INTERNAL MIGRATION FROM TURKISH BIRTH HISTORIES

ABSTRACT

Considering residence at the time of each birth for Turkish women increases lifetime moves by 43 percent over a baseline comparing only residence in childhood and at survey, both by capturing multiple moves for some women and by revealing interim moves for women who were interviewed in the same provinces where they had lived in childhood. Reduced-form and full-form models exhibit very similar effects for important determinants of the risk of migration, including educational attainment of women and their husbands as well as salience of kin influence in couples' marriages (as measured by consanguineous marriage and payment of bride-price).

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IMPORTANCE OF MIGRATION HISTORIES

Explaining shifts in human behavior in terms of immediate social context would be much simpler if people always stayed in one place, but as Buckminster Fuller famously observed, people do not have roots. Trees have roots. People have feet, and they use them to move around. Despite the threats that mobility poses for explaining behavior in terms of context, much social research fails to take migration into account. Even when we recognize the issue, the most common difficulty in accounting for migration effects stems from lack of data (Tugault 1973, Greenwood 1985). Panel studies can track people's lives over time and record migration histories in detail, but such studies involve enormous expense and often face serious problems of attrition bias, particularly with respect to migration itself. Population registers provide even better data on migration (Goldstein 1964) but are even less common, more expensive to maintain, and have empirical limitations such as double-counting people in more than one register. A large share of social research must work with a single snapshot of survey or census data for a population. If we are lucky, respondents have been asked where they were born or where they lived as children. The researcher also normally knows where respondents live at the time the data were collected. Discrepancies between these two reports of residence provide the most common measure of lifetime migration (Hamilton 1961, Wadycki 1972, Schultz 1982).

Even this reduced-form estimate of lifetime migration has important explanatory power, but in the absence of more information about when the person moved, duration effects remain a mystery. For example, we may want to link local labor market conditions to the timing of a woman's first birth at some point in her past, but unless we know when she moved from her rural birthplace to her nation's capital city, we don't know which labor market to examine at the time of her first birth. Further, as often pointed out by researchers working with incomplete migration histories, we don't know anything about intervening moves. A person born in one place and living in another may have arrived at that destination by a long and complicated sequence of moves to many intermediate destinations, or may have moved back and forth between origin and destination many times. Even worse, some of the people who appear "stationary" in the reduced-form measure may have moved around between dates of birth and survey, only to return to their original locations in time to be interviewed. How much difference would more migration history make, in accurately measuring migration in order to examine its determinants and effects?

BIRTHS AS MOBILITY MARKERS

Many strategies have been pursued for collecting more complete migration histories (Morrison 1971, Tugault 1973, Greenwood 1985, Tunali 1997). In addition to asking place of birth, for example, censuses and surveys often ask where a person lived five years earlier, or use some other fixed interval. This analysis examines an alternative approach used in the Turkish Demographic and Health Survey of 1993, in which women were asked the standard question about where they had lived as children (before age 12), and in which their residence at the time of the survey also was known. In addition to this comparison needed for the reduced-form measure of lifetime migration, women also reported the province in which they lived at the time of each of their births. Together, these residence variables from the birth histories allow us to test whether at least some residential moves are missed by questions comparing only place of birth or childhood residence to place of current residence, and whether the determinants of migration as measured by this full-form measure look the same as determinants of the reduced-form measure. While internal migration has played a central role in social and economic transformations at many times and in many places (Todaro 1976, Anderson 1980, McQuillan 1980, Hochstadt 1981, Kelley & Williamson 1984), the Turkish survey provides a particularly important context for coming to terms with migration, because massive population movements coincided with dramatic societal change in Turkey in the second half of the twentieth century (Danielson & Keleş 1985, Gökdere 1994, Ilcan 1994).

Although measuring migration from fertility histories has the inherent problem of more and shorter intervals to capture movements for people with more children (Macisco & Myers 1975, Sabagh & Yim 1980), some things about the approach also recommend it. The birth of each child changes the constellation of opportunities and restrictions faced by a family in powerful ways, so each birth interval could be qualitatively different in terms of risk of migration. The Turkish survey only records province of residence for each birth, not details of urban versus rural location, so unfortunately we cannot explore the important urban-rural dimension of population movements (Tunali 1997, Erman 1998) with these variables. However, movements from one province to another also have been highly salient features of the economic, political and demographic evolution of Turkish society, and are worth attention in their own right (Doh 1984). The timing of moves still will not be known precisely, but locating moves during different birth intervals allows us to consider piecewise-constant or other interval-specific forms for the hazard of migration, a clear gain over the reduced-form measure. Distinguishing between open and closed birth intervals for each birth order will give some idea of the seriousness of the problem of more measurement intervals for women with more births.

All dates in the 1993 TDHS are coded in consistent century-months so all birth intervals, including the first interval from age 10 to first birth and the last interval from final recorded birth to interview date, are calculated first in months and then transformed to years with decimal remainders. Although childhood residence is based on a question about residence before age 12, the first recorded birth comes at age 10. For this reason, the age/duration measure begins at age 10 and assumes that moves between ages 10 and 12 would not distort these results.

Of the 6,519 survey respondents, 1,901 (29 percent of the sample) reported different provinces of residence in childhood and at the survey date. The remaining women were interviewed in the same province where they had lived as children. Women lived an average of 22.92 years between their tenth birthday and the interview date, so a summary or reduced-form measure of the average migration hazard would be:

$$\mu_L = \frac{M_L}{N_L \cdot d_L} = \frac{1901}{6519 \cdot 22.92} = 0.0127, \qquad (1)$$

where μ is the average hazard of migration per person-year of exposure, M counts known moves, N gives number of women at risk of moving, d represents the mean duration lived in years, and the subscript $_L$ indicates the reduced-form lifetime measure of these quantities. The value shown measures expected inter-province moves per person-year of exposure.

No cases have missing values for province of residence at time of interview. Province of residence in childhood takes the same range of values as province at interview, but with 75 missing values. Unknown childhood province for 68 women with first births was recoded to province at first birth, producing conservative estimates of migration since none of the imputed values are allowed to produce migration events. Unknown childhood province for seven women with no first births was recoded to province at interview. Again, no imputed values are allowed to generate migration events. Out of 6519 cases, six women reported at least one birth without giving a province where they lived at the time of that birth. All of these women had reported previous births in the province where they were interviewed, so births with unknown province always were higher-order, more recent births. Since earlier births took place in the province of interview, the most logical solution assumed that the most recent births also took place in that province. As always, missing data never generate migration events.

For respondents with no recorded births, the first birth interval remains open and extends from age 10 to the interview date. For respondents with one birth, the closed first birth interval extends from age 10 to the first birth, and the open second birth interval extends from that birth to the interview date. Successive births close higher birth intervals, always leaving an open birth interval that extends to the interview date. All later birth intervals following any open birth interval are coded with zero length, so that for each respondent, the sum of the lengths of all 16 possible birth intervals equals the time between the respondent's tenth birthday and the interview date. This approach insures that person-years are counted consistently across all measures of exposure.

Starting in 1989 Turkey subdivided some existing provinces, creating nine new provinces before the 1993 survey. Some women thus reported an original province as their childhood residence, but without any actual migration they could report living in a new province by one of their births or by the time of the survey. Such changes in province of residence are never counted as "migrations" when they occur after 1989. Appendix Table A1 explores such identified boundary changes, none of which are included in the full-form migration counts below.

With two measures of lifetime migration (the reduced-form measure based only on residence in childhood and at the survey date, and the full-form measure based on additional information from the birth histories), we are in a position not only to compare the volume of migration estimated by each, but also to compare how other social factors predict each of the possible measures of mobility. Following sections explore the most salient of such explanations, and evaluate the results.

CORRELATES OF LIFETIME MIGRATION

Naturally, it would be absurd to try to use number of children ever born or similar measures of reproductive history to "predict" migration outcomes in these data, since the migration count itself depended in some sense on the number of births to each woman. However, other predictors of the probability of migration identified in previous research remain available. This analysis distinguishes first of all between variables commonly used to measure economic development, and others commonly used to reflect cultural change. Economic variables ideally might include both education and labor force participation, but no life-history information on the timing or duration of jobs for respondents was collected in the 1993 TDHS. Therefore, education serves as an indicator of modernization (Shryock & Nam 1965, Lee 1970). Measures of cultural differences all concern some aspect of the salience of traditional corporate kinship groups in the lives and decisions of survey respondents (Barth 1954, Tilly & Brown 1968, Kreager 1986, Bittles 1994).

Educational Attainment

Turkish women interviewed in the 1993 survey showed clear educational hypergamy. Previous studies of these data led to post-survey coding for education that distinguishes six levels: no schooling at all, incomplete primary, complete primary, incomplete secondary, complete secondary, and any higher education beyond secondary school. Comparing the 1993 respondents with their partners, almost half of all couples (44 percent) have equivalent levels of education when coded into these six categories (see Table 1).

Table 1 Here

However, for couples that do not match, the gap in education usually favors women's partners. Less than eight percent of women were more educated than their husbands, while 19 percent of women lagged their partners by one level of education, 24 percent lagged by two levels, and six percent by three or more levels. Three cases out of 6519 were missing educational level for husbands; since a plurality of women reported the same level of schooling for themselves and their husbands, these three husbands were assigned the educational attainment of their wives. The most common combinations of education are shown in Table 2.

Table 2 Here

A number of analyses by Turkish scholars have found that further collapsing these categories yields more manageable results without significant loss of predictive detail. The collapsed categories are often designated "Low" for no education or incomplete primary, "Medium" for completed primary and incomplete secondary, and "Higher" for completed secondary and any higher levels of educational attainment. Exploration of several alternative recodes confirms that this standard recode also is most efficient for both respondents and their husbands in this study of lifetime migration. Table 3 shows the joint distribution of these two recoded education variables. Since this recode reduces the number of levels, hypergamy appears for only about a third of the sample based on this definition, while 60 percent of couples appear in matching categories.

Table 3 Here

Initial models for predicting both the reduced-form and full-form measures of lifetime migration begin with this nine-category matrix for the joint distribution of respondent's and husband's educational attainment. In the case of the reduced-form migration measure, as discussed below, couples in which either the respondent or her husband or both had completed secondary or higher education (cells shown in bold-face in Table 3) had significantly *higher* shares of women who had migrated than did couples in which neither partner had completed secondary schooling. For the full-form model, couples in which both the respondent and her husband had failed to complete primary school (cell shown in italic in Table 3) also had significantly *lower* shares of women migrating than when either partner had completed primary education. This effect did not appear for models predicting the reduced-form measure of migration.

<u>Salience of Family and Tradition</u>

Three aspects of cultural context that could be relevant for lifetime migration decisions were measured in the 1993 Turkish Demographic and Health Survey. The first of these is most widely-known and studied (Litton Fox 1975, Magnarella & Turkdogan 1977) and concerns the distinction between marriages arranged by the partners themselves and marriages in which the larger kin groups play an active role in mate selection. The 1993 survey asked the following (translated) question:

"How was your marriage with your (last) husband arranged? 1) We arranged ourselves; 2) Arranged by the families; 3) Escaped/abducted; 4) Other."

Reflecting the continuing significance of corporate kinship groups in everyday Turkish life in the 1990s, 68 percent of all women responded that their families (and/or the families of their husbands) took an active role in arranging their marriages. The remainder of the sample indicated that the woman and her husband made the decision on their own, or that other circumstances were involved that did not involve family decisions or control.

Arranged marriage correlated strongly with lower levels of education. Net of this education effect, women living in the West and North regions reported significantly less arranged marriage, compared with those in the Central, South, and particularly the East region where arranged marriage was most common.

Another question in the 1993 survey asked:

"Are (were) you blood relatives with your (last) husband? 1) Yes; 2) No.
(IF YES): 1) Father's brother's son; 2) Father's sister's son; 3) Mother's brother's son;
4) Mother's brother's son; 5) Other."

Consanguine marriage, particularly the union of cousins or second cousins, has been documented in many studies of Middle Eastern cultures for centuries as a stable and effective strategy by which kin groups attempt to prevent the dispersal of family wealth through inheritance (Barth 1954, Tunçbilek & Ulusoy 1989, Bittles 1994). Consanguine marriages reflect the continuing power and interests of kinship groups with respect to the marriage decisions of their members. Nearly one-fourth of the 1993 TDHS sample (24 percent) indicated that they were related to their husbands in some fashion. Paternal parallel cousins (father's brother's son) accounted for 4.9 percent, while paternal cross cousins (father's sister's son) accounted for another 3.6 percent of unions. Mother's sister's sons were married to 3.4 percent of respondents, along with mother's brother's sons for 3.2 percent of the women.

Consanguine marriages, like family-arranged unions, more often involved women with lower education and living in the Central, South and particularly the East regions of Turkey.

The final measure of traditional kin influence in the marriage decision concerns the payment of bride-price (Kressel 1977, Remez, 1998)--traditionally, money paid by the groom and/or his family to the family of the bride, in compensation for the loss of her services to her family once she marries and goes to live with her husband. The 1993 survey asked:

"Did your (last) husband or his family pay bridesmoney to your family? 1) Yes; 2) No."

Nearly one-fourth (23 percent) of the 1993 TDHS respondents reported that their husbands or their husbands' families paid bride-price when they got married. This traditional practice also correlated strongly with lower educational levels for women. Even after controlling for education, bride-price was reported most commonly by women living in the East region and least by women in the West region, as for other measures of traditional kinship influence. However, the regional distribution of bride-price differed from that of arranged and consanguine marriages. While the latter two practices were common for women in the South as well as the Central and East regions, bride-price was significantly less likely to be reported in the South region. In fact, women in the South region (net of educational and age effects) were less likely to report payment of bride-price than women in the North.

The reasons for varying geographic distributions of different measures of traditional kinship influence require further exploration in the future, but they give a hint of an important fact with direct bearing on the current study of lifetime migration. As might be suspected from these geographic differences, the overall correlation of the three different measures of kin involvement was low, as reported in Table 4. Clearly, each of the three factors measures something about kinship and culture that is distinct from the others, so that it is appropriate to include all three as simultaneous predictors in models.

Table 4 Here

PATTERNS IN REDUCED-FORM LIFETIME MIGRATION

Both educational attainment and salience of traditional kinship practices related strongly to the chance that a woman moved from one province to another between childhood and the date of the 1993 Turkish survey. Table 5 shows the coefficients from a logistic regression of lifetime migration on all these factors, with ages of women added as a covariate to control for duration of exposure to the risk of moving. The age variable in this model is not significant. All variables except age are zero-one binary variables.

Table 5 Here

Since the logit model presented in Table 5 is additive only in the logarithms of odds ratios, changes in probabilities of moving derived from the model apply only to the reference point specified by the model (women at the mean age for the sample, living in the Central region in couples without completed secondary education, with none of the measured forms of traditional kin involvement in their marriages). From these women's perspective, the coefficients shown in Table 5 mean that unit changes in each variable would produce changes in the probability of moving as shown in the last column. Changes in this probability starting from other reference points would be different, but could be derived from the model coefficients shown.

All considered predictors except age and kin involvement in arranging marriages show statistically significant effects on the log odds of lifetime moves measured by the reduced-form indicator. The largest single effect appears for residence in the West region, where women are more than twice as likely to have moved as women living in the Central region (the reference category). All other regions also differ significantly from the Central region, with migration more likely for women in the South, and less likely for women in the North and East. These regional patterns make obvious sense, because most migration in recent Turkish history has been from the North, Central and particularly the East regions, to the South and particularly to the West (Treadway 1972, Doh 1984, Ulusoy 1993, Gökdere 1994). It is no surprise to find the greatest share of lifetime migrants currently living in the West region.

As noted above, the most efficient distinction for predicting reduced-form migration from education of respondents and their husbands reduces to a single dichotomy. Women in marriages where either they or their husbands (or both) had completed secondary school were nearly twice as likely to move between childhood and the survey date (31 percent) as were women in couples where no one had completed secondary education (less than 18 percent). Greater movement by more educated couples may have reflected greater opportunities, perhaps due to wider and more effective communication and information about life-chances in other provinces.

While arranged marriage was not important for predicting lifetime migration using the reduced-form indicator, the other two measures of kin involvement did reveal statistically significant effects. An estimated 22 percent of women with consanguine marriages and 27 percent of women whose husbands paid bride-price would be expected to move between childhood and survey, compared to less than 18 percent of women with neither form of traditional kin involvement. Since arranged marriage included more than three-fourths of all respondents while less than one-fourth had experienced the latter two kin-related features, it may be that bride-price and consanguine marriage become more likely when the husband already resides or is likely to move far away from the woman's home and family. The two less frequent kin-based traditions may represent a sort of "insurance" for the woman and her family when inter-province migration seems likely. If such an interpretation turned out to be valid, causal direction in this case could run from the choice of marriage partner to the forms of marriage adopted, rather than the other way around; care always must be taken in interpreting observed statistical associations.

PATTERNS IN FULL-FORM LIFETIME MIGRATION

The preceding section presents patterns in lifetime migration as measured by the reduced-form measure, using only childhood residence compared to residence at the survey date. If each respondent could only make one move, considering residence at each successive birth for women would contribute no new information except to pinpoint the timing of a move more accurately (still a major gain in knowledge in itself). The province of residence would be different at the start and end of one interval containing the move. All other intervals would have identical values for province of residence at the beginning and end of the period. This hypothetical pattern of results forms a "null hypothesis" in the present study—that the reduced-form estimate of lifetime migration is complete and accurate. In this case, the average hazard of migration shown in equation 1 above should equal the weighted average of all interval-specific hazards, where the weights are the fraction of the entire sample's total person-years found in each considered interval. The subscript *i* refers to birth intervals and the subscript *j* refers to open versus closed intervals:

$$H_o : \mu_L = \overline{\mu}_{ij} = \sum_{i,j} \left(\mu_{ij} \cdot \frac{N_{ij} \cdot d_{ij}}{N_L \cdot d_L} \right). \qquad (2)$$

On the other hand, if values for province of residence change more than once as we move from residence in childhood, through residence at each birth, and ending with residence at survey, information from the birth histories will capture moves that were invisible to the reduced-form measure comparing only residence in childhood and at survey. In this case, the reduced-form average hazard of lifetime migration will be less than the weighted average hazard from birth interval data shown in equation 2. The ratio of these two average hazards will be the same as the ratio of actual movements counted by the two alternate methods.

As it turns out, nearly 86 percent of women who lived in a different province at interview than in childhood (1,566 out of 1,901) do show a change of residence in only one of all possible considered birth intervals. However, 272 women reveal multiple moves (see columns of the second row in Table 6 for the number of women reporting each number of moves). At the same time, 63 of the women whose residence at survey differed from where they lived in childhood are <u>not</u> counted as migrants by the full-form measure of lifetime migration. These women were interviewed in one of the nine new provinces created out of older provinces after 1989, and the timing of the birth interval in which the shift occurs allows us to guess that the "move" was really just a boundary change (see Appendix A) so these are not counted as lifetime migrations. In total, the migratory 1,901 respondents provide evidence of at least 2,234 moves, one-sixth more than the reduced-form count of respondents with any lifetime movement.

Table 6 Here

In addition, consideration of where respondents lived at the time of each birth brings to light a number of residential changes among women who were interviewed in the same provinces where they had lived as children--see the top row of Table 6 above. One woman interviewed in the same place where she had lived as a child reported only one intervening move. While this seems impossible (after one move she would have to be somewhere else) in fact she was "re-defined" into one of the new provinces and then moved back into the original province to a new address. Other supposedly immobile respondents were most likely (206 cases) to report exactly two moves. By definition, they moved away somewhere long enough to have a child in the new location and then later returned. Another 12 respondents reported three moves, and seven reported four moves. This less-obvious form of undercounting of mobility means that consideration of residences at each birth reveal an additional 477 moves, so that the total number of documented moves reaches 2,711 rather than the reduced-form count of only 1,901. More detailed birth order information thus counts almost 43 percent more moves than the reduced-form method. Similarly, the actual value of the weighted average of interval-specific hazards of migration (see equation 3 below) is 43 percent larger than the reduced-form average hazard shown in equation 1 above.

$$\overline{\mu}_{ij} = \sum_{i,j} \left(\mu_{ij} \cdot \frac{N_{ij} \cdot d_{ij}}{N_L \cdot d_L} \right) = 0.0181. \qquad (3)$$

Examining the actual hazards by birth interval allows us to see that this hazard declines monotonically (in fact, nearly linearly) from one birth interval to the next, as would be expected from Turkish research on migration and the family life cycle (Koç 2001). Estimated rates become very erratic at the highest birth orders, as shown in Figure 1, due to rapidly-shrinking numbers of cases. For example, only 339 women out of the original 6519 closed the eighth birth interval with a birth and went on to the ninth birth interval.

Figure 1 Here

Since the full-form measure of lifetime migration produces more than one move for some women, however, it is no longer suitable to apply a logistic regression approach such as that used for modeling effects on the reduced-form measure. The reported numbers of women with different numbers of moves in Table 6 above strongly suggest that this count variable may be well-suited to Poisson regression analysis. The overall mean number of moves in the population (m_{av}=0.4159) can be modeled as a poisson process, such that the probability of each number of moves is a function of both m_{av} and m_i, where i is each order of move:

$$\Pr(m_i \mid m_{av}) = \frac{e^{-m_{av}} \cdot m_{av}{}^{m_i}}{m_i!}. \qquad (4)$$

In fact, a Poisson distribution of numbers of moves from the observed overall mean number matches the observed distribution very closely, as shown in Figure 2 below. Since the computed variance of the full-form migration measure is larger than its mean, negative binomial regression also was considered as an alternative to Poisson regression, which can under-predict the number of cases with zero values (here no migration). The more complex negative binomial model did not improve results enough to warrant its use.

Figure 2 Here

For Poisson regression, we assume that the number of moves follows a Poisson distribution with a conditional mean that depends on selected characteristics of respondents. The only value to be estimated by the regression is this mean. Once we know the mean, the distribution of women over different numbers of moves is determined as shown in equation 4 above. Poisson regression models treat the natural logarithm of this mean as the result of a linear model that includes a vector of values for the selected characteristics, and their coefficients as estimated by maximum likelihood. The exponential relation between the conditional mean number of moves and the linear model forces all estimated numbers of moves to be positive, as required for a Poisson process (Long, 1997). Within this linear model of the logarithm of the mean, other predictors have coefficients and effects just as in a linear model predicting the logarithm of odds ratios from a binary outcome such as the reduced-form measure. In this way, effects for the other predictors can be compared across models.

As Table 7 shows, most effects observed in the logit regression of the reduced-form measure on predictor variables also appear with similar signs and magnitudes in the poisson regression of the full-form measure on those same predictors.

Table 7 Here

Regional differences resemble those in the reduced-form model, with the West and South once again appearing as net destinations for migrants, and the North, Central and East identified as sending regions. The full-form model also confirms that more education meant more moving from one province to another, and again this effect is observed whether it is the woman herself, her husband, or both who completed secondary education (or perhaps even more schooling). In addition to this effect, also observed for the reduced-form model, Table 7 includes a new significant effect for couples in which neither husband nor wife completed primary schooling. Women living in such couples were significantly less likely to have moved from their childhood provinces of residence. This difference was not significant for the reduced-form measure of lifetime migration. Finally, the effects of various forms of kin involvement in marriages again suggest that the widespread reliance on kin-arranged marriages is not an important predictor of migration. However, special kinds of kin involvement in the marriage (such as consanguineous marriage and payment of bride-price by the husband and/or his family) again appear significantly more frequently for women who have moved between provinces. Finally, in the poisson regression of the full-form measure on predictor variables, age itself emerges as positively correlated with lifetime moves as one might have expected from the outset. This result is not surprising at all, since multiple moves almost by definition require more time to perform. The mean number of moves increases with age of woman. For the reference group of women from the model in Table 7 (women living in the Central region with the medium level of education and no kin involvement in their marriages) the expected mean number of moves would be 0.16 for the youngest woman, 0.26 for a woman at the mean age of 32, and 0.36 for the oldest woman in the sample. As always, the actual number of women with each number of moves (including zero) is assumed to follow a poisson distribution in each case, based on these mean values.

COMPARISON OF MIGRATION MEASURES

The comparison of models based on the reduced-form and full-form calculations of lifetime migration may be facilitated by reference to Figure 3 below, presenting a graphic representation of the coefficients from the models in Tables 5 and 7. The basic similarity of the two models appears clearly in this Figure.

Figure 3 Here

However, some differences between these models deserve attention. The explanatory power of consanguine marriage and payment of brideprice drops in the full-form model. No doubt inclusion of the newly-significant “low-education” variable in this model produced this effect, due to the strong correlation of these traditional kin-based practices with lower educational attainment. Regional differences in migration also become less dramatic for the full-form measure incorporating information from birth histories. With the exception of the East region, regional contrasts (with the Central region as the reference category) decline by more than half from those observed for the reduced-form measure. The regions did not differ significantly in shares of migrants with at least one move who had actually moved multiple times, so the difference between the two measures themselves does not explain the reduced regional effect. A model omitting the newly-significant "low-education" variable also replicates these reductions in regional contrasts, so the effect does not stem from introduction of that variable. The reason for reduced regional contrasts in the full-form model will be investigated in more detail in future research.

The 1993 Turkish Demographic and Health Survey offers a unique alternative for measuring lifetime migration and its determinants. In addition to the standard baseline or reduced-form comparison of residence at time of survey to residence earlier in life (such as at birth or in childhood), the Turkish survey captured residence at the time of each birth for women. Taking these observations into account increases the measured volume of migration by 43 percent over the baseline level, both because it captures multiple moves by some women and because it reveals interim moves by women who were interviewed in the same provinces where they had lived in childhood. The number of moves in the full-form model closely follows a Poisson distribution, allowing Poisson regression to identify coefficients for covariates that are analogous to coefficients from a logistic regression for the reduced-form model. The resulting reduced-form and full-form models exhibit very similar effects for important determinants of the risk of migration, including educational attainment of women and their husbands and also including the salience of kin in couples' marriages (as measured by consanguineous marriage and payment of bride-price). The hazard of migration within birth intervals declines linearly with increasing parity. This hazard also shows very similar levels and trends for open compared to closed birth intervals, suggesting that typically longer durations for open birth intervals do not seriously distort estimates of migration risk. While migration histories more commonly ask about mobility in fixed time intervals (most commonly, the five years before the survey date), measurement based on birth histories may actually offer some methodological advantages. First, this approach combines measurement of births and migration in a single retrospective calendar, streamlining interviews. Second, since births dramatically change the context of constraints and opportunities for a family, birth intervals may provide intrinsically homogenous time periods within which to identify piece-wise constant risks of migration. While this approach does create the theoretical risk that we may measure more migration for women with many children simply because we observe them more times, in practice most of the migration events in the Turkish sample were observed in the earliest birth intervals (see Figure 1 above). While uncovering considerably more population movement, the full-form measure based on information from birth histories preserves the same basic picture of determinants of migration seen in the reduced-form lifetime measure, and may offer an efficient and informative way to increase our awareness of migration in some research contexts.

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Table 1

Respondent's Educational Level Minus Husband's Educational Level

Difference	Cases
-5	18
-4	65
-3	306
-2	1,559
-1	1,219
0	2,858
1	341
2	140
3	9
4	1

Source: calculated from 1993 Turkish Demographic and Health Survey

Table 2

Most Common Combinations of Educational Attainment

(hypergamy combinations in italics)

Respondent	Husband	Cases	Percent
primary	primary	1976	30.3%
none	*primary*	*1064*	*16.3%*
primary	*< secondary*	*591*	*9.1%*
none	none	382	5.9%
primary	*secondary*	*365*	*5.6%*
< primary	*primary*	*265*	*4.1%*
higher	higher	176	2.7%
secondary	secondary	152	2.3%
secondary	*higher*	*151*	*2.3%*
<secondary	<secondary	141	2.2%
primary	*higher*	*141*	*2.2%*
<secondary	*primary*	*140*	*2.1%*
all other	all other	972	14.9%
Total	Total	6516	100.0%

Source: calculated from 1993 Turkish Demographic and Health Survey

Table 3
Recoded Education for Respondents and Husbands

Respondent	Husband Low	Medium	Higher	Total
Low	*562*	1535	**104**	2201
Medium	126	2848	**686**	3660
Higher	**2**	**137**	**516**	655
Total	690	4520	1306	6516

Source: calculated from 1993 Turkish Demographic and Health Survey
(*italics* = significantly less migration in full-form model)
(**bold** = significantly more migration in both models)

Table 4
Inter-correlation of Measures of Traditional Kinship Influence

	arranged	consanguine	brideprice
arranged	1		
consanguine	0.0789	1	
brideprice	0.1919	0.0845	1

source: calculated from the 1993 Turkish Demographic and Health Survey

Table 5
Logit Regression Coefficients for
Reduced-Form Model of Lifetime Migration

	Coefficent	Std. Err.	z-score	Pr>|z|	lower 95% CI	upper 95% CI	Pr (Move)
Intercept	-1.5458	0.0849	-18.2	0.000	-1.7123	-1.3794	0.176
Region							
west	1.0279	0.0798	12.9	0.000	0.8715	1.1842	+0.198
south	0.2442	0.0885	2.8	0.006	0.0707	0.4177	+0.038
central	0.0000						
north	-0.3883	0.1047	-3.7	0.000	-0.5935	-0.1832	-0.049
east	-0.2565	0.1055	-2.4	0.015	-0.4633	-0.0497	-0.034
High Educ	0.7499	0.0681	11.0	0.000	0.6165	0.8834	+0.135
Arranged	-0.0589	0.0631	-0.9	0.350	-0.1826	0.0648	-0.008
Consanguine	0.2982	0.0682	4.4	0.000	0.1645	0.4320	+0.047
Brideprice	0.5398	0.0697	7.7	0.000	0.4032	0.6764	+0.092
Age (years)	0.0022	0.0034	0.6	0.530	-0.0046	0.0089	+0.000

Source: calculated from the 1993 Turkish Demographic and Health Survey
(reference group = women at mean age in Central region, medium education or medium+low for both partners, no kin influence in marriage decision)

Table 6
Lifetime Moves with and without Information from Births

Lifetime Moves	women	none	one	two	three	four	five	moves
none (red. form)	4618	4,392	1	206	12	7	0	477
any (red. form)	1901	63	1,566	175	76	15	6	2234
Total	6519	4,392	1,567	381	88	22	6	2711

Source: original tabulations from 1993 Turkish Demographic & Health Survey

Table 7
Poisson Regression Coefficients for
Full-Form Model of Lifetime Migration

	Coefficient	Std. Err.	z-score	Pr>|z|	lower 95% CI	upper 95% CI	Pr (Move)
Intercept	-1.3199	0.0591	-22.3	0.000	-1.4357	-1.2041	0.267
Region							
west	0.5190	0.0534	9.7	0.000	0.4143	0.6238	0.182
south	0.0951	0.0617	1.5	0.123	-0.0259	0.2161	0.027
central	0.0000						
north	-0.2342	0.0735	-3.2	0.001	-0.3783	-0.0902	-0.056
east	-0.2494	0.0775	-3.2	0.001	-0.4012	-0.0976	-0.059
Education							
both low	-0.2416	0.0812	-3.0	0.003	-0.4007	-0.0825	-0.057
all others	0.0000						
either higher	0.6457	0.0435	14.8	0.000	0.5604	0.7310	0.242
Arranged	-0.0248	0.0426	-0.6	0.560	-0.1082	0.0586	-0.007
Consanguine	0.2545	0.0459	5.5	0.000	0.1644	0.3445	0.077
Brideprice	0.3017	0.0477	6.3	0.000	0.2083	0.3952	0.094
Age (years)	0.0222	0.0024	9.4	0.000	0.0176	0.0269	0.006

Source: calculated from the 1993 Turkish Demographic and Health Survey
(reference group = women at mean age in Central region, medium education or medium+low for both partners, no kin influence in marriage decision)

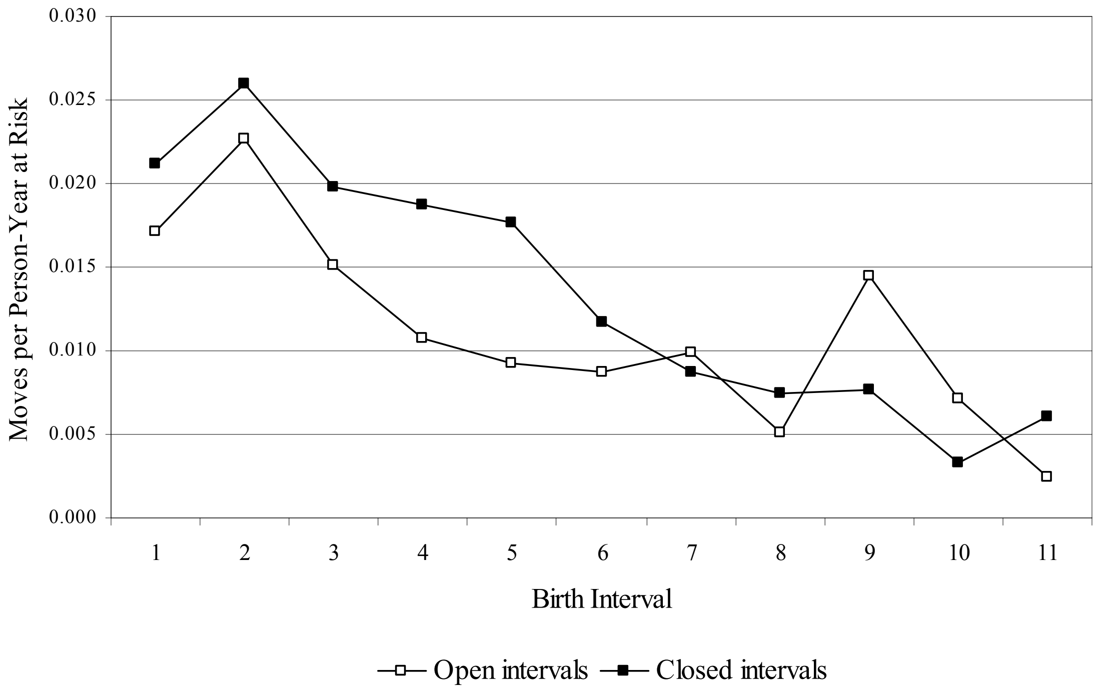

Figure 1 - Full-Form Hazard of Moving by Birth Interval
(Open and Closed Intervals)

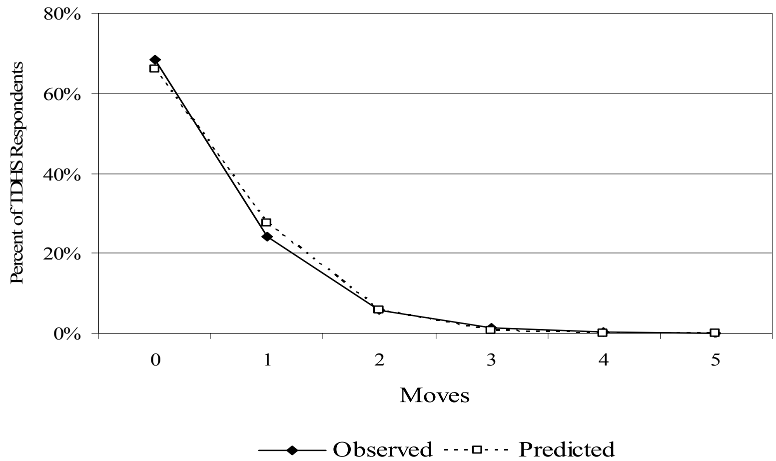

Figure 2 - 1993 TDHS Respondents by Number of Moves
(Observed & Poisson distributions)

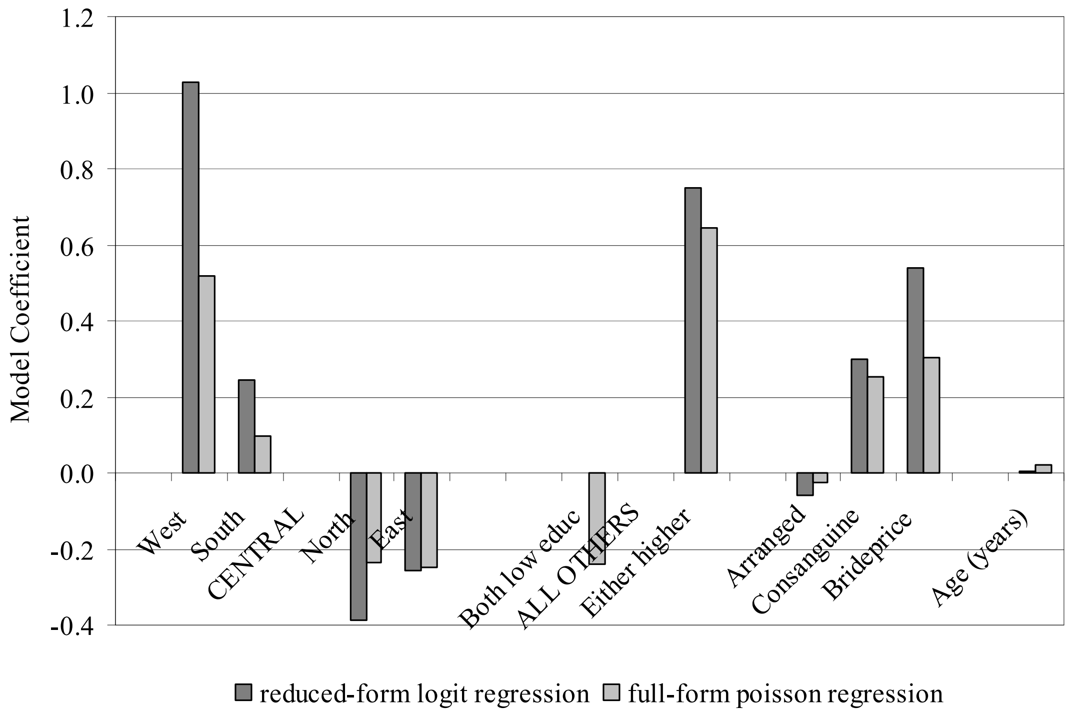

Figure 3 – Coefficients Predicting Migration from Logit and Poisson Regressions
(Reduced-form versus Full-form Measures of Lifetime Mobility)