Bringing Intergenerational Social Mobility Research into the Twenty-first Century: Why Mothers Matter

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Conventional social mobility research, which measures family social class background relative to only fathers' characteristics, presents an outmoded picture of families—a picture wherein mothers' economic participation is neither common nor important. This article demonstrates that such measurement is theoretically and empirically untenable. Models that incorporate both mothers' and fathers' characteristics into class origin measures fit observed mobility patterns better than do conventional models, and for both men and women. Furthermore, in contrast to the current consensus that conventional measurement strategies do not alter substantive research conclusions, analyses of cohort change in social mobility illustrate the distortions that conventional practice can produce in stratification research findings. By failing to measure the impact of mothers' class, the current practice misses a recent upturn in the importance of family background for class outcomes among men in the United States. The conventional approach suggests no change between cohorts, but updated analyses reveal that inequality of opportunity increased significantly for men born since the mid-1960s compared with those born earlier in the century.

To understand current inequality, social mobility researchers must bridge a longstanding gap between theory and practice that increasingly distorts social mobility and stratification research findings.[1] A gap exists because, in theory, class background (i.e., childhood class position) is a family-level variable, but the conventional research practice equates class background solely with a father's class position. This assumes that mothers' economic participation is not common or important to class background and that father-headed families are the norm. Yet in the United States, rates of labor force participation among mothers have steadily increased since intergenerational class mobility models were first developed. Figure 1 illustrates this trend.

The bias that comes from excluding mothers' class characteristics is increasingly important but not widely recognized or understood. Scholarly debate over the conceptualization and measurement of family-level class position has waned since the early 1990s, with a general consensus that the conventional mobility

Direct all correspondence to Emily Beller (ebeller@berkeley.edu). This research was supported by a National Science Foundation Dissertation Improvement Grant (0403401). An earlier version of this article was presented at the 2007 meeting of the International Sociological Association Research Committee 28 on Social Stratification and Mobility (RC28), Montreal, Canada. I thank Claude Fischer, Harry Ganzeboom, Caroline Hanley, Michael Hout, Jerome Karabel, Sam Lucas, Robert Mare, Donald Treiman, Jane Zavisca, and four anonymous reviewers for helpful comments, advice, and suggestions on this research. The views expressed are solely those of the author and do not represent the views of the U.S. Government Accountability Office.

[1] The concept of social mobility rests on the idea that social positions exist and can be differentiated from one another, but there are a variety of ways to do so. Here, I define social position in terms of occupational classes, but the ideas discussed are applicable to other definitions (e.g., socioeconomic status).

Figure 1. Percent of Respondents with a Mother Who Worked Outside the Home, by Year of Birth

Source: General Social Surveys 1994 to 2006.

research practice remains adequate. Furthermore, there are practical incentives, such as data considerations, to follow the conventional practice. For example, until 1994 the General Social Survey did not ask respondents about their mothers' occupations. In this article, I demonstrate empirically that fathers' class is an increasingly poor proxy for family social class background in the United States, and research conclusions can be distorted when it is used as such.

Given practical considerations such as limited data on mothers' occupations, it is important to emphasize that adequately defining and measuring class background to bring it in line with theory is essential for substantive, rather than simply methodological, reasons (e.g., improving explained variance or model fit in and of itself). To illustrate this point, I examine changes in intergenerational class mobility between recent birth cohorts in the United States and show that the typical measurement strategy masks important findings. Without updating intergenerational class mobility models to incorporate mothers' class characteristics, a recent, significant upturn in the importance of class background for men's class destinations in the United States is not evident. The conventional approach indeed suggests there was no change in the extent of mobility between cohorts. Updated analyses, however, reveal that class mobility declined for men born since the 1960s compared with those born earlier in the century.

BACKGROUND

Structural Mobility and Social Fluidity

Intergenerational class mobility research analyzes the strength of the association between individuals' class background or childhood class position (class origin) and their current individual or family-level class position (class destination), as well as patterns of immobility or movement between particular origins and destinations. Such intergenerational class mobility has two components worth examination. The first, often called social fluidity, refers to the extent to which an individual's chances of reaching a particular class destination are constrained by class background. The second component, often called structural mobility, captures shifts in the distribution of class origins and destinations that affect everyone's mobility, regardless of class background (e.g., upgrading of the economy toward better jobs). In the analyses that follow, I hold structural mobility constant, thus centering attention on social fluidity. Social fluidity is generally considered indicative of equality of opportunity; more fluidity is therefore "better." However, perfect or very high social fluidity—the absence or near absence of association between origins and destinations—is neither plausible nor, arguably, desirable, given that some of the processes leading to the intergenerational persistence in class position, such as inherited cognitive ability, may be legitimate (Harding et al. 2005; Roemer 2004).

Because there is no external benchmark, such as perfect mobility, to aid in interpretation, estimates of social fluidity mean little in and of themselves. Yet such estimates become instructive in comparative context. For example, researchers compare the social fluidity levels of different periods or cohorts to assess whether equality of opportunity is increasing or decreasing over time (e.g., Breen and Jonsson 2007; Hout 1988). Researchers also often compare social fluidity levels between countries (e.g., Breen 2004; Erikson and Goldthorpe 1992). The comparative nature of social mobility research is an important reason why moving the measurement of class background more closely in line with theory is a matter of more than simply technical or methodological interest. When measurement is biased, the extent to which it is biased can vary between the cohorts, countries, or other groups being compared. As a result, researchers could misinterpret differences in measurement error between groups as substantive differences, or a lack thereof, in social fluidity. In this article, I use the example of change in social fluidity between successive birth cohorts, about which little is known, to illustrate this very concern.

Comparing Social Fluidity Rates over Time

Research on change over time in social fluidity typically uses either a period or a cohort approach. In the period approach, researchers compare social fluidity levels between different survey years (i.e., periods). In the cohort approach, survey data collected in different years or periods is pooled, but respondent birth cohort is held constant. Period-related shifts in mobility apply to individuals across the board at a given point in time, independently of their birth cohort, while cohort-related shifts in mobility trends arise from the different experiences of individuals born in specific cohorts. Breen and Jonsson (2007) argue convincingly that changes over time in social fluidity are more likely to be cohort-driven than period-driven.

Research on change over time between periods and cohorts demonstrates that social fluidity increased over the course of the past century in the United States, until about the mid-1980s (from a cohort perspective, for individuals born up until about 1960; DiPrete and Grusky 1990; Featherman and Hauser 1978; Hout 1988). Trends in social class fluidity after the mid-1980s are unclear, in part because changes to the census coding of occupations in the 1980s made it impossible to directly compare new survey data with older data (Vines and Priebe 1988). Some research suggests a possible slowing of the trend of increasing social fluidity after the mid-1980s (Hout 1996); others predict a continued trend of increasing fluidity due to a growing proportion of individuals raised in non-intact families, who appear more mobile than their peers raised in intact families (Biblarz and Raftery 1999).

Defining Social Class

My own analyses of the possibilities suggested above employ Erikson and Goldthorpe's (1992) class schema, which is widely used in social mobility research and often called the EGP class schema in reference to an early explication of it (Erikson, Goldthorpe, and Portocarero 1979). The theoretical basis for the EGP class schema has been linked to the Weberian view that classes can be meaningfully differentiated according to the market resources and, consequently, the life chances of their members (see Breen 2005).[2] Classes, then, are not defined

[2] Erikson and Goldthorpe (1992:37) cite both Weber and Marx as sources for the principles of class differentiation on which the EGP schema is based.

with respect to particular workplace tasks, roles, or experiences per se, but are defined according to the resources that are consequences of work. Correspondingly, the EGP schema defines classes primarily in terms of the types of employment relationships that characterize them—with the logic that different employment relationships entail different rewards, opportunities, and constraints—and the authors emphasize that class experiences are not restricted to the workplace (Erikson and Goldthorpe 1992:236).

The EGP schema starts with a basic distinction between the self-employed and employees; among employees, a key additional distinction is made between two types of relationships with employers. Employment in a professional or service context, in which employees have some degree of autonomy and advantageous resources such as employment security, career advancement prospects, pensions, and salary increments, is distinguished from employment regulated by labor contracts, which is under close supervision, in return for piece wages, and lacks the advantages of the service relationship (see Goldthorpe [2000] for discussion of how occupational conditions lead to these two employment relationships).

The distinction between occupations involving service versus labor-contract relationships is further refined in two ways. First, professional occupations involving a service relationship are divided into two classes (higher- versus lower-level professionals, managers, and administrators) based on the extent of expected advantageous resources. Similarly, occupations regulated through labor contracts are divided into two classes of skilled and unskilled manual workers in recognition that some beneficial modifications to the labor contract are likely for skilled workers. Second, occupations in which the distinction between service versus labor-contract employment relationships is blurred are also included in intermediate classes (e.g., one class includes the administrative positions that support professional bureaucracies; another covers supervisory manual and technical occupations). The fully elaborated schema includes 11 class categories that may be collapsed into fewer categories for research purposes (Erikson and Goldthorpe 1992:35–47).

CONCEPTUALIZING AND MEASURING CLASS AT THE FAMILY LEVEL

Erikson and Goldthorpe, as noted above, argue against a workplace centered view of class; they also maintain that the family, rather than the individual worker, is the unit of class "fate" (Erikson and Goldthorpe 1992:233). While class experiences follow from family members' involvement in different types of employment relationships, they are not limited to the workplace; they also include, for example, broader experiences of economic security or insecurity, affluence or poverty, and prospects for improvement in economic circumstances. Members of a family who live together experience similar resources and future life chances. The class position of family members without occupations, including children, and wives who are not employed outside the home, can thus be considered dependent on the resources accrued from the household head's employment relationship. In addition, spouses in dual-earner families, whose individual occupations may be substantially different in terms of employment relationships, also share one class position because of their shared experiences of resources or constraints. While some argue that the idea of a shared family-level class position should be abandoned in favor of measuring class in terms of individual occupations (Acker 1973; Stanworth 1984), the individual approach is not possible when applied to class origins (given that children do not have occupations).

Debate over Joint versus Conventional Measurement of Shared Family Class Position

Scholars who agree theoretically that families share both class experiences and a class position still debate how shared family class should be measured (Sorensen 1994). Goldthorpe (1983, 1984) initially argued that the shared family-level class position is determined by a father's or husband's class (this was termed the conventional view of family class). This position sparked debate, with others arguing for a joint approach to measuring the shared family class position, in which both spouses, if employed, contribute to family class (Britten and Heath 1983; Davis and Robinson 1988, 1998; Heath and Britten 1984). Research offers mixed support for both the conventional and joint viewpoints, leading to a conclusion that the conventional measurement of class may not strongly distort research results (Sorensen 1994). Conventional scholars remain in favor of measuring the family-level class with respect to only one spouse (rather than jointly based on two spouses' occupations) but allow that the "dominant" class position—that of the spouse with the stronger labor force attachment and higher individual class position—could potentially determine family class position, rather than necessarily the husband (Erikson 1984; Erikson and Goldthorpe 1992). They argue that the joint approach to determining the family-level position, while attractive in principle, may blur class boundaries and create too many possible class positions (Erikson and Goldthorpe 1992:238).[3]

Importantly, conventional practice has more potential to distort social mobility research conclusions than is clear from the previous debate, which had important limitations. Core questions addressed in the debate regarding family-level social class include: (1) how to measure the proportion of mixed-class families to determine the significance of the problem they might pose to researchers, and (2) how mixed-class families might affect substantive research results, particularly given findings that the subjective class identification, class related behavior (e.g., voting), and life chances of married women can be better predicted by their husbands' than by their own occupations (Baxter 1994; Erikson and Goldthorpe 1992; Goldthorpe 1983, 1984; Heath and Britten 1984; Stanworth 1984).

One limitation of the prior debate is that it focused on adults' class positions without also considering the position of children—despite the fact that intergenerational class mobility research is centrally concerned with the influence of family class position on children's future life chances. For example, as Erikson and Goldthorpe (1992:250, note 16) note, researchers critiquing the conventional view of class destinations nonetheless rely on conventional measures of class origins. Given a focus on how class affects future life chances, the adequacy of different approaches to measuring the shared family class position must be evaluated from the perspective of children, as well as adults.

Another limitation of the prior debate is that it focused on families in which adults were employed in different classes ("mixed-class" families) over class-consistent families or families with only one spouse in the labor market (single-earner families). Families with only one employed spouse, or with both spouses employed in the same class, were not considered problematic. Furthermore, the class position of a dual-earner family with both spouses employed in a particular class was considered equivalent to the position of a single-earner family with one spouse employed in that class. Sorensen (1994:43) characterizes this assumption as surprising, noting that a major reason for developing new measures of families' class positions is the hypothesis that women's employment makes a difference for families' material circumstances and life chances. That is, if the joint perspective is correct, it should logically apply not only to mixed-class but also to class-consistent families. For example, if individual spouse class characteristics jointly define family class position, dual-earner families where both adults are employed in the higher professional class might be expected to have a more advantaged class position than would families with one spouse employed in the higher professional class and another spouse who is either not in the labor market or is employed in a less advantaged occupation. The idea that each spouse's employment relationships could produce cumulative class resources or constraints illuminates the implications of the conventional versus joint measurement.

Theoretical Implications of Conventional and Joint Measurement of Family Class Origin

Breen and Jonsson (2007) propose a theoretical model of social mobility in which arrival at a particular class destination depends on class-related parental resources that can be either directly (e.g., genetics or property) or indirectly transmitted between generations. The role of indirect transmission in this process reflects the idea that parents' class experiences and consequent class-related resources influence the extent to which the next generation can accumulate assets, such as higher education, which,

[3] See Breen and Rottman (1995) and Sorensen (1994) for more detailed reviews of this debate.

in turn, generate particular returns in class destinations. This model helps clarify the theoretical implications of conventional and joint measurement of the shared family class position with respect to class origin in particular. Joint measurement of family class origin accounts for the possibility that each parent's employment relationships result in class-related resources and assets that may accumulate—regardless of whether the parents' occupations fall into different classes or are class-consistent. By contrast, the conventional measurement practice assumes that, net of the key (father's or higher) class position, a second parent's employment does not result in additional class-related resources. If this assumption is incorrect, conventionally measured class position could be understood to serve as a proxy for a more complex set of family class resources that are products of both the measured class position and an unmeasured second parent's class position.

EVIDENCE OF A CUMULATIVE IMPACT OF PARENTS' CLASS-RELATED RESOURCES

Theorized mechanisms of the intergenerational class transmission process are consistent with the joint view of family class origin. Class-related economic resources clearly accumulate and play a key role in indirect and direct transmission of resources and assets (Conley 2001; Hill and Duncan 1987), but economic resources are only part of the story (Mayer 1997). Class-related noneconomic resources, such as occupational prestige and parent education (often termed cultural resources), might also play a role in indirect transmission. Individual parent cultural resources could accumulate; many theorize that advantaged parents provide children with advantageous cultural resources through interactive processes (Bourdieu and Passeron 1977; Lareau 2003). For example, middle- and upper-class parents may intentionally cultivate children's social skills such as addressing and negotiating with authority figures (Lareau 2003), have greater knowledge of educational bureaucracies (Deil-Amen and Rosenbaum 2003; Lareau 1989; Lucas 1999), and hold high aspirations for their children (Hauser, Tsai, and Sewell 1983; Sewell, Haller, and Portes 1969). The impact of cultural resources could add up, particularly if both parents spend time with their children.

Prior research indeed provides empirical evidence that both parents' class characteristics influence children's class-related resources, assets, and eventual class destinations, even if the parents' class characteristics are the same. Among employed parents, for instance, both parents' occupations independently shape children's educational outcomes (Kalmijn 1994; Korupp, Ganzeboom, and Van Der Lippe 2002), just as both parents' education levels do (Mare 1981). Models of occupational mobility better predict class destinations for both sexes (Khazzoom 1997) when the models include mothers' occupations. Although many mothers do not have occupations outside the home, the theorized role of parent–child interactions in the intergenerational transmission process raises the question of whether the joint view of family class origin might apply even to single-earner families—that is, homemaker mothers may contribute their own class resources despite not having an individual employment-based class position. Dynamic views of class (Marshall, Roberts, and Burgoyne 1996; Plutzer and Zipp 2001) posit that a series of experiences such as childhood class background, education, and previous employment or unemployment spells contribute to one's class and associated noneconomic class resources. In either case, and given the theorized role of noneconomic or cultural class resources in the mobility process, it is reasonable to test whether nonemployed parents may contribute to the transmission of class-related resources, rather than assume they do not.

IMPLICATIONS FOR INTERGENERATIONAL SOCIAL MOBILITY RESEARCH

The theoretical and empirical evidence described above suggests that parental class resources may jointly determine family class origin. Because the prior debate over the joint versus conventional measurement strategies did not problematize class-consistent families, assortative marriage patterns (which result in a high prevalence of class-consistent families) may appear to justify conventional measurement. Conventional practice, however, would actually be less problematic if marriage were random with respect to class; class-based assortative marriage patterns mean that the measurement error produced by the conventional practice is not random; it probably changes over time and differs among groups.

If parent class characteristics and associated class resources indeed jointly determine the family class position, conventional estimates of the strength of father–child association in class position will include the correlated but unmeasured effects of mothers' class resources on the process. Because the correlation between mothers' and fathers' individual class positions is not perfect, the conventional measurement of family class origin will underestimate the total origin–destination association. Furthermore, if marital sorting by class differs between comparison groups (e.g., nations, cohorts, or racial/ethnic groups), the use of conventional origins measures could lead analysts to erroneously interpret changes in the degree of measurement error to be substantive differences between groups in social fluidity levels.

ANALYTIC STRATEGY

DATA AND METHODS

The data from which I draw are a compilation of the available years of the General Social Survey (GSS) that include mothers' occupational data—1994, 1996, 1998, 2000, 2002, 2004, and 2006 (Davis, Smith, and Marsden 2007). Occupational data collected in these survey years were recorded in 1980 basis census codes. I conduct separate analyses for men and women, and I restrict analyses to respondents who were ages 25 to 64 and in the labor force at the time of the survey. I also restrict the analyses to respondents with valid data for both their own and two parents' (or parental figures') occupations, or, in the case of mothers, homemaker status. Missing data for parent occupation due to item nonresponse is minimal, affecting approximately 2.5 percent of the otherwise eligible sample. However, the GSS, like most surveys, does not ask about the occupations of noncustodial parents. Therefore, without making the strong assumption that noncustodial parents do not shape class background, respondents who lived with a single parent (or in an institution, in which case no parent occupation data was collected) at age 16 cannot be included in analyses that focus on comparing two-parent versus one-parent measures of family class position. To partially address this limitation, I include single-parent families in the final portion of the analysis, which examines change over time in social fluidity rather than comparing one-parent versus two-parent measures of class origins.

I adopt a six-category version of the Erikson and Goldthorpe class schema, described above, to define classes. I generated the EPG classes on the basis of the GSS International Standard Classification of Occupations (ISCO) 88 codes for occupations, together with self-employment information, from a widely used conversion algorithm (Ganzeboom and Trieman 2003).[4]

The class categories are the following:[5]

- I Professionals, administrators, officials, and managers, higher level;
- II Professionals, administrators, officials, managers, lower level;
- IIIab Routine nonmanual and service workers, higher and lower levels;
- IVab Self-employed, with or without employees;
- V/VI Technical specialists and supervisors of manual workers, skilled manual workers; and
- VIIab Semiskilled and unskilled manual workers, nonfarm and farm.

Separately for men and women,[6] I organize these data into a three-way intergenerational class mobility table by cross-classifying the mother's class category variable by the father's class category variable by the respondent's class category variable (Table A1 in the Appendix shows the distributions of class positions among

[4] The GSS data from 1994 through 2006 include detailed occupational information recorded in 1980 basis U.S. census codes. The GSS also converts these 1980 basis census codes into ISCO 88 codes. An algorithm (unpublished, available from the author) to directly convert occupational data recorded in 1980s basis census codes into EGP classes produces similar results. Erikson and Goldthorpe (1992:315–16) encountered some problems applying their schema to United States occupational data recorded in 1960 basis census codes, but the 1980 basis census codes do not present these problems.

[5] Apart from a small number of farm workers (class VIIb), I exclude agricultural classes due to limited data. Some research combines class IIIb with class VII instead of IIIa; this alternate categorization does not change the key results.

[6] A study of the validity of the class schema reports it to be valid for both men and women (Evans and Mills 1998).

male and female respondents of the appropriate age and labor force status, and those of their mothers and fathers).

I use Goodman's (1979) log-multiplicative RC association model (also called the RC-II model) to analyze the mobility tables described above. To illustrate the RC model, consider a simpler two-way contingency table such as the conventional intergenerational mobility table of father's class (i) by class destination (j). The RC model simultaneously estimates row scores (μ_i) that rank father's class (origin) categories and column scores (ν_j) that rank class destination categories, along with an intrinsic association parameter (Φ). The association parameter conveys the overall strength of the relationship between the ranked class origin and destination categories, and it is interpreted similarly to a regression coefficient in that a larger value means greater association (Hout 1983).[7]

While typically used to analyze grouped data such as mobility tables, the RC model can be extended to incorporate individual-level covariates using various techniques such as including stereotype ordered regression (SOR) parameters (Breen 1994; DiPrete 1990; Hendrickx and Ganzeboom 1998). I incorporate SOR parameters in RC models in some analyses to control for age, which becomes important in models that compare fluidity between cohorts. The SOR parameter is analogous to the RC association parameter (Φ), in that it parsimoniously expresses the overall effect of an independent variable k on all categories of the dependent variable in a single parameter—where the RC association parameter indexes the strength of association between μ_i and ν_j, the SOR parameter indexes the strength and direction of association between k and ν_j.

The mobility tables I analyze are somewhat sparse, due primarily to clustering of women in certain classes, so I assess overall model fit using the Pearson chi-squared goodness of fit statistic X^2, rather than the likelihood-ratio goodness of fit statistic L^2, but I compare nested models using L^2 (Agresti and Yang 1987).[8] To incorporate the GSS case weight variable without distorting these model fit statistics, the counts in the mobility tables are the unweighted frequencies and the models include weight vectors containing average cell weights (Clogg and Eliason 1987). X^2 and L^2 are appropriate fit statistics given grouped data, but they are not applicable to the models with SOR parameters given that such models include individual-level data. I also use the BIC criterion (Raftery 1995), in concert with the other fit statistics as applicable, to adjudicate among models. Given the sample sizes, marginal differences in BIC (of fewer than approximately 10 points) can be considered equivalent (Wong 1994). I use the LEM program (Vermunt 1997) for the analyses of grouped data, and Stata for the SOR analysis.

MEASURES OF CLASS ORIGIN

I analyze several different measures of class origin that fall into three categories: those based on

[7] I use log-multiplicative RC association models rather than log-linear models because the association parameter Φ of the RC model is readily interpretable as a descriptor of the overall strength of association between class origins and destinations. This feature of the RC model is key to illustrating the consequences of various origins measures. One limitation of the RC model, however, is that, in summarizing origin and destination categories in terms of ranked scores, it analyzes only one hierarchical dimension of origin-destination association (multidimensional RC(m) models are possible, but are not as easily interpretable). I replicated the analyses presented in this article using both log-linear models, which do not impose a unitary hierarchical dimension of association, and Erikson and Goldthorpe's (1992) core social fluidity model, which includes multiple nonhierarchical and hierarchical log-linear parameters, to describe origin-destination association, with substantively similar results (see Tables S1 and S2 in the Online Supplement on the *ASR* Web site: http://www2.asanet.org/journals/asr/2009/toc070.html).

[8] There are 15 to 22 cells that contain sampling zeros for men, and 22 to 36 cells for women, in the analyses of grouped data (the number of sampling zeros and the total number of cells varies depending on whether and how homemaker mothers are included; the total number of cells in the mobility tables ranges from 216 to 324). There are no zero margins. To detect potential problems due to sparseness, I examined the standard errors of the log-linear parameters; none are unusually large. I did not add a constant (e.g., .5) to the cell counts.

Table 1. Equations: RC Association Models with Various Measures of Class Origin

One-Parent Measures		
1	Father-Only [a]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{ij} + \Phi u_i \nu_j$
2	Mother-Only [b]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_2 D_{hj} + \Phi u_h \nu_j$
3	Higher Class Dominance [c, e]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{hij} + \Phi u_{hi} \nu_j$
4	Lower Class Dominance [d, e]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_2 D_{hij} + \Phi u_{hi} \nu_j$
Joint-Parent Measures		
5	Mother + Father [a, b, f]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{ij} + \delta_2 D_{hj} + \Phi u_{hi} \nu_j$
6	Higher Class + Lower Class [c, d, e, g]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{hij} + \delta_2 D_{hij} + \Phi u_{hi} \nu_j$
7	Equal Mother + Father [a, b, h]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{ij} + \delta_2 D_{hj} + \Phi u_{hi} \nu_j$
8	Equal Higher + Lower Class [c, d, i]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{hij} + \delta_2 D_{hij} + \Phi u_{hi} \nu_j$
Joint-Parent Measures with Interactions		
9	Full Interaction [a, b, e]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{ij} + \delta_2 D_{hj} + \Phi u_{hi} \nu_j$
10	Class Interaction [a, b, e]	$\text{Log } F_{hij} = \lambda_0 + \lambda_h^M + \lambda_i^F + \lambda_{hi}^{MF} + \lambda_j^D + \delta_1 D_{ij} + \delta_2 D_{hj} + \Phi u_{hi} \nu_j$
Extensions of Selected Equations to Control for Cohort and Age		
11	Father-Only [a, j]	$\text{Log } F_{ghij} = \lambda_0 + \lambda_g^C + \lambda_h^M + \lambda_i^F + \lambda_j^D + \lambda_{gh}^{CM} + \lambda_{gi}^{CF} + \lambda_{hi}^{MF} + \lambda_{ghi}^{CMF} + \lambda_{gj}^{CD} + \delta_{1c} D_{ij} + \nu_j (\Phi_0 u_i + \Phi_{1(cohort)} u_i + \Phi_{2(age)} u_i + \Phi_{3(age^2)} u_i + B_{1(age)} + B_{2(age^2)})$
12	Mother + Father [a, b, f, j]	$\text{Log } F_{ghij} = \lambda_0 + \lambda_g^C + \lambda_h^M + \lambda_i^F + \lambda_j^D + \lambda_{gh}^{CM} + \lambda_{gi}^{CF} + \lambda_{hi}^{MF} + \lambda_{ghi}^{CMF} + \lambda_{gj}^{CD} + \delta_{1c} D_{ij} + \delta_{2c} D_{hj} + \delta_{3c} D_{ij} + \nu_j (\Phi_0 u_{hi} + \Phi_{1(cohort)} u_{hi} + \Phi_{2(age)} u_{hi} + \Phi_{3(age^2)} u_{hi} + B_{1(age)} + B_{2(age^2)})$

Notes: h indexes mother's class (M), i indexes father's class (F), j indexes destination class (D), and g indexes birth cohort (C). For identification, λ parameters sum to zero in all models; in all models as applicable, origin and destination scores are identified using the constraints $\Sigma_i u_i = \Sigma_h u_h = \Sigma_{hi} u_{hi} = \Sigma_j \nu_j = 0$ and $\Sigma_i u_i^2 = \Sigma_h u_h^2 = \Sigma_{hi} u_{hi}2 = \Sigma_j \nu_j^2 = 1$.

[a] where $\delta_1 D_{ij} = 1$ if i = j, 0 otherwise.

[b] where $\delta_2 D_{hj} = 1$ if h = j, 0 otherwise.

[c] where $\delta_1 D_{hij} = 1$ if the higher of h or i = j, 0 otherwise.

[d] where $\delta_2 D_{hij} = 1$ if the lower of h or i = j, 0 otherwise.

[e] where origin scores u_{hi} are constrained as shown in the Online Supplement, Table S3.

[f] where $u_{hi} = u_i + u_h$.

[g] where $u_{hi} = u_{hi}$ as defined in equation 3 + u_{hi} as defined in equation 4.

[h] where u_{hi} = mean u_i, u_h from equation 5.

[i] where u_{hi} = mean u_{hi} as defined in equation 3, u_{hi} as defined in equation 4.

[j] where $\delta_{1c} D_{ij}$ and $\delta_{2c} D_{hj}$ are cohort-specific $\delta_1 D_{ij}$ and $\delta_2 D_{hj}$, $\delta_{3c} D_{ij}$ is single-earner family, cohort-specific $\delta_1 D_{ij}$, Φ_0 is the baseline association between u_i or u_{hi} and ν_j, Φ_1 gives the impact of birth cohort on the association, $\Phi_2 + \Phi_3$ give the impact of age on the association, and $B_1 + B_2$ give the association between age and class destination.

only one parent's class position (these include conventional father-only and dominance measures of class origins), joint measures of class origin determined by both parents' individual class positions, and joint measures of class origin determined by both parents' individual class positions that also include interaction effects between parent gender and parent class position. Table 1 presents equations for RC association models that fit the partial association between class destination and various measures of class origin, net of dummy variables for "diagonal" immobility effects. These immobility parameters capture respondents' tendency to cluster along the diagonal cells of the mobility table, where origin = destination, over and above what RC models would otherwise predict.[9]

The one-parent measures of class origin are the conventional *father-only* model, a *mother-only* model, a *higher class dominance* model in which the higher class position solely defines the family class position regardless of parent gender (Erikson 1984), and a *lower class dominance* model, in which the lower class position solely defines the family class.[10] The higher and lower class dominance models constrain the origin scores *u* such that the scores are based on the parent with the higher or lower class position (Table S3 in the Online Supplement shows the design matrices for these constraints).

The joint-parent measures of class origins include a *mother + father* model and a *lower + higher class* model (the latter differentiates parents by relative individual class position rather than by gender [Korupp et. al 2002]). These models combine the mother-only and father-only models and the higher and lower class dominance models described above. In addition, two related models, the *equal mother + father* and *equal lower + higher class* models are constrained versions of the former models; they specify that the effects of the father's and mother's class or the higher and lower class are held equal in determining the joint family position.

Finally, two additional models jointly measure class origin with respect to both parents' individual class positions and also include parent interaction effects. The first is a *full interaction* model that allows each combination of parent class and gender to result in a unique class origin category. The second is the *class interaction* model, which is a constrained version of the full interaction model—it includes interaction effects between pairs of parent class positions, but not between parent class position and parent gender. In other words, in the class interaction model the origin score for a family with a professional mother and a self-employed father is held equal to the score for a family with a professional father and a self-employed mother (see Table S3 in the Online Supplement for the design matrix for the constraints). By contrast, in the full interaction model, the scores for these two combinations of parent class positions are free to differ.

Analysis Steps

ADJUDICATING BETWEEN CLASS ORIGIN MEASURES. The first component of the analysis adjudicates between all the measures of class origin described above, using the RC association models shown in Table 1, for men and women who were raised in dual-earner families (i.e., respondents who reported both a mother's and a father's occupation). I begin by assessing whether models that use the conventional father-only class origin measure adequately fit the mother's class by father's class by class destination mobility table. If conventional measures do not provide an adequate fit, the next question is whether joint-parent measures provide any improvement in overall fit, and if so, which joint-parent measure is preferable (Breen [2005:47–49] suggests this type of empirical strategy). Finally, I evaluate whether the use of different origins measures affects substantive analytical results—most importantly, the estimated degree of social fluidity as indexed by the association between class destinations and each measure of class origins.[11]

CAN HOMEMAKER MOTHERS BE INCORPORATED INTO JOINT-PARENT MEASURES OF CLASS ORIGIN? In the second step of the analysis, I evaluate whether including respondents with homemaker mothers, about one-third of all respondents, changes the results. I examine whether the meas-

[9] The inclusion of diagonal immobility parameters is standard when RC models are used to analyze mobility tables (e.g., Gerber and Hout 2004; Goodman and Clogg 1992), but their inclusion affects the interpretation of the association parameters, which no longer index the total origin-destination association. I estimated all models without immobility parameters with the same results in regard to relative model fit; however, none of the models that lack controls for father–son diagonal immobility fit the mobility data well among men.

[10] Ideally, the dominance models should also take into account which parent has the more enduring attachment to the labor force. While doing so might improve the performance of these measures, it is not possible in this analysis.

[11] I standardized the association parameters so that they are comparable across models with different numbers of origin categories (Clogg and Shihadeh 1994:51).

ures of class origin that were preferred for respondents raised in dual-earner families remain preferred for a larger population that includes respondents raised in single-earner families (i.e., with homemaker mothers). The equations are the same as those used in the prior analysis and shown in Table 1, but with a modification to the immobility parameters—the father-respondent immobility parameter, δ1, is permitted to vary in strength for respondents raised in dual-earner versus single-earner families.

I evaluate three alternative ways of incorporating respondents with homemaker mothers. First, I employ a hybrid of the father-only model and the mother + father model. In this hybrid model, the joint mother + father approach applies only to respondents with employed mothers, while the father's class alone defines class origin for respondents with homemaker mothers. In other words, a mother's position enters the model only for mothers with an occupation; the effect of homemaker mothers is set equal to zero. Second, I fit a model that specifies that homemaker mothers make up a uniform class category that may shape class origin. Third, I fit a model that specifies that homemaker mothers may influence class origin differently depending on their class resources, as indexed by their education levels.

CHANGE IN SOCIAL FLUIDITY BETWEEN SUCCESSIVE BIRTH COHORTS. In the final component of the analysis, I evaluate the importance of using adequate class origin measures for research conclusions about trends in social fluidity. I use a cohort perspective to analyze change over time, comparing social fluidity levels for men and women born in 1945 to 1954, 1955 to 1964, and 1964 to 1979.[12] To control for differences in the age distributions of the survey respondents between cohorts, I extend the RC models used previously by including SOR parameters for age and age-squared. The equations for these extensions of the key models used in the first two steps of the analysis are shown in Table 1 and specify that, net of cohort-specific immobility parameters and cohort differences in the distribution of class origins and destinations (e.g., cohort change in structural mobility), both class origin and respondent age may affect class destination. Furthermore, the strength of the association between class origin and destination may vary by birth cohort and age.

The cohort-specific analyses include respondents with homemaker mothers and use methods for incorporating homemaker mothers that the second step of the analysis demonstrates work best for each gender: for men, homemaker mothers make up one class category; for women, homemaker mothers are differentiated by education level. Additionally, to examine cohort change in social fluidity for a more comprehensive population, the final models in this section also include respondents raised in single-parent families. Given the problem of missing data for noncustodial parent occupation, one of the more straightforward ways to include noncustodial parents in class origin measures is to add "class" categories representing noncustodial mothers and fathers. Instead of six individual class positions for fathers, for example, there are now seven—one of which captures noncustodial fathers (whose occupational class positions are not known). The model therefore estimates an origin score representing the impact of having a noncustodial parent alongside scores for having parents of particular class positions; this accounts for the average impact of noncustodial parents' positions.

RESULTS

Evaluating Conventional versus Joint-Parent Class Origin Measures

In prior debates about joint versus conventional measurement of class, proponents of the conventional approach argued that empirical support for the joint approach to measuring family class must go beyond demonstrating a significant net impact of wives' individual class on the family class position: making an empirically compelling case for the joint view requires demonstrating that conventional measurement is empirically inadequate and that joint measurement can change the substance of research

[12] I exclude respondents from survey years prior to 1994, when the GSS began to collect data on mothers' occupations. This restriction prevents the inclusion of respondents born during the first half of the twentieth century. I define the youngest cohort more broadly than the others to obtain a comparable sample size.

Table 2. Fit Statistics and Parameters for RC Association Models with Various One-Parent and Joint-Parent Class Origin Measures

	Model Description	X^2	L^2	*df*	BIC	*p* vs. 1	*p* vs. 9	Model *p*	Φ	δ_1	δ_2
Men											
1	Full Interaction	159.5	166.2	134	–891.3		0	.07	.26	.31	.03
2	Class Interaction	165.3	172.3	149	–1003.6	.98	0	.17	.23	.31	.03
3	Mother + Father	177.8	186.0	159	–1068.9	.76	0	.15	.21	.31	.03
4	Higher Class + Lower Class	189.2	193.4	159	–1061.5	.35	0	.05	.22	.17	.22
5	Equal Mother + Father	178.2	186.1	164	–1108.2	.92	0	.21	.21	.30	.03
6	Equal Higher + Lower Class	193.7	198.9	164	–1095.4	.34	0	.06	.20	.18	.21
7	Higher Class Dominance	265.2	261.0	165	–1041.2	0		0	.22	.18	
8	Lower Class Dominance	243.9	242.8	165	–1059.4	0		0	.21	.25	
9	Father-Only	228.0	231.9	165	–1070.3	0		0	.18	.30	
10	Mother-Only	364.9	354.0	165	–948.2	0		0	.23	.05	
11	Quasi-independence	309.6	319.4	174	–1053.9	0		0		.60	
12	Independence	498.9	493.5	175	–887.6	0		0			
Women											
1	Full Interaction	134.9	145.5	134	–919.2		0	.46	.29	.02	.07
2	Class Interaction	146.4	158.6	149	–1025.4	.60	0	.54	.25	.02	.10
3	Mother + Father	171.6	183.6	159	–1079.8	.05	0	.23	.20	.03	.08
4	Higher Class + Lower Class	167.1	176.7	159	–1086.7	.18	0	.31	.21	.02	.15
5	Equal Mother + Father	171.6	183.7	164	–1119.5	.15	0	.33	.20	.03	.08
6	Equal Higher + Lower Class	172.1	184.7	164	–1118.4	.12	0	.32	.18	.03	.21
7	Higher Class Dominance	212.3	217.5	165	–1093.6	0		.01	.18	–.01	
8	Lower Class Dominance	186.5	196.4	165	–1114.7	.01		.12	.19	.21	
9	Father-Only	220.3	225.9	165	–1085.1	0		0	.19	.05	
10	Mother-Only	210.8	223.3	165	–1087.8	0		.01	.20	.03	
11	Quasi-independence	283.4	293.5	174	–1089.1	0		0		.24	
12	Independence	302.5	310.6	175	–1079.9	0		0			

Notes: N = 2,676 (men); N = 2,824 (women). Φ is the association parameter and δ_1 and δ_2 are model-specific immobility parameters defined in Table 1.

findings (see Sorensen 1994). The results of the first step of the analysis demonstrate that both of these standards are met in the case of family class origin.

Table 2 shows the results of mobility models employing the various one-parent and joint-parent class origin measures.[13] The model significance statistics (Model *p*) illustrate which models best summarize the actual patterns observed in the data (if Model *p* is significant at $p < .05$, the models' predictions differ significantly from the observed mobility data and model fit is not considered adequate). None of the models employing one-parent class origin measures (Models 7 through 10 in Table 2) accurately summarize observed mobility patterns, apart from the lower-class dominance model among women. By contrast, all of the models using joint-parent measures (Models 1 through 6 in Table 2) do accurately account for the patterns observed in the data. These results indicate that each parent's class resources shape family class origin, and using one parent's class position as a proxy for the family-level class position is not an empirically adequate approach.

The next question, then, is which of the multiple joint-parent models, all of which adequately summarize observed mobility, are the

[13] Among the one-parent measures, the higher-class dominance model, suggested as an improvement over the conventional father-only model, has a poorer overall fit than the father-only model among men and, notably, a poorer fit than the lower-class dominance model among both men and women. The father-only model is strongly preferred over the mother-only model for men, but the two are equivalent for women.

best? While L^2 comparisons show that all of the models that employ joint-parent class origin measures (Models 1 through 6 in Table 2) provide a significantly better account of mobility than does the father-only model, the full and class interaction models are needlessly complex compared with more parsimonious joint-parent models, such as the mother + father models (the BIC criterion, which emphasizes model parsimony, rejects the full and class interaction models compared with the father-only model). Both the higher + lower class and mother + father models remain as appropriate joint-parent measures of class origin; for the remainder of the study, I focus on the mother + father conceptualization because it is more straightforward than the higher + lower class measure.

Not only do conventional origin measures fail to serve as an adequate proxy for family-class position from the perspective of accurately summarizing observed mobility patterns, but the results of this analysis also illustrate how their use could distort the substance of mobility research findings. Generally, mobility studies address questions about the extent of equality of opportunity in a society by examining the estimated extent of social fluidity or class immobility. The δ parameters in Table 2 show that the extent of class immobility estimated by the father-only model approximates that of the mother + father models.[14] Turning to estimated social fluidity, on the other hand, the Φ parameters in Table 2 for the father-only model are smaller than the Φ parameters estimated by the joint-parent models from the same data.[15] This is particularly true for men. The estimated association between class origins and destinations is 15 percent higher among men and 5 percent higher among women given either of the mother + father models, compared with the conventional model.[16]

That the different models estimate somewhat different levels of association given identical data supports the possibility, further evaluated below, that, when both parents' resources determine family class position but it is measured with respect to fathers only, measurement error due to ignoring mothers' resources could depress the estimated intergenerational association in class position, potentially distorting research findings about social fluidity levels. This is of particular concern because the extent of the measurement error and consequent distortion of estimated social fluidity likely varies between comparison groups, such as nations or cohorts.

Can Homemaker Mothers be Incorporated into Joint-Parent Measures of Class Origin?

For simplicity, the first step of the analysis excludes respondents who reported that their mothers worked in the home rather than in the labor force (i.e., 33 percent of otherwise eligible respondents). The results of this second step of the analysis, shown in Table 3, demonstrate that the preferred joint-parent measures from the prior analysis can be extended to a population that includes single-earner families. The joint-parent measures continue to provide a more accurate depiction of observed mobility in this larger population than do conventional class origin measures.

First, it is important to reevaluate the conventional father-only model in the context of the current sample. Despite the fact that one-third of the sample reported that their fathers were the only employed parent, the father-only model (Model A1 in Table 3) does not adequately summarize observed mobility patterns. By comparison, the hybrid of the father-only model and the mother + father model (where the joint mother + father approach applies only to respondents with employed mothers, while the father's class alone defines class origin for respondents with homemaker mothers) provides a significantly better account of mobility, compared with the poorly fitting father-only

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[14] This is not the case when comparing higher class dominance with the higher + lower class model, the latter estimates nearly twice as much class immobility as the former.

[15] LEM does not provide standard errors for the Φ parameters; that the association is significantly higher than 0 is demonstrated by the rejection of the independence model (Model 12 in Table 2).

[16] The father-only model also underestimates the association compared with the mother-only model. Note, however, that the mother-only model does not include the father-respondent immobility parameter, which accounts for the difference.

Table 3. Methods for Including Homemaker Mothers

Model	Description	X^2	L^2	*df*	BIC	*p* vs. 1	Model *p*	Φ	δ_{1_a}	δ_{1_b}	δ_2
Men											
A: Treating the Homemaker Mother Class Category as a Single Category											
1	Father-Only	277.5	282.5	195	–1338.0		0	.19	.32		
2	Employed Mother + Father	226.6	235.0	188	–1327.4	0	.03	.21	.25	.46	.05
3	Equal Employed Mother + Father	227.1	236.3	193	–1367.6	0	.05	.22	.27	.48	.02
4	Mother + Father	226.5	235.0	187	–1319.1	0	.03	.21	.25	.46	.05
5	Equal Mother + Father	227.2	236.4	192	–1359.2	0	.04	.22	.27	.49	.02
B: Differentiating the Homemaker Mother Class Category by Education Level											
1	Father-Only	391.9	404.9	255	–1714.2		0	.19	.32		
2	Mother + Father	294.9	306.7	245	–1729.3	0	.02	.24	.29	.42	.04
3	Equal Mother + Father	296.7	308.0	251	–1777.9	0	.03	.23	.27	.40	.05
Women											
A: Treating the Homemaker Mother Class Category as a Single Category											
1	Father-Only	279.1	284.7	195	–1338.6		0	.20	.05		
2	Employed Mother + Father	232.1	240.8	188	–1324.2	0	.02	.21	–.05	.27	.02
3	Equal Employed Mother + Father	233.1	242.6	193	–1364.0	0	.03	.21	–.04	.30	.01
4	Mother + Father	226.4	234.3	187	–1322.3	0	.03	.21	–.05	.26	.04
5	Equal Mother + Father	227.7	237.2	192	–1361.0	0	.04	.21	–.03	.30	.03
B: Differentiating the Homemaker Mother Class Category by Education Level											
1	Father-Only	361.6	376.4	255	–1746.3		0	.20	.05		
2	Mother + Father	265.4	282.7	245	–1756.8	0	.18	.27	–.02	.17	.03
3	Equal Mother + Father	271.0	287.0	251	–1802.4	0	.18	.25	–.05	.15	.04

Notes: N = 4,066 (men); N = 4,123 (women). Φ is the association parameter. In Model 1, δ_{1a} represents father-respondent immobility. In Models 2 to 5, δ_{1a} represents father-respondent immobility in dual-earner families and δ_{1b} represents father-respondent immobility in single-earner families (i.e., families with homemaker mothers). δ_2 represents mother-respondent immobility in all models. See Table S4 in the Online Supplement for class origin and destination scores for selected models.

model. However, while it comes close, this hybrid mother + father model (Model A2 in Table 3) still does not quite fit the mobility table (although among men, the more parsimonious hybrid equal mother + father model [Model A3] does fit).

Considering that the mother + father model adequately depicts mobility when the analysis is restricted to respondents raised in dual-earner families, this change for the worse in the fit of the model to the data suggests that homemaker mothers may indeed shape family class origin. To further investigate this idea, Models A4 and A5 in Table 3 include homemaker as a possible class category for mothers. Among men, including the single broad class of homemaker mothers adds little to the picture of class origin—the effect of the homemaker mother class category is very close to zero, even though it was not constrained to zero as in Models A2 and A3. Among women, however, the models that include a class of homemaker mothers (Models A4 and A5) are marginally preferred over models that set the effects of homemaker mothers to zero (Models A2 and A3).

One interpretation of the finding that models ignoring homemaker mothers fit the mobility table equally well among men, and almost as well among women, as models that include a class category for homemaker mothers, is that the joint view of family class may apply only to dual-earner families. Another possible interpretation, however, is that the measurement error inherent in the single broad category of homemakers with diverse class resources could depress the observable effects of such resources toward zero. The second interpretation is consistent with the results of the models in terms of overall fit—none of the models summarize the mobility data as well as the same models in the previous analysis, which was restricted to respondents raised in dual-earner families. This suggests that homemaker mothers shape class origin, but in a fashion that is not well measured by the inclusion of a single homemaker class category. A second series of models (B, in Table 3) therefore differentiates the homemaker category by education level, as a proxy for homemaker mothers' class resources. Among men, differentiating the homemaker mother class category with respect to education still does not adequately describe mobility patterns. Among women, on the other hand, the joint-parent models in which homemaker mothers are differentiated by education level (Models B2 and B3 in Table 3) fit the mobility table well overall. Mothers' nonemployment based class resources may be more salient for women than for men.[17]

For both men and women, the results confirm that the substantive findings of the previous analysis continue to apply when respondents with homemaker mothers are included in the mobility data set. Joint-parent models provide a substantially better summary of observed mobility than does the conventional model, although the diversity within the homemaker category may be problematic with respect to model fit. The findings also reconfirm those in the previous analysis that the father-only model produces a lower estimate of intergenerational association in class position. Finally, in this step of the analysis, the inclusion of mothers' characteristics has implications for the estimated immobility between father's class and class destination. Both father–son and father–daughter class immobility is stronger in single-earner families than in dual-earner families. This weaker father–respondent immobility when mothers are employed outside the home further illustrates how mothers' employment matters for intergenerational mobility.

Change in Social Fluidity between Successive Birth Cohorts

The results thus far demonstrate that class origin measures that jointly capture both parents' class positions provide the most accurate picture of intergenerational mobility. Furthermore, defining class origin with reference to the father's position alone understates intergenerational inequality. Consequently, conventional class origin measures could very well produce misleading comparisons of social fluidity levels between groups. This final section of the analysis evaluates the research consequences of class origin measurement choices for detecting whether social fluidity levels have increased, decreased, or stayed constant across birth cohorts.

Model 1 in Table 4, the conventional father-only model, indicates that no significant change has occurred in social fluidity levels between the cohorts for either gender (similarly, diagonal immobility parameters show no significant change between cohorts). This provides a misleading picture of social fluidity trends, however, for when joint-parent class origin measures are employed in Model 2, conclusions about change in social fluidity are dramatically different. The mother + father model shows that origin-destination association has strengthened over the cohorts, and therefore social fluidity has declined, particularly for the most recent cohort.[18] As Figure 2 illustrates, the conventional father-only class origin measure (Model 1) underestimates the origin-destination association for the younger cohorts, masking a reduction in fluidity between cohorts that is revealed by joint-parent class origin measures (Model 2).[19] The reduction in fluidity experienced by the most recent cohort compared with the earliest cohort is statistically significant for men but not for women.[20]

[17] I tested origins measures that differentiated homemaker mothers with respect to their husbands' class (in other words, I tested an interaction between the homemaker category and husband's class). The results are similar to differentiating homemaker mothers by their education.

[18] I employ the mother + father model, rather than the more parsimonious equal mother + father model, in the cohort analysis to capture how the relative importance of parents' class characteristics has changed between cohorts.

[19] Similarly, the conventional model is unable to detect a change in cohort fluidity, whereas joint-parent origin models detect such change, when cohort change is assessed without controlling for age.

[20] Note that the models for men and women differ in that homemaker mothers make up one broad class category in men's models but are differentiated by education in women's models, in accordance with results from the previous analysis section. The degree of measured association is therefore not direct-

Table 4. Cohort Change in Origin-Destination Association (Social Fluidity)

Model	1: Father-Only Model		2: Mother + Father Model		3: Model 2 + Single Parents	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Men						
Baseline Association (Φ)	.21*	.04	.21*	.04	.23*	.04
Cohort 2(1955 to 1964)[a]	0	.06	.04	.06	.04	.05
Cohort 3(1965 to 1979)[a]	.06	.07	.16*	.08	.12†	.07
Age [a]	.02	.02	.04*	.02	.04*	.02
Age^2 [a]	0	0	.00*	0	.00*	0
Baseline Father–Son Immobility	.34*	.08				
Cohort 2(1955 to 1964)[a]	0	.11				
Cohort 3(1965 to 1979)[a]	.17	.11				
Baseline Father–Son Immobility, Dual-Earner Families			.36*	.10	.37*	.10
Cohort 2(1955 to 1964)[a]			–.02	.13	–.03	.13
Cohort 3(1965 to 1979)[a]			.02	.13	–.01	.13
Baseline Father–Son Immobility, Single-Earner Families			.36*	.12	.37*	.11
Cohort 2(1955 to 1964)[a]			.02	.16	–.01	.16
Cohort 3(1965 to 1979)[a]			.48*	.17	.45*	.17
Baseline Mother–Son Immobility			–.01	.12	–.01	.11
Cohort 2(1955 to 1964)[a]			.01	.15	0	.14
Cohort 3(1965 to 1979)[a]			.18	.15	.25†	.13
Age	–.14*	.05	–.18*	.05	–.11*	.04
Age^2	.00*	0	.00*	0	.00*	0
N	3,580		3,580		4,406	
Log Likelihood	–6000.9		–5963.1		–7317.1	
LR Chi Square	1183.9 (25)		1259.5 (32)		1493.3 (32)	
BIC'	–979.4		–997.6		–1224.8	
Women						
Baseline Association (Φ)	.22*	.04	.26*	.04	.30*	.04
Cohort 2(1955 to 1964)[a]	–.01	.06	.06	.06	.03	.06
Cohort 3(1965 to 1979)[a]	.02	.08	.14	.09	.06	.08
Age[a]	.03	.02	.02	.02	0	.02
Age^2 [a]	0	0	0	0	0	0
Baseline Father–Daughter Immobility	.13	.09				
Cohort 2(1955 to 1964)[a]	.04	.13				
Cohort 3(1965 to 1979)[a]	–.08	.14				
Baseline Father–Daughter Immobility, Dual-Earner Families			–.10	.12	–.06	.11
Cohort 2(1955 to 1964)[a]			.22	.15	.13	.15
Cohort 3(1965 to 1979)[a]			.03	.16	0	.15
Baseline Father–Daughter Immobility, Single-Earner Families			.27*	.14	.26†	.14
Cohort 2(1955 to 1964)[a]			–.09	.20	–.10	.19
Cohort 3(1965 to 1979)[a]			–.02	.23	–.02	.23
Baseline Mother–Daughter Immobility			.30*	.09	.22*	.09
Cohort 2(1955 to 1964)[a]			–.25*	.13	–.13	.11
Cohort 3(1965 to 1979)[a]			–.29*	.13	–.17	.11
Age	–.02	.05	0	.06	–.01	.05
Age^2	0	0	0	0	0	0
N	3,620		3,620		4,570	
Log Likelihood	–5400.4		–5329.8		–6701.8	
LR Chi Square	2072.1 (25)		2213.3 (32)		2614.1 (32)	
BIC'	–1867.2		–1951.1		–2344.4	

Notes: See Table S4 in the Online Supplement for the class origin and destination scores for these models. Age was re-centered around 40. Higher class destination scores are negative (see Table S4 in the Online Supplement), therefore negative coefficients for age indicate a positive relationship between age and higher class destination.
[a] These variables denote interaction terms (e.g., to obtain the estimated association for respondents in Cohort 3, add the coefficients for the baseline association [Φ] and Cohort 3 [1965 to 1979]).
† $p < .10$; * $p < .05$.

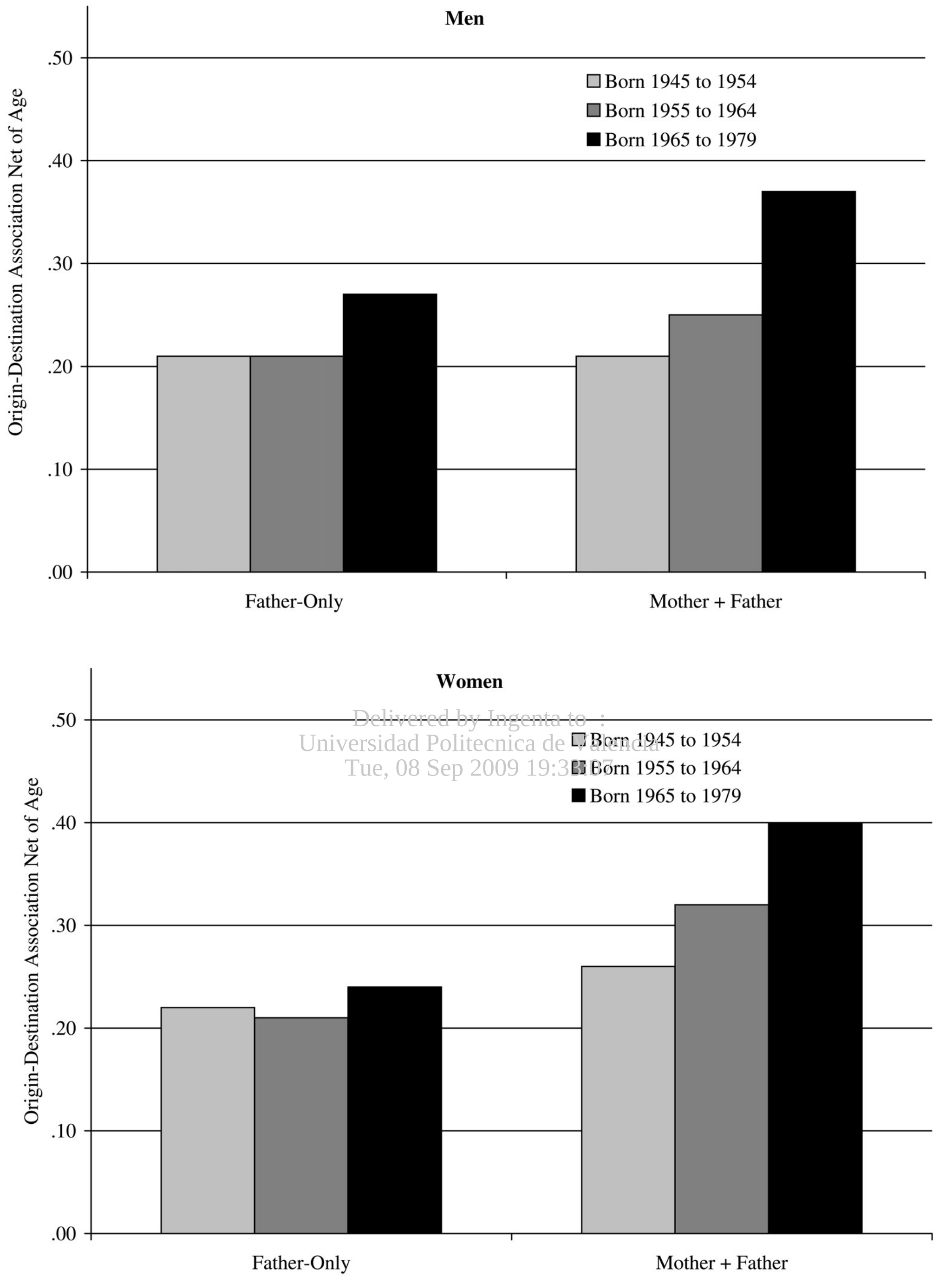

Figure 2. Association Parameter Estimates by Cohort, Net of Age Effects

ly comparable between men and women. When cohort change is assessed for women without differentiating homemaker mothers by education, the pattern of declining social fluidity between cohorts is less apparent (and it is not statistically significant in either case). Prior research did not find evidence of differences between men and women in social fluidity trends over time; however, this research employed a period comparison rather than distinguishing between cohorts (Hout 1988).

Additional differences between men and women include the finding that father–son diagonal immobility is significantly higher for men raised in single-earner families in the most recent birth cohort, compared with the earliest cohort. Among women, there is no change in father–daughter immobility between cohorts, but mother–daughter immobility is substantially reduced in the second and third cohort, compared with the first. While there is no apparent effect of age for women, respondent age among men is associated not only with higher class position, but also with stronger origin-destination association when joint-parent origins measures are employed. This means that family class origins and class destinations are more tightly linked among older than younger men within each cohort.[21] This relationship between men's age and greater intergenerational association is not detectable when conventional origins measures are used in Model 1.

Considering the dramatic results of Model 2, which indicate rising inequality, particularly among men, it is important to ask whether the finding is representative of a broader population that includes respondents raised in single-parent families. Some research suggests that respondents raised in non-intact families have higher rates of social mobility than do their peers (Biblarz and Raftery 1999). Therefore, when considering population-level estimates of social fluidity, an increasing proportion of respondents raised in single-parent families in more recent cohorts could counteract the declining social fluidity among those raised in intact families. Model 3 in Table 4 evaluates whether the results of Model 2 are maintained when the population analyzed includes respondents raised in single-parent families. Again, among men the results of Model 3 demonstrate significantly declining social fluidity between cohorts for this more representative population of respondents. Among women, however, all the cohorts appear less fluid after including respondents raised in single-parent families.

DISCUSSION AND CONCLUSIONS

The theoretical argument and empirical analyses presented in this article clearly demonstrate the importance of adequate measurement of family-level class background to social mobility research. Conventional measurement, when applied to family class origin, assumes that, net of the key measured parent's class position, a second parent's class-related resources do not affect the family class position. Yet, as theorized mechanisms of intergenerational persistence in class would suggest, this study shows that parents' class resources jointly determine family class origin. With this in mind—together with the view that class is determined by resources that follow from employment relationships—it may be helpful, in the social mobility research context, to think beyond the idea of classes as necessarily bounded positions that individuals or families occupy. Rather, class origins could be conceived of as sets of family-level economic, cultural, and other class-related resources that shape children's mobility chances and are consequences of employment relationships, occupational conditions, and other class-related experiences of adults in the family.

Empirically, this study demonstrates that joint-parent measures of class origin capture mobility patterns significantly better than do conventional measures of class origin. Although conventional measurement may provide a convenient proxy for family-level class, it is increasingly inadequate and fails to capture significant declines in social class fluidity among men born between 1965 and 1979 (compared with earlier cohorts) in the United States. This decline, as my analyses show, is a function of the growing association between mothers' class and sons' class destinations. This example of change in social fluidity between cohorts illustrates that inadequate measurement of family class origin can distort research conclusions. Moreover, the implications extend well beyond this particular example and apply to comparative mobility research more generally. For example, if mothers' class remains unmeasured and marital sorting by class differs cross-nationally, differences in measurement error could be misinterpreted

[21] Hendrickx and Ganzeboom (1998) found the opposite relationship among men in the Netherlands; more years of work experience led to lower origin-destination association. Controlling for years of work experience rather than age does not change the results of this analysis.

as substantive differences between countries in social fluidity levels. On the flip side, substantive cross-national differences in social fluidity could also be masked.[22]

Given the evidence offered regarding a dramatic decline in social fluidity among recent birth cohorts, future research should examine the social processes that may account for it. For example, women's careers and economic participation became more intangibly valued and recognized during this period, which potentially affected the extent to which children view mothers as career role models. In addition, there are at least two compelling explanations related to trends in increasing inequality between families that could play a role in explaining the recent decline in social fluidity. First, family economic resources became more unequally distributed due to rising income inequality. Second, noneconomic, cultural family class resources also became more unequally distributed due to increasing educational assortative marriage.

With respect to the first of the possibilities noted above, the explanation of rising economic inequality between families, the youngest cohort in this study grew up after the substantial increase in income inequality that occurred in the United States during the 1970s and 1980s. Greater income inequality might have heightened the advantages and disadvantages of class origin, thereby strengthening the linkage between origins and class destinations. Indeed, a similar pattern occurs in studies of intergenerational income mobility: when income inequality in the children's generation exceeds that of the parents' generation, intergenerational income mobility decreases (Solon 2001). For the rise in income inequality to be a plausible explanation for declining social class (as opposed to income) mobility depends, however, on whether income inequality increased more between, as opposed to within, class categories and also on the extent of assortative marriage by income in dual-earner families.

Like economic resources, a key cultural class resource—parental education—also became more unequally distributed among families between the successive birth cohorts analyzed in this study. Educational assortative marriage has increased since the 1960s, following a period of decline. While this trend was initially due to an increasing propensity among those with a college degree to marry one another, by the 1970s there was also a strong decline in the extent to which individuals with low levels of education married upward with respect to education (Schwartz and Mare 2005). Individuals born in the most recent birth cohort analyzed in this study were thus more likely than their predecessors to have parents with similar levels of education. The distribution of parental education between families therefore became more unequal. To the extent that education is related to occupational class position, this pattern might correspond with an increasing social divide between family class origin categories.

While this article provides initial steps toward modernizing family-based stratification research, I do not measure class destinations at the family-level, and the current line of inquiry will be logically incomplete until the same scrutiny I apply to class origin is extended to class destination. Future research might evaluate, for instance, how inequalities of family class origins may or may not be compounded given family-level, as opposed to individual, measurement of class destinations.

Finally, it is worth noting that limited data presents a substantial challenge for updating the practice of intergenerational social mobility research and stratification research more broadly. This article, for instance, illustrates the importance of measuring both parents' class resources to adequately define class origin, yet most surveys do not collect information on noncustodial parent occupation and many still do not collect mothers' occupations. For the time being, stratification and mobility research should, of course, proceed despite data limitations. This article demonstrates that the best practice is to measure class origin as jointly determined by both parents' class characteristics, but this is often not possible with current data sources. Researchers can move forward despite limited data by, on the one hand, measuring class origin as comprehensively as pos-

[22] In addition, the argument presented here in favor of family-level measurement of childhood social position is not necessarily limited in applicability to the particular class schema employed. It could also apply to other class categorizations or methods of differentiating between social positions.

sible, when possible, and on the other hand, by carefully considering in the interpretation of research findings the potential for omitted variable bias given that important aspects of children's class origins may remain unmeasured.

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Table A1. Class Distributions of Fathers, Mothers, and Respondents (Weighted Percent)

	Fathers	Mothers	Men	Women
Respondents Raised in Two-Parent Families (N = 8,189)				
I	22.63	5.04	25.29	19.39
II	12.01	14.40	18.71	32.31
IIIab	4.79	24.56	8.28	30.65
IVab	10.96	5.03	7.82	6.09
V/VI	24.89	4.27	19.26	2.99
VIIab	24.72	13.34	20.65	8.58
Home		33.35		
Respondents Born 1945 to 1954 (N = 2,522)				
I	18.34	3.58	30.13	19.76
II	9.05	11.20	18.86	30.83
IIIab	4.75	22.33	5.34	29.46
IVab	9.70	5.25	9.11	7.47
V/VI	22.76	4.28	15.68	3.21
VIIab	22.75	14.92	20.87	9.27
Home		37.05		
Noncustodial	12.65	1.39		
Respondents Born 1955 to 1964 (N = 3,364)				
I	19.81	5.15	24.73	18.51
II	10.40	12.98	16.17	29.99
IIIab	4.02	27.34	8.51	31.46
IVab	9.25	5.36	6.85	5.68
V/VI	20.10	4.87	21.36	3.93
VIIab	20.13	13.69	22.38	10.43
Home		28.01		
Noncustodial	16.29	2.59		
Respondents Born 1965 to 1979 (N = 3,090)				
I	18.21	7.38	17.94	18.26
II	10.69	19.24	20.91	32.56
IIIab	3.31	27.76	10.44	31.72
IVab	8.03	5.21	5.79	4.53
V/VI	19.92	3.93	23.76	2.95
VIIab	19.34	13.27	21.16	9.98
Home		20.58		
Noncustodial	20.49	2.64		

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