
**Aggregation and Threshold Models of Intersectoral Labor Migration: Evidence from the United States and Japan**

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# Aggregation and Threshold Models of Intersectoral Labor Migration: Evidence from the United States and Japan*


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**Abstract**

Models of intersectoral labor migration focus on expected wage differentials as the primary cause of migration from one sector to another. Empirical applications typically assume that migration occurs as soon as the return differential exceeds Marshallian migration costs, but recent work has focused on embedding the migration decision in a real options framework. If migrants consider real options, then option value is incorporated in the migration decision as an additional cost. While some authors argue that real options imply threshold behavior in aggregate migration equations, we show that there is limited theoretical justification or empirical support for this conclusion. Using data from the United States and Japan, we find little evidence of non-linearities in the aggregate migration equations. The non-linear models generate unrealistically large elasticities on returns to labor. Our results suggest that linear migration equations are sufficient for capturing key features of aggregate intersectoral migration and support the importance of the Marshallian trigger in the migration decision.


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## Introduction

Empirical models of intersectoral labor migration assume that labor migrates when wage differentials between sectors exceed a certain threshold (Harris and Todaro, 1970; Mundlak and Strauss, 1978). Such models are predicated on an economic model of agent behavior where agents consider the expected utility derived from working in mutually exclusive sectors (Larson and Mundlak, 1997; Barkley, 1990). Several studies have suggested that agents also consider real options when deciding whether to migrate (Richards and Patterson, 1998; Dennis and İşcan, 2007; Önel and Goodwin, 2014). Real options imply a discontinuity - or zone of inaction - in the migration behavior of the agent. The value of waiting adds an additional cost to the agent's reservation wage.

Even if individual agents are subject to threshold behavior, in the sense that the migration decision is a binary choice, aggregate migration equations at the sector level will not necessarily exhibit the same properties. Similar issues around aggregation and functional form have been noted in other economic contexts (Houthakker, 1955; Berck and Helfand, 1990). Because aggregation is a feature of most empirical analyses in intersectoral migration, this raises the question of whether empirical specifications incorporating non-linearities or thresholds provide improved inference. Nor are there clear theoretical criteria for selecting among alternative specifications. These problems are compounded by the important role that specification choice and functional form can play in the magnitudes and statistical significance of economic parameters of interest.

The contributions of this article are threefold. First, we show that consideration of real options by agents will not generally induce a jump threshold in the aggregate migration equation. There is little theoretical justification for the use of threshold models in this setting. Second, we empirically test for non-linear behavior in the aggregate migration equation. We then compare the economic implications of a variety of linear and non-linear models. There is little empirical evidence of lon-linearities. Lastly, we present results from both the United States and Japan for all farm labor and hired farm labor. Recent studies suggest that stronger non-linearities may be observed for hired labor. The U.S. and Japan provide interesting comparative settings as the agricultural sectors in the countries have distinct structures and face different economic conditions.

U.S. agriculture has become highly mechanized, increasingly concentrated, and more efficient in the last 80 years. Yields for many crops have increased more than three times. Against this backdrop, the U.S. agricultural sector has seen a major release of labor. Using aggregate data for the U.S., Barkley (1990) showed that the supply of agricultural labor in the U.S. between 1940 and 1985 was responsive to both the relative returns to labor and land values. Similar findings appear in Emerson (1989) and Perloff (1991) who considered migratory farm labor in Florida and workers surveyed in the Current Population Survey, respectively. In all cases, U.S. farm labor migrated in response to expected differentials in the returns to labor.

Farm labor from Mexico has historically been a major source of hired labor for U.S. farms. Recent work on migrant labor has attracted attention in the U.S. given increasing wage rates in the H-2A program. Economic development in Mexico has led to decreased supply of migrant labor (Charlton and Taylor, 2016). Taylor et al. (2012) suggest that the U.S. farm sector will need to adjust to tightening labor markets. Evidence of labor shortages for certain types of workers in California was noted by Richards (2018). While we do not distinguish between domestic and international labor in this study, patterns of economic growth occurring in Mexico, and their relation to U.S. farm labor markets, appear to follow general trends in labor release discussed by Mundlak (1978).

In Japan, land reform following the Second World War caused declines in land tenancy, an increase in owner cultivators, and is widely thought to have resulted in increased productivity (Kawagoe, 1999). At the same time, the wider Japanese economy experienced rapid economic growth; growth in gross domestic product averaged well over five percent per year in the 1960s and as high as nine percent in the late 1980s. Wage differentials between the agricultural and non-agricultural sectors were large and labor moved from the agricultural sector to non-farm employment in what Mundlak (1979) termed intersectoral labor migration.

Early work on labor migration using aggregate data from Japan was conducted by Minami (1967) and Minami (1968). His contributions were followed by Mundlak and Strauss (1978) who showed that, indeed, income differentials affect migration and that such a conclusion can only be reached through empirical validation. There was healthy debate in the 1970s as to whether the miracle of Japanese growth was achieved primarily through flows in labor from agriculture to non-agricultural sectors or in flows of savings; Mundlak (1979) demonstrated that the effects of intersectoral labor flows dominated those of savings flows. Recently, Takayama et al. (2020) showed that direct government payments to agriculture had the effect of preventing farmland abandonment and farm exit, but that such effects were relatively modest; similar conclusions were reached by Ito et al. (2019). These findings contrast with D'Antoni et al. (2012) who found that government payments increased out-migration in the United States.

While much of the literature on intersectoral labor migration assumes that households devote their labor to a single sector, a large literature has also studied the behavior of households with one foot in agriculture and one foot out. This is particularly important given the now high number of farm households with off-farm income in both the U.S. and Japan. Kimhi and Lee (1996) and Ahituv and Kimhi (2002) showed, using data from Israel, that the off-farm work decisions in a household are affected by the composition of the household and exhibit state dependence. At least in the U.S., greater off-farm work was associated with reduced farm efficiency (Goodwin and Mishra, 2004) and government payments decreased off-farm work (Mishra and Goodwin, 1997). Goodwin and Holt (2002) suggested that, much like aggregate migration equations, off-farm labor supply equations are subject to possible misspecification with standard approaches. Although there is less research on farm labor decisions in Japan, Ramsey et al. (2019) found that intersectoral migration was related to the share of farm income relative to household income for part-time farm households.

There is now little doubt that wage differentials affect labor migration. However, recent work by Dennis and İşcan (2007) and Önel and Goodwin (2014) casts doubt on the suitability of linear models as statistical representations of migration behavior. They develop a real options approach to the migration decision and suggest that nonlinear models - particularly threshold regressions - may provide a more accurate description of aggregate migration. While individuals agents may consider real options, we argue that this concept is not generally testable in an aggregate migration equation.[1] A threshold in the aggregate migration equation is not an implication of the real options model. Furthermore, the specification of the migration equation can have major impacts on parameter estimates of economic interest.

[1]Real options have been considered in agricultural labor markets and are a central feature of Richards and Patterson (1998). However, the existence of real options is empirically assessed by examining wage distributions. Similarly, real options can be incorporated into search models of the type found in Richards (2018). In either case, the key point is that the wage gap induced by real options is not measured through an aggregate migration equation.

## Marshallian and Real Options Frameworks

Early research in intersectoral migration did not specify a theory of migration at the household level. Instead, empirical analysis began from the migration equation (referred to here as the aggregate migration equation) which is an empirical construct (Harris and Todaro, 1970; Mundlak and Strauss, 1978; Mundlak, 1978). In spite of its purely empirical historical underpinnings, the aggregate migration equation has been shown to be broadly consistent with household-level models of migration. The link between household-level and aggregate equations is usually in the form of an index function that "counts" household behavior and aggregates household decisions to the level of the aggregate migration equation. Key underlying features of this aggregation are the distributions associated with the households being aggregated.

A household-level model is developed to varying degrees in Barkley (1990), Larson and Mundlak (1997), and Mundlak (1979). These models are all partial equilibrium as they do not take into account all parts of the economy. Let the indirect utility for the household be given by

$$v(g,z,j) = v(p_j, w_j, g, z, c_j(d_j, g, z, y) \tag{1}$$

where $g$ is the individual's age, $z$ are individual demographics, $j$ indexes over non-farm or farm employment, and $c_j(\cdot)$ is a term giving the cost of migrating from one sector to another. In addition, $w_j$ is the expected wage in occupation $j$, $p_j$ is the price of consumer goods, $d_j$ is the distance to other employment, and $y$ is the general level of infrastructure development in the country. Obviously, there are no costs if the household remains in the same sector. The discounted stream of utility is then given by

$$V(g,z,j) = \int_g^T \exp(-\rho t) v(g,z,j)(t) dt \tag{2}$$

so that an individual migrates from a farm to non-farm occupation when

$$V(g,z,nf) > V(g,z,f) \tag{3}$$

where $nf$ denotes the non-farm sector and $f$ denotes the farm sector. The expected signs of the partial derivatives of equation 1 with respect to its inputs are provided in Larson and Mundlak

(1997). What bears repeating is that equation 1 is increasing in wage (or income) and decreasing in costs of migration. Therefore, an individual migrates according to equation 3 when, ceteris paribus, their wage from the non-farm sector rises relative to their wage from farming.

To arrive at the aggregate migration equation, an index function collects households so that

$$(V_i(nf) - V_i(f))h_i(f, nf) \geq 0 \tag{4}$$

where $h_i(\cdot)$ takes the value of 1 if the term in parentheses is positive. That is, it takes a value of 1 when the return from migrating to the non-farm sector is positive. A similar result holds for migrants from the non-farm sector to the farm sector. Summing over the farm and non-farm labor forces gives the total number of migrants which can be written as

$$M(f, nf) = \sum_{i=1}^{L_f} h_i(f, nf) - \sum_{i=1}^{L_{nf}} h_i(nf, f) \tag{5}$$

As explained in Mundlak (2000), the migration equation is a function of the arguments of the indirect utility functions in the sectors as well as the size of the labor force in the origin and destination.

The real options approach for intersectoral migration is developed in Önel and Goodwin (2014), although the concept was discussed much earlier in the agricultural labor context by Richards and Patterson (1998). Consideration of real options recognizes that migrants may not switch sectors even in the face of a positive wage differential. This situation arises because individuals find value in waiting to switch sectors. Waiting has value because it reduces risks over time. Consider the Larson and Mundlak (1997) model adjusted to incorporate the options value in Önel and Goodwin (2014). Individuals not only consider $V(g, z, j)$ but also the option value of waiting. The individual migrates from farm to non-farm occupations when

$$V(g, z, nf) > V(g, z, f) + O(g, z, f) \tag{6}$$

where $O(g, z, f)$ is the option value of waiting given that the individual is in the farm sector.

Accordingly households can be collected by an index function

$$(V_i(nf) - V_i(f) - O_i(f))h_i(f, nf) \geq 0 \tag{7}$$

Equation 4 is modified to accommodate the presence of real options. The indexing function serves the same purpose but with an additional term incorporating the options value. For simplicity, we will only consider farm out-migration from here forward. What implications does the addition of the real options term in equation 7 have on the aggregate migration equation? Specifically, what implications are there for the conditional relationship between migration rates and the returns to labor?

This question was not completely ignored in earlier work, although it appears to have been only lightly treated. Larson and Mundlak (1997) estimate a non-linear aggregate migration equation that includes a wedge. With a wedge, migration does not stop when the Marshallian wage ratio is unity. The wedge essentially allows for households to consider real options in their migration decisions. There are a number of other factors besides option value that can generate the wedge in the Larson and Mundlak (1995) framework. Even in their basic household model, the aggregate migration equation cannot be used to test exclusively for consideration of real options by economic agents. Empirically, a key finding of Larson and Mundlak (1997) is that the estimated value of the wedge is small and statistically insignificant. In essence, it is the Marshallian effects that matter.

Consider the statistical migration equations found in Barkley (1990) and Önel and Goodwin (2014) which are are based on linear Harris-Todaro type models and discussed in more detail below. Without loss of generality, normalize the entire set of workers to unity. Suppose that the only terms entering equation 6 are the wages and the option values so that we can dispense with age, demographics, etc. Then equation 5 can be written as a continuous function with

$$M(f, nf) = \int_0^r f(r)dr \tag{8}$$

where $r$ is the reservation wage accounting for the option value and $f(\cdot)$ is the density function of reservation wages in the population. The relationship between the reservation wage and migration is then given by the cumulative distribution function. The distribution function, which may exhibit nonlinearities, is unlikely to exhibit a jump discontinuity. A uniform distribution of the reservation wage will clearly have a linear distribution function. A normal distribution of the reservation wage will have a nearly linear distribution function; at least in a large range around the mean reservation wage. In any event, under reasonable assumptions about the distribution of reservation wages, the aggregate migration equation will not be discontinuous or exhibit a threshold.

A similar conclusion can be reached if we allow the migration decision to follow a latent variable model at the individual level. Let $M^*$ be the latent variable with household migration $M = 1$ if $M^* > 0$. For simplicity, we consider only the Marshallian reservation wage resulting in a model such as

$$M* = \beta_0 + \beta_1 r + \epsilon \tag{9}$$

where $\beta_1 > 0$ and $\epsilon$ is distributed standard normal (without loss of generality). Then the probability $p(M = 1) = F(\beta_0 + \beta_1 r)$ where $F(\cdot)$ is the standard normal distribution function. This implies that the migration function is a continuous and concave function of the reservation wage. Therefore, the aggregate migration equation will be nearly linear (depending on the value of $\beta_1$) and will not exhibit any jump over the distribution of the reservation wage.

Applying a similar discussion to the real options approach, allowing for option value does not imply a discontinuity in the migration equation. It simply implies a change in the definition of $r$ in equation 9 from the Marshallian reservation wage to one that includes option value. Incorporating the value of waiting increases the value at which $M^*$ becomes positive. That is, it increases the value of the argument to the standard normal distribution function. Introducing real options value increases the condition for the migration decision, or threshold value, in the latent variable equation. It does not imply threshold behavior in the aggregate migration equation.

Our point is that the conditional relationship between returns to labor and migration is not likely to exhibit a discontinuous threshold because reservation wages follow their own continuous distribution. Option values also follow their own distributions. Or rather, the individual threshold behavior in equations 3 and 6 does not imply that the aggregate equation exhibits a jump threshold. A related point is made in Berck and Helfand (1990) in the context of aggregate production functions. In that case, non-linear individual production functions did not necessarily imply aggregate production functions of the same functional form.

Dennis and İşcan (2007) consider agricultural outmigration in a general equilibrium model. They present results both with and without migration costs. A key point in their discussion of migration with costs (or zones of inaction) is that there will be periods during which *some* workers do not migrate even in the face of a positive wage differential. The maximum wage gap - for which the worker will not migrate - is the same for all workers in their model. Dennis and İşcan (2007) assume identical workers and a constant population. Introducing distributions of migration costs, which the authors hint at in their discussion, results in some migration. While Önel and Goodwin (2014) empirically test for the present of non-linearities in aggregate migration equations, they do not provide a theoretical justification for such a test.

## Econometric Approaches to Intersectoral Labor Migration

Ultimately, the presence of non-linearities in aggregate migration equations can be considered an empirical issue even if real options do not provide a theoretical justification. Non-linearities could result from other structural changes in agricultural labor markets. There is still value in estimating a wide variety of model specifications to better understand the implications of researcher model choice on empirical estimates. The threshold regression motivated in Önel and Goodwin (2014) is an extension of the Harris-Todaro human capital migration model. There is an extensive literature on the use of Harris-Todaro models in agricultural economics and related issues of variable specification and measurement; problems of variable measurement can be non-trivial in these settings. Although we spend some time discussing those issues here, our statistical models do not depart from previous work in this area. The primary difference is choice of the likelihood function where we also consider the regression kink model which admits a continuous threshold change in the effect of wage gaps on migration.

The basic Harris-Todaro type model for the migration rate from the farm sector to the non-farm sector is

$$m_t = \theta_0 + \theta_1 r_{t-1} + \theta_2 u_{t-1} + \theta_3 g_{t-1} + \theta_4 l_{t-1} + \epsilon_t \tag{10}$$

where $m$ is the migration rate, $r$ is a measure of returns to labor in the sectors (the ratio of value added per worker or the ratio of hourly wage rates non-farm to farm), $u$ is the unemployment rate in the non-farm sector, $g$ is the relative labor force non-farm to farm, $lv$ is the value of agricultural land, and $\epsilon$ is a random error. Possible endogeneity is mitigated by using lagged values of all explanatory variables in the model. The model is estimated using ordinary least squares or maximum likelihood estimation.

Önel and Goodwin (2014) propose a threshold model to capture non-linear behavior in migration based on Hansen (1996) and Hansen (2000) who developed inferential procedures for such models. Their model allows all parameters to vary across regimes with the threshold depending on the value of the returns to labor. We consider a similar threshold model given by

$$m_t = (\gamma_0^0 + \gamma_1^0 r_{t-1}) I(r_{t-1} > \tau) + (\gamma_1^0 + \gamma_1^1 r_{t-1}) I(r_{t-1} \leq \tau) + \theta_2 u_{t-1} + \theta_3 g_{t-1} + \theta_4 l_{t-1} + \epsilon_t \quad (11)$$

with the variables as defined above. In this case the intercept and slope coefficient on the returns to labor are allowed to switch between regimes, but the other parameters are regime invariant. Note that $\tau$ is the estimated threshold value.

The regression kink model, conceptually quite similar to the threshold regression, is developed in Hansen (2017). An earlier example is given in Chan and Tsay (1998). As opposed to the threshold regression, where a discontinuity is assumed, the regression kink function is continuous with a change in slope at the threshold point. Hansen (2017) suggest that the regression kink approach is most suitable in cases where there is no reason to expect a discontinuity and the threshold only occurs in one variable. Fashioning the migration equation in the form of the regression kink results in

$$m_t = \beta_1 (r_{t-1} - \lambda)_- + \beta_2 (r_{t-1} - \lambda)_+ + \theta_2 u_{t-1} + \theta_3 g_{t-1} + \theta_4 l_{t-1} + \epsilon_t \quad (12)$$

Here the notation $(\cdot)_-$ and $(\cdot)_+$ denote $\min(\cdot, 0)$ and $\max(\cdot, 0)$ respectively. The regression kink model is a restricted or nested version of the threshold regression.

As noted in the preceding section, even if household behavior follows the theory of real options, it is unlikely that the relationship between migration and returns to labor will exhibit a discontinuous jump. Therefore, from a theoretical perspective, we suggest that the regression kink model is more appropriate among the two threshold models. Given that the regression kink is nested by the threshold regression, it may appear that the differences between the two are likely to be trivial. But existing literature on the two threshold models has assumed either discontinuity or continuity at the threshold point. As noted in Hansen (2017), testing for continuity is a difficult problem.

In the application to follow, we estimate linear and threshold (continuous and discontinuous) models applied to a variety of returns and labor metrics for both the U.S. and Japan. While both the threshold and regression kink approaches can be modified to include nonlinear response within regimes, we use linear regression segments. This choice is motivated by the moderate sample sizes encountered in aggregate studies of migration. Our results show that the choice of model can have large impacts on the estimated responsiveness of migration to returns to labor. Moreover, we find little evidence to support the use of the threshold models

## Results and Discussion

Our data are drawn from several sources in the United States and Japan with summary statistics for each variable shown in table 1. The time series runs from 1955 to 2017 for the United States and 1958 to 2017 for Japan. Based on results from Önel and Goodwin (2014), we collected data representing all farm labor (including unpaid family labor, hired labor, etc.) as well as hired farm labor. In the Japanese data, the hired farm labor are simply referred to as employees. We also consider two definitions of the returns to labor as the returns measure used in the analysis is likely to affect any tests of threshold behavior; returns to labor are measured as value added per worker and hourly wage rates.

Agricultural labor data in the United States is taken from the Economic Report of the President (ERP) which included - in Table B-100 - measurements of all farm labor and hired farm labor up to 2011. This table has since been eliminated from the ERP. The Japanese data are taken from the Labour Force Survey conducted by the Statistics Bureau of Japan for the agriculture and forestry industry. Non-farm employment statistics for the U.S. are total, private employment taken from the Current Employment Statistics Survey of the Bureau of Labor Statistics while the non-farm labor force in Japan is taken from the Labour Force Survey for non-agricultural industries.

Returns to labor are measured as value added per worker and hourly wage rates. Value added for the non-farm sector is that of private industry minus value added in the farm sector. Both of these statistics are taken from the Bureau of Economic Analysis Gross Domestic Product by Industry tables. In this case, the farm sector includes agriculture, forestry, fishing, and hunting.

The Japanese data are taken from the National Accounts of Japan published by the Cabinet Office. Hourly wage rates for Japan are taken from the Statistical Survey on Commodity Prices in Agriculture while hourly wage rates for U.S. hired farm workers are taken from Farm Labor Survey compiled by the U.S. Department of Agriculture.

Nominal land values for the U.S. were obtained from the National Agricultural Statistics Service while those for Japan were taken from the Japan Real Estate Institute. These are converted to real values using producer price indices (PPIs) for farm goods. The PPI for U.S. producers is taken from the Bureau of Labor Statistics Producer Price Indexes Database while the Japanese PPI is taken from the Ministry of Agriculture, Forestry, and Fisheries. Unemployment rates for the non-farm sector come from Federal Reserve Economic Data and Japan's Labour Force Survey. The Japanese unemployment rate includes the agricultural sector; Japan has not historically delineated non-farm and farm employment rates.

Figure 1 shows labor shares by sector and migration rates of all and hired farm labor. The farm sector comprises a decreasing amount of all labor with farm labor's share less than two percent of all labor in the U.S. and less than four percent of all labor in Japan at the end of their respective series. The share of farm labor in Japan was substantially larger than in the U.S. in the immediate post-war period (roughly double) and experienced a more rapid decline. In both countries, migration rates for hired labor are generally larger in magnitude than migration rates for all farm labor. In addition, the migration rate for all farm labor in Japan is positive over the entire time series with much less variability than the migration rates for either of the U.S. series or hired farm labor in Japan.

Plots of the two measures of returns to labor, shown in figure 2, show a relatively stark difference between Japan and the United States in the postwar period. Unlike the U.S., where value added per worker has generally increased in both sectors, the gap between the non-farm and farm sectors in Japan is relatively large. Through the 1970s and 1980s, the returns to labor grew slowly in the Japanese agricultural sector. Also in contrast to the U.S. experience, value added per worker in both sectors has been nearly flat, and declining in the farm sector after the mid-1990s.

Results for value added per worker are reflected in wage rates which have also stagnated for both non-farm and farm workers in Japan. One might expect wages and productivity to move in similar directions absent any frictions that prevent adjustment (Becker, 1962). A sharp decline in growth rates of agricultural labor productivity in the second half of the twentieth century was documented by Kuroda (1995). Ruttan (2002) notes that as the agricultural labor force declines, consolidation in farms results in a rise in the ratio of land to labor and thus in a rise in labor productivity. The lack of increased labor productivity and wages in the agricultural sector in Japan may be attributable to frictions in the market for agricultural land that prevent consolidation (Kawagoe, 1999). Land fragmentation has been cited as a major barrier to efficiency in Japanese agriculture (Kawasaki, 2010). From a comparative point of view, the lack of farmland consolidation in Japan is a major difference with the environment in U.S. agriculture.

Plots of the observed migration rates against lagged measures of returns to labor are shown in figure 3. A locally weighted scatterplot smoothing (LOESS) curve is also shown. Following Mundlak and Strauss (1978), we expect to find positive relationships; these are observed for all specifications in the United States. Discontinuities in the plots may be evidence of threshold behavior, but should only be considered suggestive. And as noted above, it does not constitute evidence of real options as other structural changes can drive similar empirical results. Some apparent non-linear behavior is seen in the plot of hired labor on wage rates. However, it is not clear whether the empirical relationship would best be described by a discrete threshold model or kinked regression. Nor do these visualization account for the impacts of other related variables.

We also observe possibly negative relationships between migration and the returns to labor in the Japanese data. Lagged wage ratios in Japan are on average much smaller than those observed in the U.S. The most extreme wage ratios are visually associated with decreased migration rates (although still positive in the case of all farm labor). As mentioned above, the plots are only suggestive of non-linearities and do not imply the use of any specific non-linear model.

We estimate the linear regression and two threshold regressions for both countries and for three different specifications of the migration variable and returns to labor. Table 2 shows the regression of the migration of all farm workers on the log ratio of value added per worker. Tables 3 and 4, respectively, contain regressions of all farm workers and hired farm workers on the log ratio of wages. All of the linear regressions, save the regression of Japanese hired labor on wages, have the expected positive sign on returns to labor. Farmers are responsive to returns to labor and migrate away from the farm sector when a positive returns to labor differential exists between the non-farm and farm sectors.

The coefficient on the unemployment rates bears the expected negative sign for the linear models of all farm labor in both the U.S. and Japan. The effect is much larger and statistically significant in the Japanese case. Although not statistically significant at conventional levels in either case, the coefficient on relative labor force is negative for the U.S. and positive for Japan. The coefficient on real farmland values is not significant in either case, but the estimate is economically large for the the U.S. Önel and Goodwin (2014) and Barkley (1990) note that there are two effects at play with real land values. The difference in estimated coefficients in the U.S. and Japan may reflect frictions in the land market which make it costly buy, sell, and rent out land. These frictions are likely to play a greater role in Japanese agriculture (Kawagoe, 1999).

We find similar results when regressing migration rates on wages. When regressing all farm workers, the coefficient on the wage ratio is not statistically significant at conventional levels. Of course, we expect that all farm labor responds only in part to wage differentials as compared to hired labor. Indeed, this is the case for hired farm workers where the linear model shows a large and significant response in the U.S. The linear model for Japan leads to a curious result in that the coefficient on the wage ratio is large and negative. We address this finding shortly after discussing the threshold regressions.

Table 5 shows the elasticities for the estimates equations evaluated at the sample means. For both the threshold regressions and the kink regressions, the elasticities are evaluated at the sample means for the respective regression segments. In terms of the linear regressions, the elasticity for all farm labor is similar for the United States and Japan with respect to value added. The estimated elasticity of 1.95 in the United States is lower than 3.25 reported in Önel and Goodwin (2014) and 4.5 reported by Barkley (1990). However, we are using a larger sample with more recent data.

The threshold results are reported in the same tables for both the jump threshold and kink regressions. We conducted non-linearity tests following Hansen (2017). The test is a bootstrap test of the linear model versus a threshold. The test does not distinguish between the jump threshold and kink regressions but is robust to heteroskedasticity. We took 10,000 bootstrap replications in conducting the tests with trimming percentage of 15%.. The p-values for the tests are reported in table 6. While there is some evidence of non-linear behavior, the p-values do not support the threshold models at conventional levels of statistical significance.

Examining the elasticities from the threshold and kink regressions in table 5, we find that they are in general much larger than the elasticities from the linear regressions. In some cases, the elasticities are unrealistically large. In the case of all labor in Japan regressed on the wage ratio, the threshold regression actually implies negative elasticities which are at odds with the expected sign and the results of the linear regression. Taken together with the nonlinearity tests, there is little support for the use of threshold regressions as statistical models for aggregate intersectoral migration.

There are several summary findings with respect to the threshold models. First, the threshold models generally support the Harris-Todaro hypothesis with the exception of the Japanese models that use wage differentials. In some cases, the coefficients on the threshold models are significant where the linear models are not. Second, signs of coefficients on auxiliary variables are almost uniformly the same as those in the linear regressions. Therefore, the linear and threshold regressions are generally in agreement except for the magnitudes (and in some cases statistical significance) of the coefficients. Lastly, agricultural labor is more responsive to wage rates in the U.S. case, but less responsive in the Japanese case.

The last point bears some discussion. The results for U.S. hired labor are in line with those of Önel and Goodwin (2014) in the sense that hired labor is more responsive than all labor. We find the opposite case for Japan, and in fact there is a negative relationship for hired labor. We suggest that this discrepancy may stem from potentially several sources. First, it may be that the statistical models do not adequately account for structural change in Japanese agriculture. Most observations with low migration rates and high wage ratios (thus driving the negative estimates) are from the first decade of the sample. When we split the sample between years prior to the Asian Financial Crisis and those after, we found visual evidence of a clear positive relation between lagged productivity and migration prior to 1990. The positive relationship is unclear after 1990. Second, given the moderate sample sizes in this analysis, it is possible that a handful of outliers are causing the large negative estimates.

Lastly, the relative lack of response to the wage rate could also be related to the nature of agricultural employment in Japan. Unlike U.S. farm labor, where labor markets are relatively open and the labor force includes large numbers of migrant laborers, the Japanese agricultural labor force is much smaller and more highly regulated. This may introduce frictions into labor markets such that labor does not respond quickly enough to changing wage rates to be captured by our one lag model. It could also be that agriculture labor decisions in Japan are also driven more by non-pecuniary factors. Both of these considerations may affect the results for hired labor in Japan.

## Conclusion

Economic development occurs alongside out-migration from the farm sector. At the household level, migration results from differences in wage rates and, ultimately, differences in net present values of working in a specific sector. Recent research has suggested that households also consider option value and sunk costs when deciding whether to migrate (Önel and Goodwin, 2014). This leads to household decision functions where households may wait (beyond a positive Marshallian wage differential) to migrate. We investigate these issues in the context of both U.S. and Japanese agriculture with focus on all agricultural labor and hired agricultural labor.

We allow for possible non-linear relationships in intersectoral labor migration. The non-linear behavior is captured by both a threshold regression and a kinked regression (Hansen, 2017). We show that 1) even if households consider real options, the aggregate migration equation will not generally exhibit a discrete jump threshold; 2) there is little empirical evidence for the use of threshold regressions in our settings; and 3) the responsiveness of migration to returns to labor depends crucially on model specification. Our results ultimately support the empirical approach suggested by Mundlak (1979); it is the Marshallian returns to labor that are most important in determining intersectoral labor migration.

From a comparative point of view, distinctions between U.S. and Japan intersectoral migration raise several important questions. While elasticities on returns to labor have the correct signs in the U.S. case, they are negative for Japanese hired labor. We suggest that this peculiarity likely arises from unique structural features of Japanese agricultural labor markets. It may be that structural changes have occurred that are not accounted for in our models. This might serve as impetus for more detailed examinations of structural change in Japanese agriculture. Particularly, why have productivity gaps between non-farm and farm sectors remained so large following the collapse of the bubble economy in the early 1990s?

We have made extensions to existing partial equilibrium models to demonstrate that the aggregate migration equation will not contain a jump threshold. It may be useful to extend such an analysis to a general equilibrium framework and also specify the time path and uncertainty in wages. As all models are only approximations to reality, such extensions would also need to make assumptions that would have implications for empirical specification. For instance, whether the population is constant, growing, or shrinking over time. Nonetheless, a detailed agent-based model may provide some additional insight into this empirical issue.

Detailed household level panel data would allow for proper measurement of real options. Unlike aggregate data, individual panel models would not be subject to randomness in the distribution of reservation wages which ultimately weakens any threshold behavior in aggregate models. Unfortunately, most data sets on agricultural workers only track individuals while they are working in the agricultural sector. Individuals are usually not followed once they stop working in agriculture, or if they change locations. From a data collection point of view, it would be helpful if individual surveys began to follow agricultural workers after they leave the agricultural sector. Some studies are already doing so (Dustmann et al., 2020).

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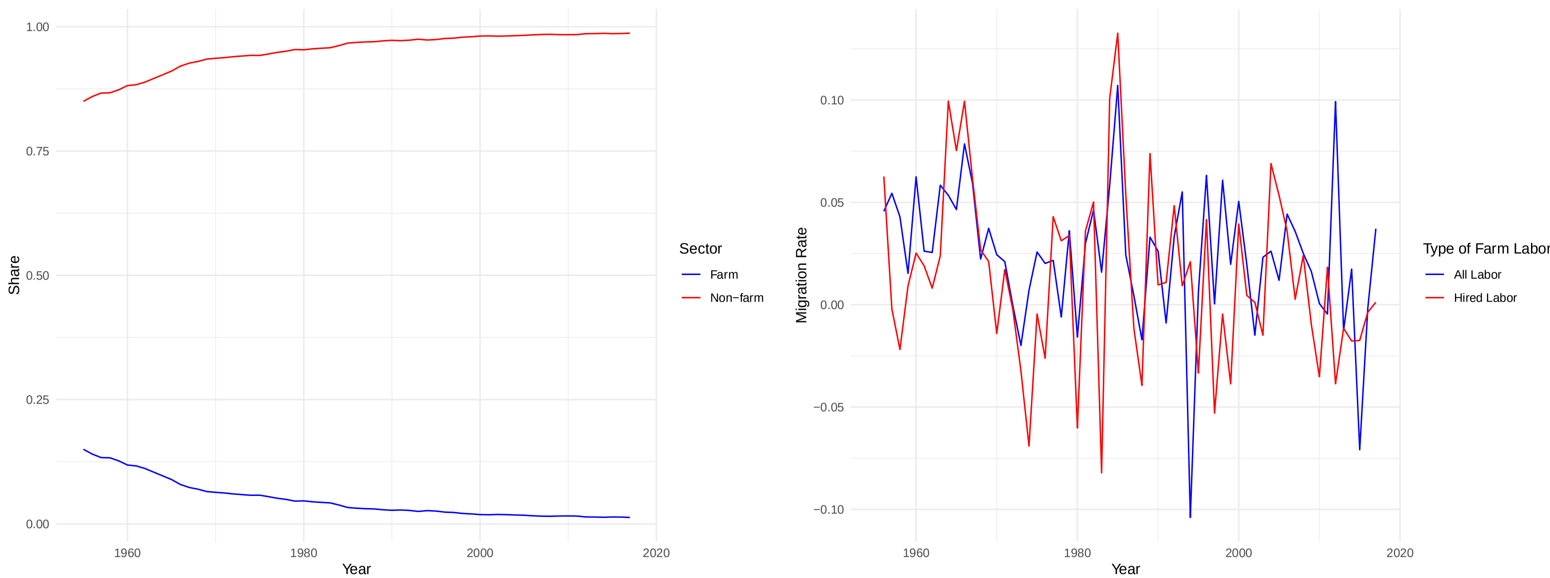
Figure 1: Labor Shares and Migration Rates

(a) United States




(b) Japan

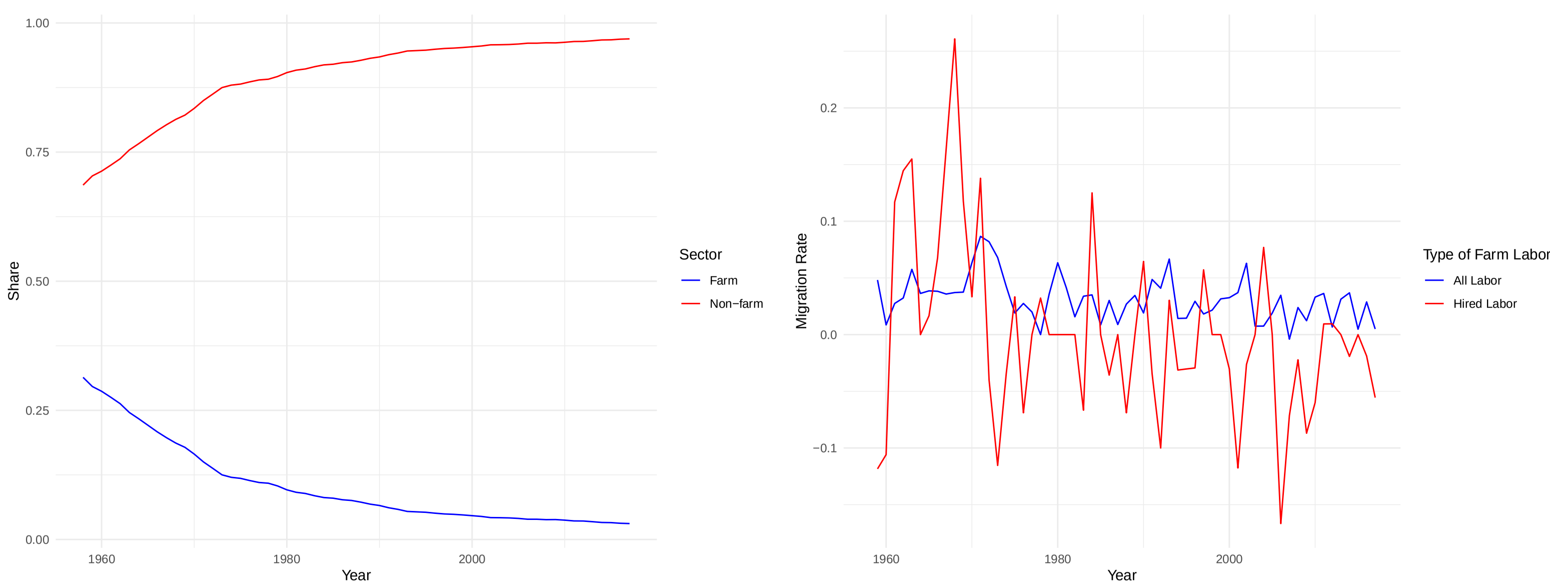
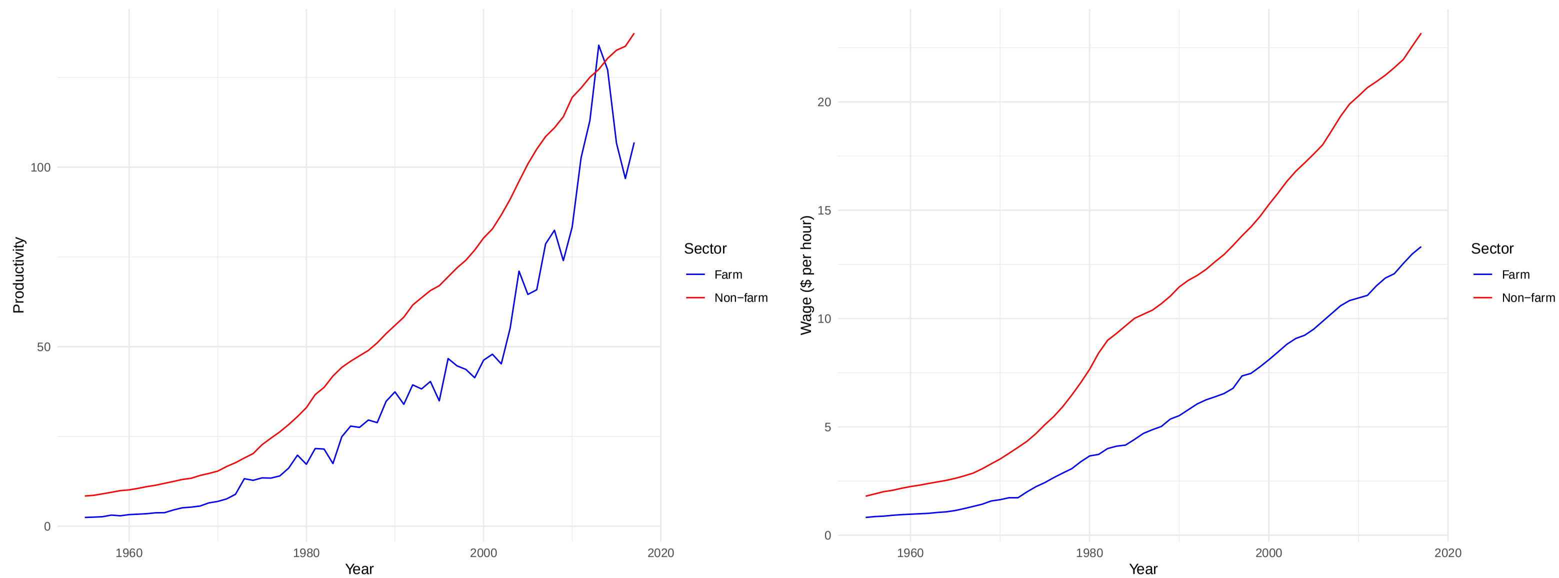


Figure 2: Returns to Labor

(a) United States




(b) Japan

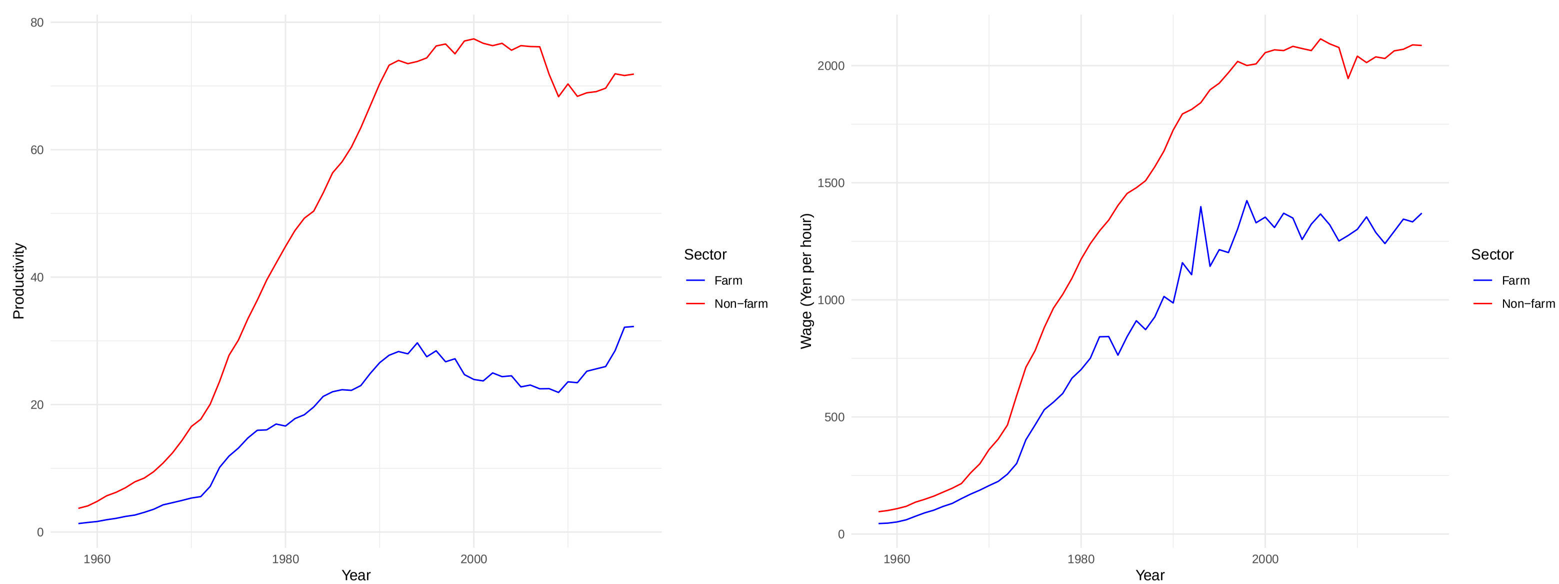
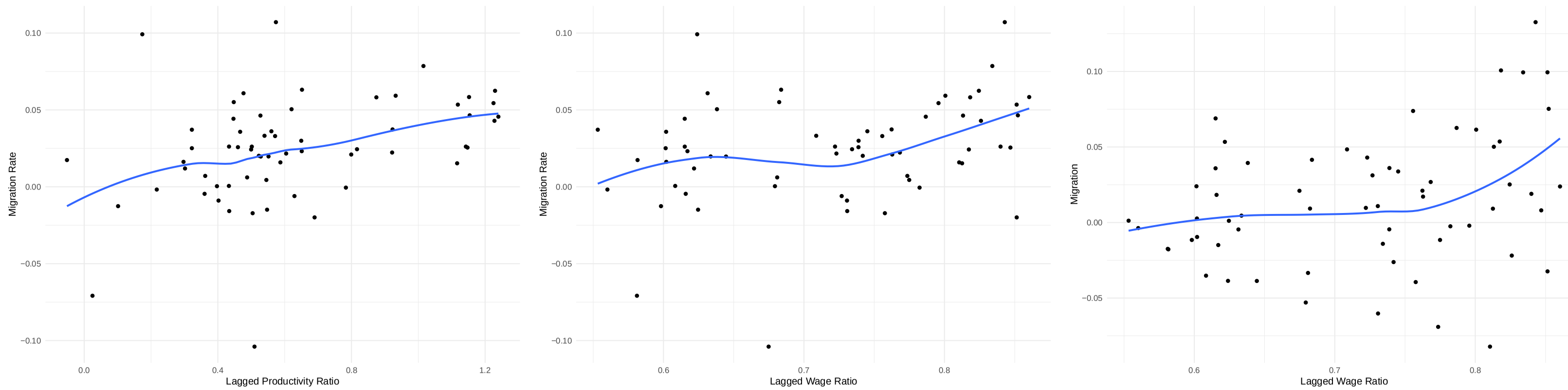


Figure 3: Migration and Returns to Labor

(a) United States




(b) Japan

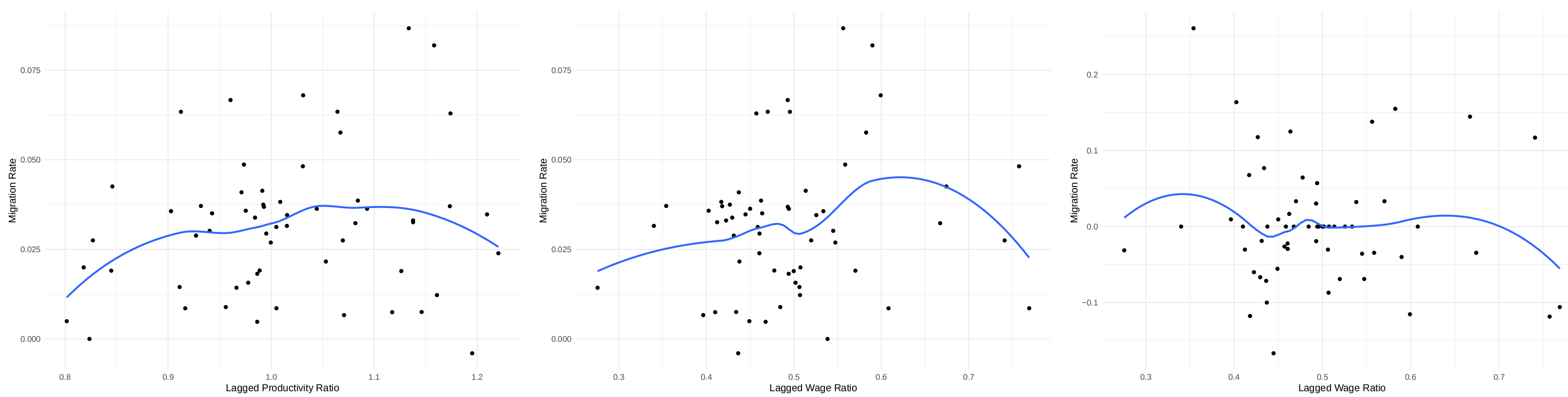


Table 1: Summary Statistics and Variable Description

| Variable | Description | United States Mean | United States Std. Dev. | Japan Mean | Japan Std. Dev. |
|---|---|---|---|---|---|
| $L_{nf}$ | Number of employees in non-farm sector (thousands) | 82,428 | 25,992 | 52,131 | 10,316 |
| $L_f$ | Number of employees in farm sector (thousands) | 3,365 | 1,604 | 5,422 | 3,323 |
| $L_{hired}$ | Number of hired employees in farm sector (thousands) | 1,148 | 381 | 416 | 160 |
| $y_{nf}$ | Value added by non-farm sector (million \$, billion ¥) | 5,591,194 | 5,137,551 | 286,750 | 175,950 |
| $y_f$ | Value added by farm sector (million \$, billion ¥) | 82,331 | 53,068 | 6,846 | 2,665 |
| $APL_{nf}$ | Labor productivity in non-farm sector ($y_{nf}/L_{nf}$) | 55.55 | 41.37 | 49.83 | 27.02 |
| $APL_f$ | Labor productivity in farm sector ($y_f/L_f$) | 37.11 | 35.03 | 18.08 | 9.61 |
| $w_{nf}$ | Wage rate in non-farm sector (\$ per hour, ¥ per hour) | 10.47 | 6.80 | 1,324.27 | 755.57 |
| $w_f$ | Wage rate in farm sector (\$ per hour, ¥ per hour) | 5.41 | 3.89 | 611.16 | 494.22 |
| $un$ | Unemployment rate | 5.96 | 1.55 | 2.78 | 1.29 |
| $nlv$ | Nominal farmland value (\$ per acre, ¥ per hectare) | 927.56 | 858.84 | 787,102.45 | 359,448.99 |
| $PPI$ | Producer price index for farm goods (base = 1982, base = 2015) | 96.13 | 44.23 | 81.67 | 28.80 |
| $rlv$ | Real farmland value (nlv/ppi) | 7.87 | 4.50 | 9,173.34 | 1,876.99 |
| $M_{all}$ | Out farm migration rate | 2.41% | 3.39% | 3.16% | 1.96% |
| $M_{hired}$ | Out farm migration rate | 1.37% | 4.31% | 0.165% | 7.80% |
| r | $\log(APL_{nf}/APL_f)$ | 0.618 | 0.318 | 1.011 | 0.106 |
| w | $\log(w_{nf}/w_f)$ | 0.717 | 0.092 | 0.495 | 0.094 |
| u | $\log(ur)$ | 1.753 | 0.253 | 0.911 | 0.486 |
| g | $\log(L_{nf}/L_f)$ | 3.246 | 0.781 | 2.400 | 0.790 |
| lv | $\log(rlv)$ | 1.897 | 0.594 | 9.102 | 0.216 |

Table 2: Migration of All Farm Workers: Regressed on Productivity

| | Linear Regression | | Threshold Regression | | Kink Regression | |
|---|---|---|---|---|---|---|
| Parameter | Estimate | Std. Error | Estimate | Std. Error | Estimate | Std. Error |
| | | | United States | | | |
| $\theta_0$ | 0.007 | 0.080 | - | - | 0.0327 | 0.212 |
| $\boldsymbol{\theta_1}$ | 0.047 | 0.031 | - | - | - | - |
| $\theta_2$ | -0.006 | 0.018 | -0.001 | 0.018 | -0.0063 | 0.022 |
| $\theta_3$ | -0.024 | 0.027 | -0.024 | 0.030 | -0.0259 | 0.048 |
| $\theta_4$ | 0.041 | 0.028 | 0.040 | 0.028 | 0.0419 | 0.045 |
| $\gamma_0$ | - | - | -0.112 | 0.080 | - | - |
| $\boldsymbol{\gamma_1}$ | - | - | 0.088 | 0.052 | 0.0547 | 0.037 |
| $\gamma_2$ | - | - | -0.004 | 0.128 | - | - |
| $\boldsymbol{\gamma_3}$ | - | - | 0.082 | 0.033 | 0.0434 | 0.088 |
| $\tau$ | - | - | 0.477 | - | 0.4431 | - |
| | | | Japan | | | |
| $\theta_0$ | -0.113 | 0.128 | - | - | 0.0138 | 0.1098 |
| $\boldsymbol{\theta_1}$ | 0.084 | 0.022 | - | - | - | - |
| $\theta_2$ | -0.040 | 0.010 | -0.034 | 0.007 | -0.0380 | 0.0111 |
| $\theta_3$ | 0.008 | 0.006 | 0.007 | 0.005 | 0.0092 | 0.0061 |
| $\theta_4$ | 0.008 | 0.146 | -0.004 | 0.142 | 0.0027 | 0.0124 |
| $\gamma_0$ | - | - | -0.112 | 0.118 | - | - |
| $\boldsymbol{\gamma_1}$ | - | - | 0.206 | 0.079 | 0.1985 | 0.3458 |
| $\gamma_2$ | - | - | -0.004 | 0.128 | - | - |
| $\boldsymbol{\gamma_3}$ | - | - | 0.082 | 0.033 | 0.0655 | 0.0341 |
| $\tau$ | | | 0.961 | | 0.9031 | |

*$\theta$ are fixed coefficients. $\gamma$ are regime-specific coefficients. Bold parameters ($\boldsymbol{\theta_1}, \boldsymbol{\gamma_1}, \boldsymbol{\gamma_3}$) signify slope coefficients on returns to labor.

Table 3: Migration of All Farm Workers: Regressed on Wages

| | Linear Regression | | Threshold Regression | | Kink Regression | |
|---|---|---|---|---|---|---|
| Parameter | Estimate | Std. Error | Estimate | Std. Error | Estimate | Std. Error |
| | | | United States | | | |
| $\theta_0$ | -0.013 | 0.103 | - | - | 0.095 | 0.054 |
| $\boldsymbol{\theta_1}$ | 0.137 | 0.105 | - | - | - | - |
| $\theta_2$ | -0.026 | 0.019 | -0.025 | 0.018 | -0.028 | 0.020 |
| $\theta_3$ | -0.040 | 0.022 | -0.043 | 0.025 | -0.054 | 0.028 |
| $\theta_4$ | 0.060 | 0.027 | 0.061 | 0.033 | 0.078 | 0.036 |
| $\gamma_0$ | - | - | -0.308 | 0.194 | | |
| $\boldsymbol{\gamma_1}$ | - | - | 0.642 | 0.307 | 0.646 | 0.341 |
| $\gamma_2$ | - | - | -0.029 | 0.125 | | |
| $\boldsymbol{\gamma_3}$ | - | - | 0.162 | 0.129 | 0.080 | 0.177 |
| $\tau$ | - | - | 0.638 | - | 0.623 | |
| | | | Japan | | | |
| $\theta_0$ | -0.059 | 0.032 | - | - | -0.06095 | 0.1282 |
| $\boldsymbol{\theta_1}$ | 0.003 | 0.032 | - | - | - | - |
| $\theta_2$ | -0.026 | 0.010 | -0.012 | 0.013 | -0.24 | 0.013 |
| $\theta_3$ | 0.003 | 0.007 | -0.005 | 0.008 | 0.0000 | 0.008 |
| $\theta_4$ | 0.012 | 0.017 | 0.008 | 0.0166 | 0.0128 | 0.015 |
| $\gamma_0$ | - | - | -0.017 | 0.141 | - | - |
| $\boldsymbol{\gamma_1}$ | - | - | -0.008 | 0.049 | 0.0728 | 0.059 |
| $\gamma_2$ | - | - | 0.092 | 0.162 | - | - |
| $\boldsymbol{\gamma_3}$ | - | - | -0.167 | 0.091 | -0.0246 | 0.047 |
| $\tau$ | | | 0.547 | | 0.45 | |

*$\theta$ are fixed coefficients. $\gamma$ are regime-specific coefficients. Bold parameters ($\boldsymbol{\theta_1}, \boldsymbol{\gamma_1}, \boldsymbol{\gamma_3}$) signify slope coefficients on returns to labor.

Table 4: Migration of Hired Farm Workers: Regressed on Wages

| | Linear Regression | | Threshold Regression | | Kink Regression | |
|---|---|---|---|---|---|---|
| Parameter | Estimate | Std. Error | Estimate | Std. Error | Estimate | Std. Error |
| | | | United States | | | |
| $\theta_0$ | -0.206 | 0.135 | - | - | 0.00088 | 0.062 |
| $\boldsymbol{\theta_1}$ | 0.274 | 0.136 | - | - | - | - |
| $\theta_2$ | -0.020 | 0.024 | -0.22 | 0.023 | -0.02177 | 0.023 |
| $\theta_3$ | 0.018 | 0.016 | 0.009 | 0.016 | 0.01596 | 0.019 |
| $\gamma_0$ | - | - | -0.049 | 1.143 | - | - |
| $\boldsymbol{\gamma_1}$ | - | - | 0.086 | 0.150 | 0.14912 | 0.165 |
| $\gamma_2$ | - | - | 0.272 | 0.660 | - | - |
| $\boldsymbol{\gamma_3}$ | - | - | -0.248 | 0.773 | 0.66557 | 0.347 |
| $\tau$ | | | 0.813 | | 0.77516 | |
| | | | Japan | | | |
| $\theta_0$ | 0.313 | 0.086 | - | - | 0.1723 | 0.059 |
| $\boldsymbol{\theta_1}$ | -0.346 | 0.117 | - | - | - | - |
| $\theta_2$ | -0.023 | 0.037 | 0.018 | 0.047 | 0.0027 | 0.046 |
| $\theta_3$ | -0.050 | 0.025 | -0.074 | 0.032 | -0.0673 | 0.033 |
| $\gamma_0$ | - | - | 0.326 | 0.099 | - | - |
| $\boldsymbol{\gamma_1}$ | - | - | -0.330 | 0.185 | -0.1833 | 0.265 |
| $\gamma_2$ | - | - | 0.644 | 0.248 | - | - |
| $\boldsymbol{\gamma_3}$ | - | - | -0.821 | 0.354 | -0.5575 | 0.295 |
| $\tau$ | | | 0.547 | | 0.4925 | |

*$\theta$ are fixed coefficients. $\gamma$ are regime-specific coefficients. Bold parameters ($\boldsymbol{\theta_1}, \boldsymbol{\gamma_1}, \boldsymbol{\gamma_3}$) signify slope coefficients on returns to labor.

Table 5: Estimated Elasticities at Sample Mean

| Specification | Linear Regression | Threshold Regression | Kink Regression |
|---|---|---|---|
| | United States | | |
| $M_{all} \sim r$ | 1.95 | 5.30/2.92 | 4.45/2.68 |
| $M_{all} \sim w$ | 5.68 | 36.93/6.01 | 20.91/18.83 |
| $M_{hired} \sim w$ | 11.33 | 19.45/-5.09 | 38.16/19.07 |
| | Japan | | |
| $M_{all} \sim r$ | 2.66 | 7.44/2.49 | 0.51/1.23 |
| $M_{all} \sim w$ | 0.09 | -0.28/-3.90 | -2.43/-1.92 |
| $M_{hired} \sim w$ | -10.97 | -551.10/-152.09 | |

Table 6: Bootstrap Linearity Tests

| Specification | Statistic | P-Value |
|---|---|---|
| United States | | |
| $M_{all} \sim r$ | 8.48 | 0.41 |
| $M_{all} \sim w$ | 8.48 | 0.41 |
| $M_{hired} \sim w$ | 6.19 | 0.71 |
| Japan | | |
| $M_{all} \sim r$ | 6.51 | 0.85 |
| $M_{all} \sim w$ | 6.51 | 0.86 |
| $M_{hired} \sim w$ | 7.23 | 0.46 |