)	(0.102)	(0.101)	(0.100)
ln (% Emigration)(t-1)	-0.055*	-0.055*	-0.055*	-0.046	-0.058*	-0.062*
	(0.029)	(0.029)	(0.029)	(0.027)	(0.029)	(0.030)
Male migrants / total migrants (t-1)	0.232***	0.231***	0.234***	0.238***	0.233***	0.216***
	(0.063)	(0.064)	(0.064)	(0.063)	(0.052)	(0.053)
ln [% Old(age>64)]		-0.006	-0.007	-0.011	-0.003	-0.014
		(0.046)	(0.046)	(0.048)	(0.046)	(0.039)
ln [% Young (age<15)]		-0.013	-0.008	0.004	-0.001	-0.022
		(0.064)	(0.066)	(0.071)	(0.061)	(0.057)
ln (% Han)			0.083	0.068	0.066	0.066
			(0.115)	(0.117)	(0.108)	(0.127)
ln (% Born in other provinces)				-0.051	-0.047	-0.054
				(0.035)	(0.032)	(0.031)
ln (% Primary)					-0.043	-0.056
					(0.113)	(0.109)
ln (% Middle school)					-0.180	-0.170
					(0.186)	(0.176)
ln (% Higher education)					-0.087	-0.079
					(0.130)	(0.123)
ln (per capita income)						-0.112
(RMB at year 2006)						(0.077)
Observations	211	211	211	211	211	210
Number of villages	93	93	93	93	93	92
R-squared	0.06	0.06	0.06	0.07	0.09	0.10

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Appendix

Table A1: System GMM estimates of the effects of emigration on income inequality in rural China (lagged emigration only)

	Dependent variable: Gini coefficient (Household total income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Gini (Household total income)(t-1)	0.377***	0.379***	0.372***	0.371***	0.372***	0.373***
	(0.115)	(0.113)	(0.113)	(0.111)	(0.108)	(0.106)
ln (% Emigration)(t-1)	-0.025*	-0.027**	-0.026*	-0.026*	-0.029**	-0.028**
	(0.013)	(0.013)	(0.013)	(0.013)	(0.014)	(0.014)
ln [% Old (age>64)]		0.012	0.012	0.012	0.011	0.012
		(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
ln [% Young (age<15)]		-0.008	-0.009	-0.009	-0.013	-0.011
		(0.014)	(0.014)	(0.014)	(0.014)	(0.015)
ln (% Han)			-0.010	-0.010	-0.008	-0.009
			(0.008)	(0.008)	(0.007)	(0.007)
ln (% Born in other provinces)				-0.003	-0.001	-0.002
				(0.008)	(0.008)	(0.008)
ln (% Primary)					-0.002	-0.002
					(0.015)	(0.016)
ln (% Middle school)					-0.009	-0.010
					(0.016)	(0.016)
ln (% Higher education)					-0.010	-0.011
					(0.009)	(0.009)
ln (per capita income) (RMB at year 2006)						0.009
						(0.016)
Wave 2000	-0.031	-0.027	-0.025	-0.025	-0.027	-0.023
	(0.023)	(0.027)	(0.027)	(0.027)	(0.028)	(0.026)
Wave 2004	-0.043***	-0.043***	-0.042***	-0.041**	-0.042**	-0.041**
	(0.015)	(0.016)	(0.016)	(0.016)	(0.017)	(0.017)
Hansen overidentification test	14.81	15.65	15.63	15.69	15.12	15.03
P-values	0.675	0.617	0.619	0.614	0.654	0.660
Arellano-Bond test for First-order serial correlation	-2.883	-3.012	-2.987	-3.049	-2.907	-3.094
P-values	0.004	0.003	0.003	0.002	0.004	0.002
Observations	249	249	249	249	249	249
Number of villages	98	98	98	98	98	98

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.

Table A2: System GMM estimates of the effects of emigration on income inequality in rural China (contemporary emigration only)

	Dependent variable: Gini coefficient (Household total income)					
	(1)	(2)	(3)	(4)	(5)	(6)
Gini (Household total income)(t-1)	0.285***	0.274***	0.276***	0.275***	0.280***	0.280***
	(0.102)	(0.101)	(0.101)	(0.101)	(0.099)	(0.098)
ln (% Emigration)	0.015	0.015	0.013	0.012	0.013	0.014
	(0.017)	(0.017)	(0.016)	(0.016)	(0.017)	(0.017)
ln [% Old (age>64)]		0.019**	0.019**	0.019***	0.021***	0.022***
		(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
ln [% Young (age<15)]		-0.018	-0.018	-0.017	-0.015	-0.011
		(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
ln (% Han)			0.003	0.002	-0.001	-0.002
			(0.008)	(0.008)	(0.008)	(0.007)
ln (% Born in other provinces)				-0.002	-0.002	-0.003
				(0.008)	(0.008)	(0.008)
ln (% Primary)					-0.004	-0.004
					(0.012)	(0.013)
ln (% Middle school)					0.024**	0.024**
					(0.010)	(0.010)
ln (% Higher education)					0.004	0.003
					(0.007)	(0.008)
ln (per capita income) (RMB at year 2006)						0.014
						(0.013)
Wave 2000	0.028	0.025	0.026	0.026	0.021	0.019
	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)
Wave 2004	0.011	-0.003	-0.001	0.000	-0.004	-0.008
	(0.021)	(0.023)	(0.022)	(0.023)	(0.023)	(0.024)
Wave 2006	0.035	0.019	0.021	0.022	0.018	0.013
	(0.024)	(0.027)	(0.026)	(0.027)	(0.028)	(0.029)
Hansen overidentification test	20.07	21.65	21.74	21.89	21.28	21.03
P-values	0.638	0.541	0.536	0.527	0.564	0.579
Arellano-Bond test for First-order serial correlation	-3.771	-3.663	-3.693	-3.728	-3.732	-3.874
P-values	0.000	0.000	0.000	0.000	0.000	0.000
Observations	327	327	327	327	327	327
Number of villages	98	98	98	98	98	98

Robust standard errors clustered at community level in parentheses; * significant at 10%; ** significant at 5%; *** significant at 1%; Survey years (dummies) are included in all specifications.