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Research Article

Migration and marriage: Modeling the joint process

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Migration and marriage: Modeling the joint process

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Abstract

BACKGROUND
Previous research on interrelations between migration and marriage has relied on overly simplistic assumptions about the structure of dependency between the two events. However, there is good reason to posit that each of the transitions has an impact on the likelihood of the other, and that unobserved common factors may affect both migration and marriage, leading to a distorted impression of the causal impact of one on the other.

OBJECTIVE
We will investigate the relationship between migration and marriage in the United States, using data from the National Longitudinal Survey of Youth 1979. We allow for interdependency between the two events and examine whether unobserved common factors affect the estimates of both migration and marriage.

METHODS
We estimate a multi-process model in which migration and marriage are considered simultaneously in regression analysis and there is allowance for correlation between disturbances; the latter feature accounts for possible endogeneity between shared unobserved determinants. The model also includes random effects for persons, exploiting the fact that many people experience both events multiple times throughout their lives.

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RESULTS
Unobserved factors appear to significantly influence both migration and marriage, resulting in upward bias in estimates of the effects of each on the other when these shared common factors are not accounted for. Estimates from the multi-process model indicate that marriage significantly increases the hazard of migration, while migration does not affect the hazard of marriage.

CONCLUSIONS
Omitting interdependency between life course events can lead to a mistaken impression about the direct effects of certain features of each event on the other.

1. Introduction

Migration decisions are made in conjunction with decisions about other life course events (Groot, Mulder, Das, and Manting 2011; Kulu and Milewski 2007; Schachter 2001). For example, when young adults leave the parental home and settle in new places in order to pursue educational and occupational opportunities (Garasky, Haurin, and Haurin 2001), some will find partners there and form families (Kulu 2008). Changes in union status (formation or dissolution) and childbearing can also motivate changes in residence, in part due to space requirements (Flowerdew and Al-Hamad 2004; Kulu and Milewski 2007). These and other plausible life course scenarios suggest an interplay between migration and other life course transitions that is far from unidirectional (Kulu and Milewski 2007). Nevertheless, most empirical research to date has examined unidirectional effects only. This has certainly been the case with research on migration and marriage, most of which has simplified the role of the former, with migration serving as an explanatory variable that either promotes or delays union formation (Feijten and Mulder 2002; Guzzo 2006). However, we suspect that the relationship between migration and marriage is more complicated. It is plausible that migration affects marriage and marriage affects migration; ignoring this more complicated interdependency may yield biased estimates of the causal effects of interest, including the effects of other explanatory variables – for example, schooling history (Mulder and Wagner 1993; Kulu and Milewski 2007; Steele, Kallis, Goldstein, and Joshi 2005). In addition, selectivity could be an issue – for example, migrants could be persons who are prone to marry, or vice versa (Kulu and Milewski 2007; Mulder and Wagner 1993).

Despite these concerns, the comprehensive set of plausible interrelations between migration and marriage has hardly been examined in empirical research to date. And the few studies that have considered this range of interrelations have relied on data from

Europe (Michielin and Mulder 2008). While it may be reasonable to draw implications for the interdependency in the United States from findings in European countries, it is the case that migration and marriage in the U.S. are distinct. For example, rates of internal mobility in the U.S. have generally been higher than in European countries (Molloy, Smith, and Wozniak 2011) and marriage retains an importance in the U.S. (Cherlin 2004) which is not observed in some European societies (especially in northern Europe). Therefore one must question if the nature and magnitude of effects evident in Europe – for example, Michielin and Mulder 2008 for the Netherlands and Mulder and Wagner 1993 for West Germany – characterize the U.S.

In an effort to fill this gap in the literature, the current study investigates migration and marriage in the U.S., using panel data from the National Longitudinal Survey of Youth 1979 (NLSY79). Since individuals are expected to experience these two events several times over the life course, we model the series of events employing a multilevel model which accounts for the individual propensity to migrate or to marry (Steele et al. 2006). Moreover, to allow for an association between migration and marriage that is not accounted for by the measured explanatory variables, we estimate a multi-process model in which equations for both types of events are estimated simultaneously and, further, a residual correlation between the two is allowed. This type of model was introduced into the demographic literature by Lillard and Waite (1993).

This study extends the previous literature in several important respects. First, a better fix on the associations between marriage and mobility contributes to a better understanding of interrelationships between family transitions and residential changes over the life course. Although a growing body of research has investigated the association between various life course transitions, when it comes to migration most existing research has been limited to the effects of migration on reproductive transitions (Clark and Withers 2009; Kulu and Steele 2013; Michielin and Mulder 2008). Consequently, it is not clear from past research how marital transitions affect changes in residence, and in turn how mobility influences marital decisions. In the existing literature, moves have been understood as an adjustment to and a trigger of family events, especially childbirth – for example, subsequent moves in response to family growth (Kulu and Steele 2013). Parallel to this literature, we examine interrelations between migration and marriage; it cannot be assumed that these will take the same form as migration–fertility interrelations. Second, analyzing histories with multiple occurrences of migration and marriage strengthens estimates of the effects of each on the other. Both transitions – geographic moves and entrance to marital union – are affected by unmeasured individual propensities which possibly are associated with measured explanatory variables of interest. For example, some people prefer to live in specific areas (Benson and O'Reilly 2009), while some are attracted by other environments. Moreover, research on marital decisions acknowledges the likely impact of variables which are often unmeasured, such as physical appearance or personality (Lichter, Anderson, and Hayward 1995). Estimation relying on histories with multiple events of both types provides some leverage for controlling for unmeasured propensities to transition (move or marry). There is the further obstacle to estimation caused by decisions to marry and to move which are made jointly; not allowing for this can also result in biased estimates. The current research addresses both of these obstacles to recovering valid estimates of the causal effects of interest. A final contribution of this research is the empirical territory which it encompasses. We analyze longitudinal information on a nationally representative sample of individuals in the United States who were born between 1954 and 1957. These data cover each individual's lifetime from late teens to early 50s during the historical period 1979–2008. That is, we examine migration–marriage interrelations over three decades of individual life course and three decades of historical experiences.

2. Background and research questions

2.1 Migration and marriage

Mobility patterns vary by life course stage in part because changes in family size require residential adjustment (McAuley and Nutty 1982; Michielin and Mulder 2008). More fundamentally, family-building events are well recognized as a major cause of migration, and in particular migration is closely related to marriage (Guzzo 2006; Speare and Goldscheider 1987). For example, Michielin and Mulder (2008) found that the hazard of moving in the Netherlands increases within six months of becoming married. Although this study considered only first marriages, the results confirm a close relationship between migration and marriage. With regard to the association between the two life course transitions in the U.S., marriage increases the likelihood of migration as newlyweds settle in new places or at least one partner moves in with the other (Speare and Goldscheider 1987). This positive relationship is observed in both first marriages and remarriages, and it persists for several years (Speare and Goldscheider 1987). This particular study used a geographically restricted U.S. sample (Rhode Island), and the findings should be confirmed using a national sample from more recent years. On the one hand, family accumulation (partner, children) can lower the chance of migration, due to the increased financial costs of moving as well as the costs of breaking ties (Long 1973; Michielin and Mulder 2008).

Prior studies find that migration influences union formation in two contradictory ways. First, migration can encourage marriage to the extent that the move improves an individual's socioeconomic status (Cadwallader 1992; Massey et al. 1993). Empirical research confirms that migrants from non-metropolitan counties reach higher levels of educational attainment and earnings after moving to a metropolitan county (Mills and Hazarika 2001), thereby improving their marriage market prospects. More directly, migration can improve marriage market prospects if individuals move to a location where there is a larger supply of marriageable mates (Lichter et al. 1992; Lichter et al. 1995; Oppenheimer 1988; South and Lloyd 1992). If a marriage market fails to provide favorable conditions matched to a mate seeker's criteria, one strategy is to compensate by adjusting the criteria and in effect casting a wider net (Lichter et al. 1992; Qian, Lichter, and Mellott 2005). Moving to a new place may be an alternative strategy to casting a wider net in one's current marriage market, just as job seekers move from low to high wage places to improve income (Massey et al. 1993). An empirical finding of increased likelihood of marrying after migration would support the proposition that migration is, among other things, a strategy for expanding marital opportunities.

Second, even so, the short-term impact of migration on marriage may be negative, because time is required to adjust to new environments, including becoming familiar with opportunities to meet potential marriage partners (Jampaklay 2006). Unless the move occurs with a partner, meeting and courting potential partners in new places may be difficult at first. In studies on relationships between migration and fertility, for instance, the fertility levels of migrants are low in the immediate post-migration period but then eventually catch up with those of non-migrants later (Goldstein 1973; Kulu 2005). As with fertility, marital behaviors may be disrupted by the migration process temporarily. Thus it is reasonable to expect that the probability of marriage decreases within the short period of a migration event, later rebounding as the length of residence increases and possibly exceeding the probability of marrying in the place of origin.

2.2 Correlated processes between marriage and migration

Research on migration has disproportionately focused on the first move, even though most people migrate several times throughout the life course (DaVanzo 1983; Molloy et al. 2011). The initial move is, however, less likely to satisfy movers' needs, because they may have had little information about new environments and be disappointed by the discrepancy between their expected and actual gains (McHugh, Hogan, and Happel 1995). As a result, some of the initial movers seek additional destinations, while others return to their original places (McHugh et al. 1995; Wilson et al. 2009). The sequential migration patterns vary by location-specific capital and length of residence in the place (DaVanzo and Morrison 1981) and also by personal traits which are not captured by explanatory variables typically available in major datasets (DaVanzo 1983; Gabriel and Schmitz 1995). Similarly, some individuals experience multiple union transitions (formation and dissolution) over their lifetime (Cherlin 2010), and the likelihood of multiple transitions is itself associated with socioeconomic status and race/ethnicity, as well as personal values and personality factors (Smith 2005; Thornton and Young-DeMarco 2001). Such personal traits, unmeasured in most data that offer good measurement of migration and marriage transitions, affect marital choices and differentiate those who are involved in a succession of marriages from those who are not. In regression analysis with appropriate data, these unmeasured personal traits can be represented by individual-specific random effects (for the propensity to move, the propensity to marry, and the correlation between the two).

In addition, unobserved traits probably play a significant role in other life course transitions (Kulu and Milewski 2007; Mulder and Wagner 1993). For example, using data from West Germany, Mulder and Wagner (1993) investigate the interrelation between migration and marriage employing a model that allows for an interaction between the two events (largely ignored in the prior literature). They find that the estimates from a model which allows for this interaction differ in important respects from estimates that do not allow for it. More specifically, if the model does not allow for an interaction, it appears that women younger than age 25 are more likely than young males to move a short distance. However, once the interaction is allowed, sex differences in short-distance moves disappear and age-specific moving patterns become less pronounced (Mulder and Wagner 1993). In other words, the sex differences in mobility are largely attributed to its interrelation with marriage. Despite this demonstration two decades ago of the importance of allowing for residual co-occurrence of the two events, few empirical studies since then have done so.

Simultaneous relationships between migration and marriage can be estimated using a multi-process model (Lillard and Waite 1993; Steele et al. 2005). In this research, we employ a model with separate equations for migration and union transitions, with each equation having a random effect for unmeasured person-specific propensities. The model also allows for a correlation between these two random effects – that is, association between propensities to move and to marry. Consistent estimation of the two random effects and their correlation is achieved via joint estimation of the two equations. The multiple episodes for each person of risk of union transition and risk of migration transition are the basis for identification of the model (Rabe-Hesketh and Skrondal 2012; Steele et al. 2006).

2.3 Research questions

Informed by this previous literature, we pose the following research questions:

1) Does marriage affect the hazard of migration?
2) Does migration affect the hazard of marriage?
3) Do unmeasured person-specific traits affect the co-occurrence of migration and marriage?

3. Data and methods

3.1 Data and measurements

This study uses public and geocode data from the National Longitudinal Survey of Youth 1979, which contains a nationally representative sample of individuals in the United States. The NLSY79 interviewed 12,686 respondents from 1979, when they were ages 14 to 21, until their late 40s and early 50s in 2010. The analyses include longitudinal information on 9,763 respondents, omitting a military sample of 1,280 individuals having unusual moving patterns and a subsample of 1,643 economically disadvantaged non-black and non-Hispanic individuals who have not been interviewed since 1990. We restrict the migration and union transitions to survey years from 1979 to 2008, when information on both transitions is available. The NLSY provides a large amount of information on family formation, education, and employment in public files, and information about residential mobility in the geocode files. This detailed measurement is the basis for construction of the required life course histories.

For marriage, we draw information from the partner-specific characteristics files which have information on relationship status (spouse, partner, or single) every survey year (Center for Human Resources 2013). To create a complete marital history, we also use the actual starting and ending dates of marriages as provided in the public files. About 79% of the respondents experience first marriage by age 26 and about 25% of them marry again (20% of the entire sample). On average, the first marriage lasts for 12.9 years (see Table 1).

Table 1: Description of migration and marriage histories

	Proportion	Mean age		Duration	
Migration history					
Never moved	.50	-		-	
1	.21	25.48	(.18)	4.40	(.12)
2	.13	28.15	(.20)	3.77	(.11)
3	.07	30.24	(.22)	3.78	(.15)
4	.04	32.27	(.27)	3.89	(.22)
5	.02	34.27	(.40)	3.75	(.27)
More than 5	.03	-		-	
Marriage history					
Never married	.21	-		-	
1	.61	25.61	(.13)	12.86	(.16)
2	.16	33.46	(.19)	8.73	(.16)
3	.02	36.64	(.39)	7.23	(.25)
more than 3	.01				

Notes: 87,931 person-years for 7,827 respondents are included. Marital duration of those who are still married until last interview was calculated using the year of last interview (2008). All statistics were adjusted under survey setting in Stata. Numbers in parentheses are standard errors.

In contrast to marriage, the exact dates of migration are not available in the NLSY79 prior to 2000. We therefore use the county and state FIPS (Federal Information Processing Standard) codes in the geocode file and define a migration event as changes in county of residence between the annual survey interviews. Migration is therefore inferred indirectly, and some short-term moves are missed, namely those that are followed by another move before the next interview (respondent is correctly classified as a migrant, but one destination is missed) and short-term moves that are followed by return to residence at previous interview prior to the next interview (respondent is incorrectly classified as non-migrant). Moreover, the chronology of moves and marriage that occur in the same year is uncertain. This is an unfortunate limitation of the NLSY79 measurement of migration prior to 2000. We assume that responses of marriage to migration occur with some lag, and vice versa, and therefore lag each event one year with respect to the other. Moreover, it is possible that premarital cohabitation (common precursor to marriage) distorts the relationship between migration and marriage. In our data, about 7% of moves follow premarital cohabitation – that is, migration intervenes between cohabitation and marriage, and should be regarded as occurring after the partnership has started (and therefore it would be inappropriate to attribute a causal effect of the moves on marriage).[4] To avoid bias due to this phenomenon, we drop from the analysis moves which follow immediately on premarital cohabitation. In the resulting sample, about 50% of the sample have moved at least once from their county of residence in 1979 by age 26, and the duration in the new place averages about four years (see Table 1).

Since the analyses take an event history approach, data are transformed into person-year files which contain 134,204 person-years. To estimate equations for migration and marriage simultaneously, we restrict the sample to those who were age 16 at the first interview, which is taken as the onset of risk for both first migration and first marriage. The final dataset includes 7,827 individuals who provide 87,931 person-years for analysis.

Table 2 displays explanatory variables included in the analyses. These are individual, household, and county characteristics which have been identified in the existing literature as determinants of each type of transition. The NLSY offers a rich array of variables to represent hypothesized determinants.

Table 2: Description of individual, household, and county characteristics

Variable	Proportion	Variable	Proportion	
Female	.47	Childbirth	.30	
		Total number of children (numbers)	2.24	(.03)
Race				
White	.72	Household characteristics		
Black	.21	Had lived in an intact family until age 18	.64	
Hispanic	.07	Maternal educational attainment (years)	11.62	(.09)
Education		Living in metro areas	.84	
Less than high school	.22	County or residence characteristics		
High school or equivalent	.60	Female population (%)	51.25	(.11)
College or more	.18	Male population (%)	48.65	(.10)
		Population with a college degree or higher (%)	10.96	(.21)
Employment		Unemployment rates (%)	5.83	(.13)
Employed part-time	.44	Poverty rates (%)	1.62	(.04)
Employed full-time	.44	Crime rate (%)	5.86	(.14)

Notes: 87,931 person-years for 7,827 respondents are included. All statistics were adjusted under survey setting in Stata. Numbers are proportions except those indicated. Numbers in parentheses are standard errors.

[4] We thank an anonymous reviewer for drawing our attention to this.

As distinct developmental tasks are required at each stage of life course (McAuley and Nutty 1982, 1985), age effects (duration from age 16 to age at which events of interest occur) are controlled for in the models with square and cubical terms. Previous empirical research on life course events shows mixed findings about gender differences. Females marry earlier than males in virtually all societies and subgroups (Kreider and Ellis 2011), but gender differences in migration vary by study. Males are in general more likely than females to migrate, probably because historically men invest more in human capital (Quinn and Rubb 2011). On the other hand, young female adults leave their parental home earlier than their male counterparts (Buck and Scott 1993; Garasky 2002) and unmarried females move significantly more than their male counterparts, especially in the case of long-distance moves (Mulder and Wagner 1993). This is partly because a higher proportion of men are involved in post-secondary education and females marry earlier (DeJong 2000; Long 1973). To sort this out, we examine the effects of gender on migration and marriage.

Race and ethnicity also influence the timing of marriage and migration (Glick et al. 2006). Whites tend to delay marriage until they have achieved a first set of career goals, but the lifetime risk of marriage is higher for whites than for blacks and Hispanics (Goldstein and Kenney 2001). Studies on racial differences in moving patterns also suggest that mobility rates vary by the characteristics of places, such as racial distribution (Crowder 2000) and racial differences in wage rates (Wolaver and White 2006). In addition, individual socioeconomic statuses represent capability to manage life course transitions (Oppenheimer 1988). For example, those who have completed their schooling and are employed full-time are more likely to marry but less likely to move, whereas those who are in school and unemployed tend to show unstable life course transitions, including frequent moves as part of job searches (DaVanzo 1983). To the extent that socioeconomic contexts affect life course decisions, we also include local economic conditions in the current county of residence as control variables, assuming that individuals in places with worse economic conditions are more likely to move (Massey et al. 1993). Our indicators of local economic conditions are unemployment rates, poverty rates, and the share of population having a college or higher degree. We also expect migration to be influenced by comfort or concern with local public safety (Cadwallader 1992; McAuley and Nutty 1982) and therefore county-level crime rate is included in the equations. Finally, the local demographic structure may influence either migration or marriage, the latter by affecting the availability of potential mates of the opposite sex (South and Lloyd 1992); to account for this, the sex ratio (percentage of opposite-sex population in the county) is a further explanatory variable.

Household characteristics can also affect life course decisions. Greater household socioeconomic resources promote independence and autonomy for household members, resulting in earlier migration and independent living (Avery, Goldscheider, and Speare 1992). Unstable family structure encourages individuals to move frequently and form their own family earlier, although the opposite has also received empirical support such that economic and emotional support from a stable family encourages their adult children to be independent (Avery et al. 1992). We allow for these hypothesized relationships by including explanatory variables of maternal educational attainment and whether respondents had lived in an intact family until age 18. We also include an indicator of type of place of residence (non-metropolitan or metropolitan areas), because previous research reveals significant variation in life course experiences by type of place (Cromartie 1993; Snyder, Brown, and Condo 2004). A yearly time-varying measure of non-metropolitan and metropolitan residence is merged from the 2003 Urban Influence Codes from USDA ERS (United States Department of Agriculture Economic Research Service) due to changes in standards measuring the metropolitan and non-metropolitan statistical areas in the NLSY79.[5]

3.2 Analytical strategies

We estimate equations for the discrete-time hazards of migration and marriage (Allison 1984). The fact that many NLSY respondents experience more than one marital and migration event provides a basis for identification of person-specific random effects. The equations for each process can be specified as:

$$\log[h^{MIG}(t)] = \alpha_0^{MIG} D(t) + \alpha_1^{MIG} F(t) + \alpha_2 X^{MIG} + \alpha_3 Marriage_{(t-1)} + u^{MIG} \quad (1)$$

$$\log[h^{MARR}(t)] = \beta_0^{MARR} D(t) + \beta_1^{MARR} F(t) + \beta_2 X^{MARR} + \beta_3 Migration_{(t-1)} + u^{MARR} \quad (2)$$

Equation (1) is for the hazard of migration at time t (log [h^{MIG}(t)]); D(t) represents the duration pattern of migration following the onset of the risk. Once individuals

[5] As for classification of non-metropolitan and metropolitan areas, the NLSY79 uses the 1973 City Reference File (CRF) during 1979-1982, the 1982 CRF in 1983, the 1983 CRF for 1984-1987, the 1987 CRF for 1988-1992, the 1992 CRF for 1993-1998, and a slightly different calculation process from 2000 to 2006 (Center for Human Resource Research, NLSY79 Codebook Supplement, Appendix 6 2013). Due to the changes, some respondents can appear to move from metropolitan to non-metropolitan areas though they have not changed their residence.

move, they are at risk of a next move. $F(t)$ denotes a time-varying covariate whose values change over time: educational attainment, employment status, living in metropolitan areas, and county characteristics in this study. X denotes time-constant variables such as demographic and household factors at the first interview. The equation also includes selected facets of the marriage history, namely total number of marriages and whether a marriage occurred in year t-1; both are time-varying. As described, a person-specific residual, u^{MIG}, is included to represent the person-specific propensity to move that is not captured by measured explanatory variables (Allison 1984; Rabe-Hesketh and Skrondal 2012).

As with the migration equation (1), equation (2) is for the hazard of marriage and consists of a set of terms, D, capturing the duration pattern, and a large set of time-varying, ($F(t)$), and time-constant, (X), covariates. And, analogous to the migration equation, equation (2) contains selected facets of the migration history, namely total number of moves and whether a move occurred in year t-1. In the analysis of marriage, the onset of risk for first marriage is defined as age 16. Whereas after each move an individual is at risk of another move, in the case of marriage the subsequent episode is the risk of marital dissolution. Only after marital dissolution are individuals at risk of transitioning into another marriage. A person-specific random effect, u^{MARR}, controls for any unobserved heterogeneity for the same individual, affecting marriage and being constant across subsequent marriages.

Turning to the random effects, these are assumed to follow the normal distribution, with a variance specific to each effect to be estimated from the data:

$$u^{event} \sim N(0, \sigma^2)$$

We first assume that the two person-specific propensities (to move and to marry) are independent of each other. This specification is correct if the facets of the two histories included as explanatory variables fully capture the effects of each history on the other and if there are no unobserved explanatory variables that affect both histories. These are very strong assumptions and it seems unlikely they are satisfied. Therefore it is safer to posit a correlation between these two random effects – that is, an association between the propensity to move and to marry that is not captured by the measured right-hand-side variables. This form of endogeneity can be accounted for by estimating the two equations simultaneously and allowing the two disturbances to be correlated (Steele et al. 2005; Upchurch, Lillard, and Panis 2002).

$$u = (u^{mig}, u^{marr}) \sim N(0,0) \begin{pmatrix} \sigma_{mig}^2 & \\ \sigma_{mig,marr} & \sigma_{marr}^2 \end{pmatrix}$$

3.3 Identification

Identification of a two-equation system with random effects and correlated random effects typically requires covariates that are included in one equation but excluded from the other – that is, instrumental variables (Brien, Lillard, and Waite 1999; Lillard and Waite 1993; Steele et al. 2005, 2006). But if respondents experience repeated events of each type – that is, multiple marriages and multiple moves – this provides an alternative to covariate exclusion as a means of model identification (Lillard, Brien, and Waite 1995; Steele et al. 2005; Upchurch et al. 2002). All sources of correlation between migration and marriage are accounted for by person-specific random effects that are constant across replications for the same individual. Then the remaining variation after accounting for the correlation across two processes represents the effects of previous moving (or marriage) on the current episode, which is exogenous from the other process (Upchurch et al. 2002). This means of model identification has been utilized in previous research on interdependent life course histories because it is often difficult to locate variables that satisfy the strict exclusion requirements (Steele et al. 2005).

4. Results

Table 1 describes respondents' migration and marriage experiences over survey years. About half of the respondents never move from their county of residence in 1979, while the other half migrate at least once. On average the first move takes place in their mid-20s. With regard to marriage, 79% of the respondents marry once during the survey period and about 20% experience a second marriage. Few people are involved in third or higher order marriages. The mean age at first marriage is about 26, while the second marriage takes place on average at age 33.

We consider two specifications of the two-equation model presented above: equations (1) and (2). In the first, the migration and marriage equations are estimated independently – that is, as a pair of independent processes (although the explanatory variables for both processes contain aspects of the parallel history). The second specification is a multi-process model that includes correlation between the person-specific random effects. If this correlation emerges as significant, the conclusion is that unobserved propensities to move and to marry are associated, and the coefficients from the single-process model are biased (Steele et al. 2005).

4.1 Correlation between random effects

We begin by considering the random effects. Table 3 presents the random effects estimated from the single-process and multi-process models. In the multi-process model, we find significant correlation between the migration and marriage random effects ($\sigma_{mig\ *marr}$ = .24). This indicates that some components not included in the models of migration and marriage make people more likely to both move and marry. Moreover, the coefficient on the marriage variables in the migration equations (b = .60) and that on the migration variables in the marriage equations (b = .11) are different in the multi-process model (b = .28 and .04 respectively: see top rows in Tables 4 and 5).[6] This suggests upward bias in the coefficients in the single-process model, possibly because of shared unobserved determinants. In other words, the estimated positive effects of marriage on migration, and vice versa, in the single-process specification in part reflect selection of individuals with a tendency toward migration and marriage, rather than causal impact of one history on the other. The discrepancy between single-process and multi-process models in Tables 4 and 5 indicates that not allowing for the correlation of disturbances results in a distorted impression of the causal relationships between migration and marriage.

Table 3: Estimated random effects

	Migration		Marriage	
Migration	.13	(.11, .15)		
Marriage	.24	(.22, .26)	.10	(.07, 13)

Notes: Numbers in parentheses are 95% CI.

4.2 Modeling the hazard of migration

Table 4 presents coefficients for effects on the hazard of migration. The left-hand column presents the single-process estimates and the right-hand column presents the multi-process estimates. The random effect of migration in the single-process model (σ_{mig} = .13 in Table 3) indicates that individuals who changed their residence in the past are significantly more likely to move again due to unmeasured person-specific characteristics.

[6] We test the differences in coefficients between the single-process and multi-process models using the Hausman test. Both differences are statistically significant.

Table 4: Estimates from models for migration

Variable	Single-process Coeff.		Single-process O.R.	Multi-process Coeff.		Multi-process O.R.
Marriage						
Married 1 year before migration	.60 (.16)	***	1.82	.28 (.08)	**	1.33
Total number of marriages	-.08 (.05)	†	.92	-.01 (.02)	ns	.99
Duration						
Duration to migration	-.11 (.02)	***	.90	-.09 (.01)	***	.91
Squared duration to migration	.01 (.00)	**	1.01	.01 (.00)	***	1.01
Cubical duration to migration	-.00 (.00)	***	1.00	-.00 (.00)	***	1.00
Individual characteristics						
Female	-.06 (.04)	ns	.94	-.05 (.02)	**	.95
Black	-.49 (.04)	***	.61	-.21 (.02)	***	.81
Hispanic	-.31 (.05)	***	.73	-.14 (.02)	***	.87
High school or equivalent	.52 (.04)	***	1.69	.26 (.02)	***	1.29
College or more	.27 (.04)	***	1.31	.18 (.02)	***	1.20
Employed part-time	.21 (.04)	**	1.24	.11 (.02)	***	1.12
Employed full-time	-.35 (.05)	***	.70	-.15 (.02)	***	.86
Living in metro areas	-.20 (.04)	***	.82	-.09 (.02)	***	.92
Giving a birth	-.04 (.08)	ns	.96	-.04 (.04)	ns	.97
Total number of children	-.13 (.02)	***	.88	-.04 (.01)	***	.97
Household characteristics						
R lived in an intact family until age 18	-.17 (.03)	***	.84	-.08 (.01)	***	.92
Mother's education	.05 (.01)	***	1.05	.02 (.00)	***	1.02
County characteristics						
Sex ratio (%)	-.00 (.01)	ns	1.00	-.00 (.00)	ns	1.00
Population with a college degree/higher (%)	.02 (.00)	***	1.02	.01 (.00)	***	1.01
Unemployment rates (%)	.01 (.01)	ns	1.01	.00 (.00)	ns	1.00
Poverty rates (%)	.01 (.00)	***	1.01	.00 (.00)	***	1.00
Crime rates (%)	-.01 (.00)	ns	.99	-.00 (.00)	ns	1.00
Intercept	-2.90 (.49)	***	.06	-1.31 (.19)	***	.27
Log likelihood	-22900.9			-46092.4		
Wald Chi2 (22)	1709.5			4303.4		
Person-years	87,931					
Number of observations	7,827					

Notes: Numbers in parentheses are standard errors.
ns = not significant, † p≤.10, * p≤.05, ** p≤.01, *** p≤.001

Controlling for the unmeasured personal propensity to migrate in the single-process model, marriage increases the hazard of migration by about 82%. Even in the multi-process model, marriage is a significant determinant of migration, although the size of the coefficient decreases by more than half (b = .60 and b = .28 in single-process and multi-process models respectively).[7] The positive relationship between migration and marriage is consistent with the well-known process of newlyweds establishing their household in a new place. In contrast, the total number of marriages is negatively related to the hazard of migration in the single-process model (b = -.08, p = .073). But this variable becomes much smaller and loses statistical significance once the correlation between the propensity to move and to marry is accounted for in the multi-process model (b = -.01, p = .564).

The coefficients for the other explanatory variables in the migration equation change once unobserved heterogeneity is accounted for in the multi-process model, although the signs of most coefficients remain the same. Females are less likely than males to move and the statistical significance becomes stronger in the multi-process model (from b = -.06, p = .141 to b = -.05, p≤.01). African-Americans and Hispanics are less likely than their non-black, non-Hispanic white counterparts to move, and the racial differences persist after accounting for the correlation between the migration and marriage random effects. To the extent that migration is an investment for purposes of socioeconomic gain (Cadwallader 1992), these results suggest that women and black and Hispanics are less likely to make human capital investments that require moves (Quinn and Rubb 2011; Shauman and Noonan 2007).

Greater individual and household socioeconomic resources increase the hazard of migration in both the single-process and the multi-process models. Those who have completed high school or college and who are employed part-time are more likely to move compared to those with less than a high school diploma and those not employed respectively. Full-time employment, however, decreases the hazard of migration (as compared to being unemployed). Taken together, these results are consistent with a change in residence as a job search strategy. Those who have acquired sufficient human capital (education, employment experience) tend to settle down rather than move around (Schachter 2001). The effects of household characteristics become weaker once the simultaneous relationship between migration and marriage is accounted for. Those from an intact family are less likely to move (b = -.08, p≤.001 in the multi-process model), although an increase in maternal educational attainment is positively related to the hazard of migration (b = .02, p≤.001 in the multi-process model). Finally, residence characteristics are significantly related to the hazard of migration of individuals. An increase in the share of population with a college or higher degree increases the likelihood of migration by 1% (b = .01, p≤.001 in the multi-process model). The hazard

[7] Statistically significant difference is tested according to the Hausman test.

of moving is also higher where county-level poverty rates are higher (b = .004, $p \leq .001$ in the multi-process model). However, sex ratio, unemployment, and crime rates in the county of residence do not affect the hazard of migration in either the single-process or the multi-process specification.

4.3 Modeling the hazard of marriage

The results from both the single-process and multi-process models for marriage are shown in Table 5. Again, the person-specific random effects from the equations are presented in Table 3. The random effect of marriage (σ^{marr} = .10) indicates that unmeasured factors make some individuals more prone to marry – that is, those with a higher hazard of marrying in the past have a higher hazard in the present episode of marriage.

In Table 5, we find that migration is positively associated with the hazard of marriage (b = .11, $p \leq .05$): those who moved one year ago are approximately 12% more likely to marry. However, once correlation between the random effects for migration and marriage is accounted for in the multi-process model, the significant effect of migration on marriage disappears (b = .04, p = .584). This suggests that the positive effect of past migration on the hazard of marrying, according to the single-process model, actually reflects association between the two propensities due to other unmeasured factors. The total number of moves does not appear significant in either the single or the multi-process model. Frequent movers may face various new environments to which they must adapt and this process may hinder other major decisions such as marriage, which in turn shows no association between marriage and migration in our model. We note again that the estimated causal effects of migration on marriage are upwardly biased if we regard the multi-process as preferred (because it allows for correlation between the two propensities).

Table 5: Estimates from models for marriage

Variable	Single-process Coeff.		Single-process O.R.	Multi-process Coeff.		Multi-process O.R.
Marriage						
Married 1 year before migration	.11 (.05)	*	1.12	.04 (.03)	ns	1.05
Total number of marriages	.02 (.02)	ns	1.02	.00 (.01)	ns	1.00
Duration						
Duration to migration	.18 (.02)	***	1.20	.07 (.01)	***	1.07
Squared duration to migration	-.01 (.00)	***	.99	-.01 (.00)	***	.99
Cubical duration to migration	.00 (.00)	***	1.00	.00 (.00)	***	1.00
Individual characteristics						
Female	.29 (.04)	***	1.34	.12 (.02)	***	1.12
Black	-.83 (.04)	***	.43	-.36 (.02)	***	.70
Hispanic	-.16 (.05)	**	.85	-.08 (.02)	***	.93
High school or equivalent	.34 (.04)	***	1.40	.16 (.02)	***	1.17
College or more	.26 (.04)	***	1.30	.12 (.02)	***	1.13
Employed part-time	.44 (.05)	***	1.55	.21 (.02)	***	1.23
Employed full-time	.79 (.05)	***	2.21	.40 (.02)	***	1.49
Living in metro areas	-.11 (.04)	*	.90	-.05 (.02)	**	.95
Giving a birth	.16 (.07)	*	1.17	.06 (.03)	†	1.06
Total number of children	-.01 (.02)	ns	.99	.01 (.01)	*	1.01
Household characteristics						
R lived in an intact family until age 18	-.06 (.03)	†	.94	-.03 (.01)	*	.97
Mother's education	-.02 (.01)	***	.98	-.01 (.00)	***	.99
County characteristics						
Sex ratio (%)	.02 (.01)	†	1.02	.01 (.00)	ns	1.01
Population with a college degree/higher (%)	-.01 (.00)	**	.99	-.01 (.00)	**	.99
Unemployment rates (%)	-.00 (.01)	ns	1.00	.00 (.00)	ns	1.00
Poverty rates (%)	.01 (.00)	*	1.01	.00 (.00)	*	1.00
Crime rates (%)	-.01 (.00)	*	.99	-.00 (.00)	*	1.00
Intercept	-4.05 (.49)	***	.02	-1.99 (.20)	***	.14
Log likelihood	-23270.2			-46092.4		
Wald Chi2 (22)	1396.4			4303.4		
Person-years	87,931					
Number of observations	7,827					

Notes: Numbers in parentheses are standard errors.
ns = not significant, † p≤.10, * p≤.05, ** p≤.01, *** p≤.001

Accounting for personal propensity also reduces the estimated effects of most other explanatory variables in the marriage equation, although the directions of effects are stable. In both equations, females have a higher hazard of marrying – that is, earlier marriage – than males, and African-Americans and Hispanics have a lower hazard of marrying than non-black, non-Hispanic whites. Individual socioeconomic characteristics are also significantly related to the hazard of marriage. Those having a high school diploma or a college and higher degree have a higher hazard of marrying – that is, earlier marriage – than those with less than a high school diploma. Moreover, employees (either part-time or full-time) are more likely than those not employed to marry. Indeed, employment status in the preceding year is the most potent factor affecting the hazard of marriage: the odds ratio increases by 23% and 49% for part-time and full-time employees (in the multi-process model) respectively. Childbirth increases the hazard of marriage ($b = .06$, $p = .051$ in the multi-process model) and total number of children is positively related to the hazard of marriage ($b = .01$, $p \leq .05$ in the multi-process model).

Household and residence characteristics are also significantly related to the hazard of marriage, although the effects generally are smaller in magnitude than the effects of individual characteristics. Comparing the single- and multi-process models indicates that failure to account for unmeasured person-specific traits results in an overestimation of the effects on marriage of household and county characteristics. Maternal education is negatively associated with the hazard of marriage ($b = -.01$, $p \leq .001$ in the multi-process model) and having lived in an intact family decreases the hazard of marriage ($b = -.03$, $p \leq .05$ in the multi-process model). The hazard of marriage is negatively related to the share of the population with a college or higher degree and to the local crime rate ($b = -.01$, $p \leq .01$ and $b = -.005$, $p \leq .05$ in the multi-process model respectively). On the other hand, the higher poverty rates in the county of residence, the higher the hazard of marrying ($b = .002$, $p \leq .05$ in the multi-process model). Sex ratio and unemployment rates in the county are not significantly related to the hazard of marriage in the multi-process model.

4.4 Robustness of the results

Although identification without exclusion restrictions works well, to ensure the robustness of the results we estimate the multi-process model with selected explanatory variables excluded (and therefore serving to identify the model). We also consider a religion variable as a further basis for model identification.[8] First, we assume that the sex ratio in the place of residence is directly related to individual marital decisions

[8] Because religion is only relevant to marital transitions, we exclude it in the final estimates.

(South and Lloyd 1992) but not to migration, and therefore can serve to identify the marriage equation. When considering migration, people take into account the local environment – such concerns as safety of the residence or milder weather (Chen and Rosenthal 2008; Whisler et al. 2008). Therefore we include county crime rates in the migration equation as an identifying variable. Religion can be an appropriate identifying variable if one assumes that religion is a determinant of the decision to marry but not to migrate (Lehrer 2004; Wilcox and Wolfinger 2007). As it turned out, the regression results are remarkably robust to these alternative specifications; none of them results in markedly different coefficient values. On the basis of this set of tests, we conclude that our results are robust to alternative model identification strategies.

Moreover, we examined the effect on estimates of the treatment of cohabitation-related moves. In the estimates presented above, we excluded moves occurring one year after the onset of cohabitation. Moves occurring in the same year as the onset of cohabitation may also intervene between cohabitation and marriage, and therefore should be excluded. Excluding these additional moves has a minimal impact on the estimates (coefficients differ only slightly). In short, our results are robust as to the choice of criteria for identifying cohabitation-related moves.

5. Discussion

The current study investigates relationships between migration and marriage in the United States, allowing for complex interdependency between the two life course histories. A few studies using European samples have allowed for complexity of the form incorporated in the present study, but no such study has been conducted using U.S. data. Thus the findings from this study not only improve our understanding of the association between migration and marriage, but also add a country case study that, when compared to results in other settings, will begin to illuminate cross-country variability in the nature and strength of the interdependency between these two processes. A virtue of this study is the long observation period afforded by the NLSY79 – roughly three decades.

First, findings from this study reveal that the multi-process model which accounts for unobserved propensity to move and to marry yields different estimates than a single-process model. This suggests that significant unmeasured correlation between migration and marriage exists, reflected in a tendency for individuals to migrate and marry in the same year. This is consistent with results from a previous study using a West German sample (Mulder and Wagner 1993). Accounting for endogeneity between life course transitions yields less biased estimates, whereas ignoring this form of interdependency leads to a mistaken impression that there are direct effects between certain features of each history on the other. Future research on the life course will benefit from accounting for interrelationships between life events.

Second, we find that marriage positively affects the risk of migration in the short term, but over the longer term the total number of marriages does not significantly affect the hazard of migration. This finding suggests that migration occurs in anticipation of family events but over the long haul increasing social ties via family events can lower the risk of moving. However, the sample of persons experiencing multiple marriages is small, so this result should be viewed with caution. Cohabitation is common in the United States and a large proportion of marriages begin with cohabitation, especially remarriages (Smock 2006; Cherlin 2010).There is evidence of a growing tendency in the U.S. to remain in a cohabiting status rather than move to marriage if the partnership is second or higher order and especially if one or the other partner was previously married (Smock 2006; Cherlin 2010). Which means that incorporating the cohabitation history in research on migration–marriage interrelations would be desirable. This is not feasible with the NLSY79, however: while the NLSY79 provides a range of life course transition histories, the cohabitation histories are inadequate for our purposes, for multiple reasons. First, cohabitation history in the NLSY79 was constructed through retrospective inquiry in 1990, which raises concerns about its completeness and accuracy. Second, there is non-comparability introduced by changes in the definition of 'partner' across survey years (Center for Human Resources 2013). Moreover, only from 2002 have respondents been asked detailed questions about brief cohabitation episodes (three months) (Center for Human Resources 2013). Therefore, we did not incorporate the cohabitation histories in this research, except for removing cohabitation-related migration. Certainly it would be desirable to have future research which investigates the extent to which the relationship between migration and union transitions is affected by cohabitation experiences, thereby drawing a more complete portrait of interdependency between life course transitions.

Third, regarding the effects of migration on marriage, we find no significant effects once unobserved heterogeneity has been controlled via the multi-process model with random effects. Our expectation that migration is a strategy to rectify failed marriage prospects is not supported. In the previous literature, the effects of moving on union formation have been understood as an anticipation or plan for family changes (Feijten and Mulder 2002). In our models, the possible anticipating or planning effect is partially ruled out when we discount moves which are preceded by cohabitation. Furthermore, an inclusion of unobserved personal components in the model controls for possible endogeneity of the planned behaviors. The absence of significant effects of migration on marriage, therefore, suggests that migration is not motivated much by pursuit of marriage opportunities but rather is a life course transition driven by other life decisions. This conclusion must be treated warily: the time-metric is a year, and a clear portrayal of migration–marriage interrelations may require more fine-grained measurement (months or even weeks). For example, using Finish data, Kulu and Steele (2013) found that couples are more likely to get pregnant during the first months after a move, but the risk of pregnancy decreases and becomes stable afterwards – that is, one year after the move. In addition, the relationships between migration and marriage may vary by types of mobility – for example, within-county versus between-county moves. The two different types of mobility have had different purposes in the U.S.: between-county moves are predominantly motivated by employment considerations, while within-county moves often occur as a response to changes in the family configuration, such as the birth of a child (Schachter 2001). More precise measurement of migration and union histories – for example, monthly data and clearer distinction among different forms of migration and partnership – will provide a stronger foundation for a valid assessment of the association between these two life course transitions.

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