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Explaining Change in Social Fluidity: Educational Equalization and Educational Expansion in Twentieth-Century Sweden[1]

Richard Breen
Yale University

Jan O. Jonsson
Stockholm University

The authors analyze social fluidity among Swedish men and women using a series of 24 annual surveys, 1976–99 ($N = 63{,}280$). A theoretical model suggests that changes in fluidity are normally driven by cohort rather than period effects. The results support this argument: changes in fluidity between the mid-1970s and late 1990s were due to the successive replacement of older and less fluid, by younger and more fluid, cohorts. Cohorts differed in their fluidity because the effect of class origins on educational attainment declined (an equalization effect) and because greater shares of each cohort had higher levels of educational attainment, which placed them in labor markets that operate more meritocratically (a compositional effect). The article discusses the relevance of these results for other countries and for policy.

INTRODUCTION

In studies of social mobility, Sweden has long been recognized as holding a distinctive place: class origins (the social class in which a person is

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brought up) appear to have a smaller influence on class destinations (the class which the person comes to occupy as an adult) than in most other countries (Erikson and Goldthorpe 1992; Breen 2004), and there has been a fairly lengthy period during which the impact of origins on destinations has steadily weakened. In this article we reaffirm this trend, but we go beyond most research in social mobility by developing a theoretical model of how class origins are linked to class destinations and in providing an explanation, which we test empirically, of the trends that we find. It is no secret to say that contemporary research in social mobility is, for the most part, technically sophisticated but theoretically weak. In this article, we work to strike a better balance by providing a firm theoretical footing for the empirical study of social mobility.

The term "social fluidity" is often used to refer to the degree to which class destinations depend on class origins: the weaker the statistical association between origins and destinations, the greater social fluidity is said to be.[2] Accordingly, social fluidity is often interpreted as an index of equality in the chances of access to more or less advantageous social positions between people coming from different social origins, and studies of whether societies are moving toward greater "openness" use social fluidity as a key indicator. Whether countries differ substantively in their degree of social fluidity and whether fluidity changes over time are questions that have been much debated in the stratification literature. Research in the 1980s and early 1990s (e.g., Grusky and Hauser 1984; Erikson and Goldthorpe 1992) found support for the "FJH hypothesis," or variants of it, which claims that social fluidity is similar in all industrial societies "with a market economy and a nuclear family system" (Featherman, Jones, and Hauser 1975, p. 340). Some of these studies also argued for "a high degree of temporal stability" (Erikson and Goldthorpe 1992, p. 367; also Goldthorpe 2000, chap. 11) in social fluidity. But particularly in more recent research, greater emphasis has been placed on change and variation. Ganzeboom, Luijkx, and Treiman (1989, p. 47) claimed that "there are substantial cross-national and cross-temporal differences in the extent of mobility," and Breen (2004; also Breen and Luijkx 2004*a*) documents significant variation in social fluidity among the countries of Europe and a fairly widespread temporal trend toward greater fluidity in the closing decades of the 20th century (see also the review in our earlier work [Breen and Jonsson 2005]).[3]

Our findings in this article support the latter position insofar as we report change in Swedish social fluidity over the course of the 20th century, but we also provide a theoretical framework for understanding how social fluidity may change, and we use this framework to account for the change that we observe. Hitherto, the task of explaining temporal change within a country has often been an ad hoc affair and, although some studies (esp. Hout 1988) have provided valuable insights into the mechanisms that might underlie change, it has not been possible wholly to account for temporal trends by the addition of variables which represent plausible processes by which fluidity might change. But this is what we do in the present study: we show that variables capturing two simple processes can indeed account for trends in social fluidity in 20th-century Sweden.

Previous research shows that the association between class origins and destinations in Sweden was relatively stable during the first decades of the 20th century (Carlsson 1958) but weakened during the postwar period and up to the beginning of the 1980s (Erikson 1983, 1987).[4] This trend toward increased openness continued to 1991, particularly for women (Jonsson and Mills 1993; Jonsson and Erikson 1997). In this article we first examine changes in social fluidity among Swedish men and women by comparing successive birth cohorts. This allows us to cover almost the entire 20th century, because the oldest cohorts in our data were born around 1912 and the youngest in the early 1970s. On the other hand, and in contrast to this cohort perspective, we can also compare social fluidity across the different surveys in which our data were collected. These were carried out between 1976 and 1999, and so this period perspective allows us to focus on changes that took place in the final quarter of the 20th century. But the period and cohort perspectives are different ways of examining the same data, and so the second goal of the article is to relate the two. As we explain below, we believe that there are good reasons for supposing that changes in fluidity are normally and mainly—though not exclusively—driven by cohort-related, rather than period-related, factors. If this is true we should expect that most, if not all, of any period change that we detect will prove to be a consequence of changes occurring between cohorts.

The third, and major, goal of the article is to account for the pattern of change that we observe. We look at the role of education, given the widespread recognition that it is one of the major channels through which intergenerational class reproduction occurs (Ishida, Müller, and Ridge 1995), and we identify two processes through which education might cause

[2] In the log-linear modeling context in which this article is situated, the association between origins and destinations is measured using odds ratios.

[3] And some former proponents of the FJH hypothesis now reject it (see Hauser 1995, pp. 176–77).

[4] Ganzeboom et al. (1989), using five Swedish data sets collected between 1950 and 1983, claim that Sweden shared in the general worldwide trend to increasing fluidity that they identify. Wong (1994), in his reanalysis of their data, found that this trend was evident only for Hungary and Sweden.

social fluidity to change. The first, which we call "equalization," is a decline in the association between class origins and educational attainment: in other words, class origins come to exercise a weaker effect on educational attainment. In several European countries, including Sweden (as we document below), reforms of the educational system have been undertaken with exactly this goal. Equalization affects social fluidity because, for a given association between education and class destinations, a lesser impact of class origins on educational attainment will weaken the overall association between origins and destinations. The second process we call "compositional": if there is an association between origins, education, and destinations such that the origin-destination association is weaker at higher levels of education, and if educational expansion places increasing shares of each cohort in those educational levels where the association is weakest, then this compositional change can be expected to lead to an overall reduction in the gross association between origins and destinations.[5] A three-way interaction between class origins, educational qualifications, and class destinations may be present when, for example, higher qualifications are a powerful signal for employers that leaves little leeway for social network effects, or when the job markets in which degreeholders operate are particularly meritocratic. A weaker origin-destination association at higher levels of education is in fact reported from the United States (Hout 1988), France (Vallet 2004), and Sweden (Erikson and Jonsson 1998), making it possible that an expansion of higher education across cohorts in these countries led to increasing fluidity. In our analyses we document the existence of both equalization and compositional effects, and we also seek to quantify the importance of each in accounting for cohort changes in social fluidity.

In the next section of the article we present a theoretical model of the processes that underlie change and stability in social fluidity, and we explain why these processes are more likely to manifest themselves as cohort than as period effects. Successive sections present our data and our analyses, and a final section summarizes our results and draws conclusions.

[5] A stronger result can be shown in the case of linear systems. Let X, Z and Y be continuous measures of, respectively, class origins, educational attainment, and class destinations. We then have a two-equation stochastic system: (i) $Z = a_0 + a_1X + u_1$; and (ii) $Y = b_0 + b_1X + b_2Z + b_3XZ + u_2$. By substituting (i) into (ii) and arranging terms we see that the derivative, $\partial Y/\partial X$ (i.e., the unconditional effect of X on Y), is given by $b_1 + b_2a_1 + b_3a_0 + 2b_3a_1X$, and the derivative of this with respect to a change in the overall level of educational attainment (i.e., a_0) is given by b_3. So if there is no interaction between origins and education ($b_3 = 0$), changes in the mean level of education will not change the gross association between origins and destinations. But if, as in the case under discussion, $b_3 < 0$, such a change will reduce this association (the compositional effect). But the association will always change for a change in the origin-education relationship (captured by a_1, the equalization effect).

A THEORETICAL MODEL OF SOCIAL FLUIDITY

The association between origins and destinations (parents' class and respondent's class) depends on the degree to which factors associated with class position in the parental generation can influence which classes their children come to occupy.[6] Figure 1 provides a simplified depiction of this process. In the parental and filial generations occupation of a particular class position depends on assets (such as educational qualifications) and gives rise to consequences (such as income). In the parental generation the impact of a given asset on class position is labeled α, and in the filial generation, γ, while the "return" to class in terms of consequences is given by λ and by φ in the two generations.[7] The extent to which a member of the filial generation possesses a particular asset depends on the direct transmission of that asset between generations (this effect is labeled τ_1) and on the degree to which the consequences of class position in the parental generation are associated with the securing of more of this asset in the filial generation (this effect is labeled τ_2). The parameters α, λ, γ, τ_1, and τ_2 all play a role in social fluidity because they shape the two fundamental mechanisms that underlie it: (*a*) the association between filial assets and parental class, which we call *transmissibility* (given by $\alpha\tau_1 + \lambda\tau_2$),[8] and (*b*) the association between filial assets and filial class destinations, or *class returns* (γ). Together these determine the association between class origins and destinations as mediated by a particular asset (given by $\gamma[\alpha\tau_1 + \lambda\tau_2]$). Variations in fluidity, over time or between countries, will be largely dependent on institutional arrangements and demographic circumstances that determine the transmissibility of, and the returns to, assets.

The direct transmission of assets (through τ_1) includes the inheritance of property, aspirations, and aptitudes and, perhaps most important, genetic factors. The indirect transmission, via τ_2, captures the idea that occupying a particular class position entails consequences, which themselves influence the accumulation of assets in the next generation. These

[6] This section draws particularly on ideas in Breen (1997) and Bowles and Gintis (2002).

[7] In some cases one might want to assume that one or both of these effects was constant over generations: $\alpha = \gamma$, and/or $\lambda = \varphi$.

[8] Transmissibility, as we define it, is $\alpha\tau_1 + \lambda\tau_2$ rather than simply $\tau_1 + \tau_2$, because we are interested in the intergenerational transmission of assets only insofar as it contributes to the association between parental class and filial assets. From this perspective transmissibility must include the link between parental class and parental assets.

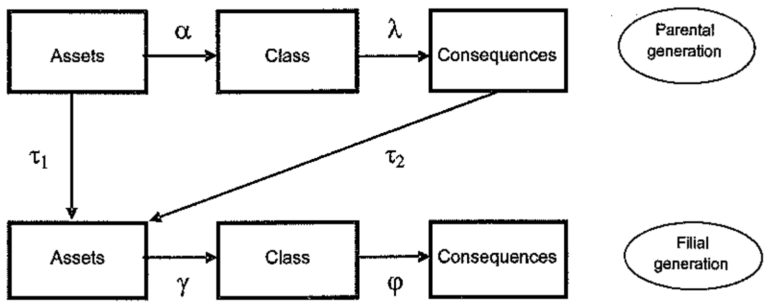

FIG. 1.—Influence of parental position on class location of offspring

consequences may be psychological, as in the case of preferences shaped by class position and transmitted to children, or material, as in the case of the rewards that accrue to particular classes. How far such consequences can influence the assets for mobility possessed by the next generation depends on the degree to which classes generate differential consequences (which is captured in fig. 1 by λ) and the extent to which these can be translated into assets for the attainment of class positions by their children (τ_2). Income is an obvious example. The greater the variation in income between classes and the more families can use their income to purchase advantages for their children, the stronger will the origin-destination association be. Such would be the case where class differences in income are great and public provision of education is poor: here high income can be used to purchase education that then functions as an asset for achieving a privileged class position. But if progressive taxation reduces the inequality between classes in their market income (λ is small) and/or the public educational system is as good as the private system (τ_2 is small), then the association between origins and destinations will be correspondingly weaker. A similar argument can be advanced in respect of other institutions of the welfare state that reduce the extent to which inequalities in rewards in the parental generation translate into a differential distribution of mobility assets among their children.

Change in fluidity can come about through changes in either or both of the two fundamental mechanisms mentioned above: the class returns to assets and their transmissibility. Educational qualifications, which we focus on in this article, may become less dependent on origins due to school reforms, for example, which would suppress transmissibility. Or education may become less important in the labor market, in which case the class returns will decrease. In both of these cases, social fluidity will increase. But in real life several assets are likely to be of importance, and so changes could be hard to predict. In the case of two assets, such as

social networks and educational attainment, we may find that even if the transmissibility of education was decreasing, this could be counteracted by increasing class returns to networks. And if the transmissibility of networks was stronger than that of education, but neither changed over time, social fluidity would in fact decline if the class returns to networks increased relative to the returns to education.

A more complex situation arises if the return to one asset depends on the returns to another asset. In this case, a shift in the distribution of one of these assets can lead to a change in social fluidity. If social networks are consequential for class position among those with low education but not at high levels of education, and if the distribution of education changes so that a larger share of the filial generation has higher education (even though the correlation between education and parental social class remains the same), then the average effect of social networks on class position will nevertheless decline, and, therefore, social fluidity will increase.

While changes in social fluidity, according to our model, follow changes in transmissibility or class returns, they could take effect in two major ways, namely as *cohort effects* or *period effects*. The former works through cohort replacement, the importance of which for social change has been stressed by several sociologists (Mannheim 1952; Ryder 1965). For example, increased social fluidity in the working-age population would come about through the entry of more fluid younger cohorts and the exit of less fluid older cohorts. Period changes, however, are differences in fluidity in the population at different historical time points that arise from some processes that affect all, or a large share, of the population, more or less simultaneously. The crucial issue, then, seems to be whether changes in transmissibility or class returns are spread across the entire working-age population or whether their influence is restricted to those cohorts entering the labor market.

We first consider changes in transmissibility: one can imagine circumstances in which these yield period effects, such as when the inheritance of property—an asset that may be transferred at any age—is suddenly discontinued (an example is the postwar collectivization of the Hungarian agricultural sector; cf. Erikson and Goldthorpe 1992, pp. 152–54). Nonetheless, transmissibility is overwhelmingly likely to lead to cohort effects for the simple reason that most transmission of assets occurs during childhood and early adulthood, manifested particularly in educational qualifications.

Changes in the class returns to assets will give rise to period effects on fluidity provided that these changes influence a sufficiently large share of the working population. Period effects are therefore most likely to be visible given the large-scale restructuring of labor markets, which involves the separation of workers from their current job and their reintegration into employment in a different class because the assets they possess no longer secure the same class positions.[9] We may observe this when, for instance, certain areas of the economy, such as manufacturing or mining in some industrialised countries, decline and their workers are relocated from jobs in the skilled manual class to unskilled work, or when new economic opportunities appear (as in the growth of the IT industry) which make certain assets more valuable than hitherto. In recent times the restructuring of the formerly state socialist economies of central and Eastern Europe has provided striking examples. Gerber and Hout (2004) show a period decline in social fluidity in Russia, and a similar finding is also reported for Hungary by Róbert and Bukodi (2004). Again, however, period change appears to follow rather uncommon events.

Changes in returns can lead to cohort, rather than period, effects when changes in the returns to an asset take effect gradually, as when the possession of an asset is associated with the occupation of a particular class or classes among older workers but not among younger ones. An example is found in the declining viability of small farms: typically, older men continue to work as farmers despite having relatively small amounts of land, whereas younger men who inherit such holdings are unlikely to become farmers. Indeed, we suspect that changes in the returns to assets, just like transmissibility, are more likely to appear as a cohort rather than a period effect. Recall that by returns to an asset we mean the returns in terms of the chances of acquiring a given class position. It is widely agreed that changes in an individual's social class position are relatively rare after the age of about 35 (Goldthorpe 1980, pp. 51–52, 69–71; Erikson and Goldthorpe 1992, p. 72; Breen 1994). In Sweden, between circa 1960 and 1990, around 80% of gainfully employed men and women were class stable between ages 35 and 40, and 80%–90% between ages 45 and 50 (Jonsson 2001, fig. 8.1). So any change in the class returns to different assets is likely to have its strongest effects on that share of the working-age population under about 35 years of age, with relatively little effect on the majority, except in unusual circumstances.

Our conclusion is that period change in fluidity is likely to be the exception rather than the rule and that most changes in fluidity that we observe in stable democratic societies will arise from processes of cohort replacement. Yet, studies of change in social fluidity using multiple observations from a country have almost always adopted a period, rather than a cohort, perspective (e.g., Luijkx and Ganzeboom 1989; DiPrete and Grusky 1990; Jonsson and Mills 1993; Vallet 1999; and the contributions to Breen 2004; exceptions include Hout [1988] and Hauser and Huang [1997]), and often there has been little explicit consideration of whether change is in fact driven by period or cohort effects or both.[10] For the Swedish case, however, we can now test the assertion that change in social fluidity is predominantly driven by cohort replacement. Because we have a series of repeated cross-sectional mobility surveys we can examine period change and cohort change simultaneously. If, indeed, fluidity is chiefly responsive to cohort-related factors, rather than to those associated with periods, then period change should disappear when we control for the former.

Given our theoretical model, and our conjecture that change in social fluidity is likely to be cohort driven, what changes do we expect to see in Sweden? First, there is reason to believe that a decline in λ, the returns to class position, has led to increasing social fluidity. Erikson (1996) showed that, over the course of the 20th century, more equal conditions among social classes in Sweden led to more equal educational attainments between children of different origins: this may have led to increasing fluidity during the 1980s and 1990s because of the more egalitarian conditions that prevailed during the childhood of the cohorts that entered the labor market then (who were born in the 1960s and the 1970s) as compared with the cohorts that they replaced (born approximately in the 1910s and 1920s). Political means toward this end include progressive taxation and redistribution mainly through the welfare state, for example by the introduction of the child allowance and various benefits in cash and in kind (Korpi and Palme 2004).

Political strategies to equalize opportunity in Sweden were not exclusively directed toward reducing inequality of condition. Policies also sought to reduce the *consequences* of such inequality (captured by τ_2 in fig. 1). These included the abolishing of fees for secondary and tertiary education, the introduction of free school meals and health care in schools, and free books and teaching aids in primary and secondary school, all of which have made educational opportunities less dependent on economic resources in the family of origin. Educational reforms, such as the comprehensive school reform in the 1950s, have aimed at increasing opportunities and educational attainment, particularly among those from less well-off families (e.g., Erikson and Jonsson 1996).

However, in Sweden as in many other European countries, for many years educational expansion per se was not an important strategy, and this contrasts with the United States, where expansion played a central role (Hout and Dohan 1996). Expansion could be seen as an attempt to

[9] Large-scale unemployment itself would not be sufficient to bring about an observed period effect because the convention in mobility research is to assign the unemployed and others without a job to the social class of their last occupation.

[10] But see Gerber and Hout (2004), whose empirical approach to distinguishing cohort from period effects is derived from our earlier work (Breen and Jonsson 2001).

weaken the importance of both τ_1 and τ_2 because parental assets and resources alike would be less decisive for admission to higher education. In our terms, expansion would promote equalization. But this supposes that children of less advantaged class backgrounds are the primary beneficiaries of the increased opportunities—something which is not necessarily the case. Jonsson and Erikson (in press) show that in Sweden expansion has been relatively unsuccessful in reducing educational inequality, probably because middle-class children with lower grade levels but high aspirations are often the first to take advantage of increasing opportunities. Nevertheless, expansion may lead to a reduction in the association between class origins and class destinations even though it does not weaken the inequality between class origins and educational attainment. This can come about if the conditions hold for the compositional effect we discussed earlier, so the returns to assets other than education (e.g., social networks) are higher at lower levels of education. Then educational expansion may lead to increased social fluidity even though it does not reduce the transmissibility of assets, but, rather, because it alters the average class return to other assets.

In sum, we expect that social fluidity has increased in Sweden and that this change has been largely driven by cohort replacement; and we believe that the process behind this reflects a combination of educational equalization and a compositional effect following educational expansion.

DATA AND VARIABLES

The data set is a compilation of the annual surveys of living conditions (ULF) 1976–99, conducted by Statistics Sweden (Vogel et al. 1998). Each survey is representative of the adult Swedish population, ages 15–75. The sample fraction is around one one-thousandth and the yearly sample sizes are around 6,000 respondents.[11] Nonresponse rates vary between 15% in the early surveys to around 22% in the later ones. Generally, the data quality is high. Some basic requirements for inferring change from repeated surveys are fulfilled: the same fieldwork and data preparation organization has been used throughout, classifications have followed national standards, and, since there has been a continuity of purpose, the

[11] The 1976–79 surveys have a household design, and from 1980 there is a panel element (where 50% of the sample is included in the sample eight years later). This may lead to within-person correlations across observations and so we use weights to adjust sample sizes accordingly. More specifically, respondents in households where both spouses are interviewed are down weighted by 0.5, as are those in a panel wave who have responded twice; those who have participated in three different waves are assigned a weight of one-third for each wave.

survey methods (face-to-face interviews) as well as the questionnaire design have been largely the same during the whole period.

The social class schema that we use (Andersson, Erikson, and Wärneryd 1981) is similar to the internationally used EGP class coding (Erikson and Goldthorpe 1992; Erikson, Goldthorpe, and Portocarero 1979), and therefore we use that notation.[12] We identify six classes. As can be seen in table 1, the distribution of classes among fathers is significantly different from that of sons and daughters.[13] The table reflects the upgrading of the social class structure with an increase in service-class jobs (classes I and II), as well as a sharp decline of the farming class.[14] Respondents' educational attainment—coded into six categories, using the CASMIN (Comparative Analysis of Social Mobility in Industrial Nations) educational classification (Müller and Shavit 1998, table 1.2b)—is shown in table 2. To exemplify the change in the educational structure we report the distribution of educational credentials for the older and the younger parts of our sample (born 1912–43 and 1944–74, respectively). The most noteworthy change is the rapidly decreasing numbers who leave school after the compulsory years and the expansion at tertiary levels—although, as can be seen, Sweden still has relatively few people with university degrees.

[12] Foremen, supervisors of manual workers, and lower grade technicians do not form a separate class (class V in the EGP schema). Instead some more qualified supervisors (such as *verkmästare*) go into class II, foremen are in general classified in class III, and blue-collar technicians (a group that is relatively uncommon in Sweden) go mostly into class VI. Occupations normally organized in the manual workers' trade union in Sweden (LO) are classified into classes VI and VII, and these include some occupations that in the EGP coding schema are found in the unqualified strata of the nonmanual classes (IIIb). Among these are lower-grade salespersons and shop assistants as well as lower-grade service workers (employed, inter alia, in hotels, restaurants, and in offices) and nurse's aides. It should be mentioned that the dividing line between class I and class II in the class schema was changed in 1982 leading to a small reduction (around two percentage points) of the percentages in the higher class (changes concerned both origin and destination classes). This is unlikely to have any major effects on the results, and our inspection of the results shows no break in the *OD* associations between the period up to 1981 and following 1982.

[13] The basis for the origin classification is a question on father's main occupation and employment status during the respondent's childhood (0–16 years of age). Unfortunately, before 1984 information about mother's occupation was collected only for respondents who did not live with their father during childhood, so only in those cases is information on the mother used to indicate origin class.

[14] The gradual decline of unskilled positions (both classes IIIb and VII) is evident from more detailed analyses. Here, we have merged the skilled workers with the unskilled, partly because we want to save degrees of freedom for our models, partly because these classes show fairly similar intergenerational mobility propensities (as well as high internal mobility). To avoid empty cells, we could also have merged farmers with the other self-employed, but previous research shows important mobility barriers to exist between jobs in the agricultural and nonagricultural sectors (Erikson and Goldthorpe 1992).

TABLE 1
MARGINAL DISTRIBUTIONS FOR ORIGIN (FATHER'S CLASS) AND DESTINATION CLASS FOR SONS AND DAUGHTERS

		ORIGIN	DESTINATION	
CLASS	DESCRIPTION	All	Men	Women
I	Professionals, higher executives or administrators, employers with 20 employees or more	6.0	14.7	8.1
II	Semiprofessionals, medium level administrators and officials	9.1	17.2	17.5
IIIa	Lower routine white-collar workers (jobs demanding some qualifications)	6.6	7.1	12.3
IVab	Self-employed, employers with 1–19 employees	14.6	10.3	4.3
IVcd	Farmers	16.6	2.8	2.1
IIIb,VI, VII ...	Unskilled workers in service, manual workers (skilled and unskilled)	47.0	47.8	55.8
Total		99.9	100.0	100.1

NOTE.—EGP social class schema; figures relate to repondents who were between 25 and 64 years of age in 1976–99.

It is notable that women, as in many other countries, have lower educational qualifications than men in older cohorts, but higher educational attainment than men in younger cohorts.

The analysis is confined to those ages 25–64 and undertaken separately for men ($n = 33{,}281$) and women ($n = 29{,}999$). We group the data into four-year tables, giving us six periods (1976–79, 1980–83, . . ., 1996–99).[15] Within each we identify 10 four-year age groups (25–28, 29–32, . . . , 61–64) which allow us to define 15 overlapping age cohorts (see table 3). The oldest cohort was born in 1912–18, the youngest in 1968–74. By definition not all cohorts can be observed in every period: the oldest and youngest are each observed only once, while the cohorts born in the 1930s and 1940s are observed in all six periods. The result is that the cross-tabulation of origin (O) by destination (D) by cohort (C) by period (P) is not rectangular: of the possible 3,240 ($= 6 \times 6 \times 15 \times 6$) cells of the table, 1,080 are structural zeroes.

[15] We used data for all men with a class code and all women who were not registered as home workers because among home workers there are some cases that might have their class position assigned to them on the basis of their spouse's occupation.

TABLE 2
EDUCATIONAL DISTRIBUTIONS FOR MEN AND WOMEN, BORN 1912–43 AND 1944–74

		MEN		WOMEN	
QUALIFICATION	DESCRIPTION	1912–43	1944–74	1912–43	1944–74
1ab	Compulsory	40.0	15.9	38.6	12.4
1c	Lower vocational	23.7	35.6	32.9	40.3
2ab	Lower secondary	4.0	4.2	7.4	5.1
2c	Upper secondary	14.6	15.8	4.4	9.8
3a	Lower-level tertiary	8.3	15.4	8.7	18.8
3b	University degree	9.3	13.2	8.0	13.6
Total		100.0	100.0	100.0	100.0

NOTE.—CASMIN educational schema.

RESULTS

Our analytical strategy is straightforward, but requires several steps. It is also rather technical, and so we here summarize the logic of the analyses and our main findings. We begin by testing for, and finding, change in social fluidity across birth cohorts, and then we carry out three analyses of changes between periods. In the first of these we ask whether, for each birth cohort, there is evidence that its fluidity changes as it ages: we find that there is not. Second, we use the panel element in the data to test whether, among those individuals who were interviewed on two occasions, patterns of fluidity differ at the later as compared with the earlier interview. Again, we find that this is not the case. Third, we make simple comparisons between each of the surveys in our data and ask whether fluidity changes between them. In this case we find that it does. Having found change between cohorts and between surveys we then ask whether the latter is in fact driven by processes of cohort replacement, and so we analyze cohort and period change in the same model. We find that when we control for cohort changes, period differences vanish. The final set of analyses then seeks to explain why fluidity changed over cohorts, and we do this by modeling the equalization and compositional effects of education and showing that, taken together, they are sufficient to account for all the observed cohort change in social fluidity.

Cohort Change

Table 4 shows the goodness of fit of six log-linear models fitted to the three-way ODC table.[16] Model 1 is the model of common social fluidity,

[16] Throughout we model social fluidity in a not very parsimonious way. As Hout (1988, pp. 1374–76) has noted, more powerful tests would use fewer degrees of freedom. But in this case, given our large sample size (63,280, almost 10 times bigger than Hout's

TABLE 3
THE RELATIONSHIP BETWEEN COHORTS, AGE GROUPS, AND PERIODS

AGE GROUPS	PERIODS 1976–79	1980–83	1984–87	1988–91	1992–95	1996–99
25–28 ...	10	11	12	13	14	15
29–32 ...	9	10	11	12	13	14
33–36 ...	8	9	10	11	12	13
37–40 ...	7	8	9	10	11	12
41–44 ...	6	7	8	9	10	11
45–48 ...	5	6	7	8	9	10
49–52 ...	4	5	6	7	8	9
53–56 ...	3	4	5	6	7	8
57–60 ...	2	3	4	5	6	7
61–64 ...	1	2	3	4	5	6

NOTE.—Nos. 1–14 in the table cells represent cohorts. Birth years are, for cohort 1, 1912–18; cohort 2, 1916–22; cohort 3, 1920–26; cohort 4, 1924–30; cohort 5, 1928–34; cohort 6, 1932–38; cohort 7, 1936–42; cohort 8, 1940–46; cohort 9, 1944–50; cohort 10, 1948–54; cohort 11, 1952–58; cohort 12, 1956–62; cohort 13, 1960–66; cohort 14, 1964–70; cohort 15, 1968–74.

stating that the OD association is the same in all cohorts: this can be rejected using the standard chi-squared goodness of fit test.[17] In model 2 a common pattern of local OD association is assumed to vary log multiplicatively over cohorts in what has come to be known as a model of uniform difference, or unidiff (Erikson and Goldthorpe 1992; Xie 1992). Specifically, letting $i = 1, \ldots, 6$ and $j = 1, \ldots, 6$ index origins and destinations, respectively, and $k = 1, \ldots, 15$ cohorts, we write

$$\ln\theta_{ij|k} = \beta_k \ln\theta_{ij}. \tag{1}$$

Here $\theta_{ij|k}$ denotes the odds ratio

$$\frac{F_{ijk}/F_{ij'k}}{F_{i'jk}/F_{i'j'k}}, \tag{2}$$

where F_{ijk} is the fitted value in the ijkth cell of the table and β_k is a cohort-specific multiplier, and we set $\beta_1 = 1$.[18] The model says that the difference in the log-odds ratios between any two cohorts is proportional to the

sample), even the global tests that we use will have reasonable power to detect change. In table 4 and all other tables we report both the deviance (likelihood-ratio chi-square) and the BIC statistic (Raftery 1986). We base our model selection on the former; we report the latter because it is a widely used criterion of fit, though, for reasons advanced by Weakliem (1999), we do not make use of it.

[17] Henceforth when we write that something is or is not significant we are referring to the $P < .05$ criterion.

[18] In table 3 and subsequently we attach a superscript to β to indicate the variable over which fluidity is changing.

TABLE 4
GOODNESS OF FIT OF MODELS FOR THE ORIGIN × DESTINATION × COHORT TABLE FOR MEN AND WOMEN

			MEN (n = 33,281)			WOMEN (n = 29,999)		
NO.	MODEL	df	Deviance	P	BIC	Deviance	P	BIC
1 ...	OD	350	402.86	.027	−3,242	423.47	.004	−3,185
2 ...	$OD\beta^C$	336	394.81	.014	−3,104	366.26	.123	−3,098
3 ...	$OD\beta^C{}_k$+diag	330	315.85	.703	−3,120	350.76	.207	−3,051
4 ...	$ODk\beta^C$+diag	343	331.12	.667	−3,240	358.79	.268	−3,177
5 ...	$ODk\beta^C$	349	398.65	.033	−3,235	379.01	.130	−3,219

NOTE.—All models include the terms OC DC. O = origin; D = destination; C = cohort; diag = 6 parameters fitted to cells on main diagonal of the O-D table; $k = 1, \ldots, 15$.

difference in their β parameters, and thus declining values of this parameter over cohorts correspond to increasing social fluidity. This model uses 14 degrees of freedom (i.e., number of cohorts minus one) more than the independence model, but reduces the deviance by only eight for men. For women, however, there is evidently more change as captured by the unidiff model—the reduction in deviance is around 57, which is clearly a significant improvement in fit. The third model introduces parameters for each of the cells on the main diagonal of the mobility table. This imposes the constraint that the sum, over all cohorts, of the frequencies in each cell of the main diagonal of the origin by destination table is fitted exactly. The log-odds ratios under this model are given by

$$\ln\theta_{ij|k} = \beta_k \ln\theta_{ij} + \delta_{ij} - \delta_{ij'} - \delta_{i'j} + \delta_{i'j'},$$

$$\delta_{ij} = 0, \quad \text{if } i \neq j. \tag{3}$$

Here δ_{ij} denotes the diagonal parameters. Because these parameters are constant over cohorts, the difference between cohorts in their log-odds ratios is still proportional to the size of their β coefficients. As equation (3) makes clear, the diagonal effects operate over and above the evolution of the pattern of local odds ratios and so, even though the diagonal parameters do not vary over cohorts, this does not mean that propensities for individuals to be found in their class of origin remain constant over cohorts: rather, this propensity changes according to the common log-multiplicative evolution of the whole pattern of local association. What does remain constant is the tendency for class inheritance that exists over and above that which is implied by this pattern.[19] Adding these diagonal

[19] We tested models in which the diagonal parameters were allowed to vary over cohorts, but this never yielded a significant improvement in fit.

parameters makes a substantial difference to the fit of the men's tables, though rather less difference for women. This is not surprising, since it is well known that men are more likely than women to be found in the same class as their father (e.g., see Jonsson and Mills 1993 for the Swedish case). In model 4 we impose a linear constraint on the evolution of the β parameters: so now we write

$$\ln\theta_{ij|k} = k\beta\ln\theta_{ij} + \delta_{ij} - \delta_{ij'} - \delta_{i'j} + \delta_{i'j'},$$

$$\delta_{ij} = 0, \quad \text{if } i \neq j, k = 1, \ldots, 15. \tag{4}$$

For both men and women model 4 is preferred to model 3 and to model 1. Last, model 5 removes the diagonal effects but retains the linear constraint. This considerably worsens the goodness of fit of the model for men, but has less impact on women, once again illustrating the lesser importance of inheritance effects among women.

Figure 2 shows the parameter values for the diagonal cells generated by model 3. As is evident, there is strong inheritance in the upper service class—for women, in particular—and in the petty bourgeoisie; and among men inheritance effects are very strong in the farming class.[20] Of more interest to us, however, are the β coefficient estimates from model 3 and also the linear β estimated from model 4, both of which are shown in figure 3. They display a trend toward increased social fluidity over cohorts, with an estimated slope for the association between origins and destinations of $-.036$ for men and $-.043$ for women, though the cohort-specific βs suggest that, from the 1948–54 cohort onwards, the trend toward more fluidity disappears.[21] It should be borne in mind that for the oldest and youngest cohorts we have only one observation, two for the second oldest and second youngest, and so on, so the end points on this and the other figures depicting cohort change should be regarded as particularly subject to uncertainty (and they also exercise less influence on the estimate of the slope). If we compare the β_k coefficients for the 1964–70 cohorts, we see that they are two-thirds or less of the value for those for the 1916–22 cohorts, indicating a substantial increase in fluidity over the 20th century.

Age Effects

Having established that fluidity changed over cohorts in 20th-century Sweden, we now turn to period change. The first question we address is whether there is any indication of substantive changes in fluidity within cohorts as they progress in their career: in other words, an age effect.[22] We cannot address this issue for our very oldest and youngest cohorts, for each of which we have only one table. Table 5 shows the result of fitting, to each of the remaining 13 cohorts, a model that assumes constancy over periods in the *OD* association (*PO PD OD*) and a model of uniform change in fluidity across periods (*PO PD OD* β^P). In every case the model of common fluidity fits the data and the model of uniform change never yields a significant improvement in fit. The result is overwhelmingly comforting for the cohort view: there seems to be no substantive change in fluidity across periods for people in a given cohort.[23]

This conclusion can be checked by using the panel element in our data, though with a smaller sample.[24] For some respondents we have information on their class position at two points, eight years apart, which we call *D*1 and *D*2. Information on *D*1 was collected at the initial ULF interview between 1979 and 1991 and on *D*2 between 1986 and 1999.

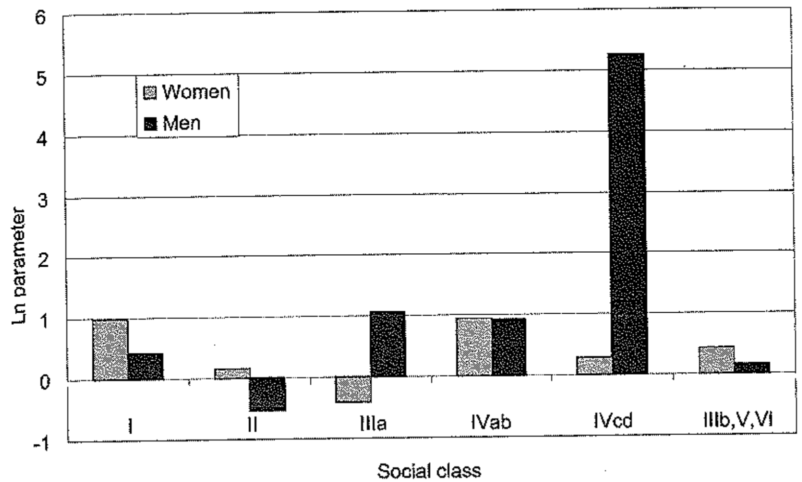

FIG. 2.—Parameter values for the diagonal cells generated by model 3

[20] There is also some "disinheritance" among men in class II and women in class IIIa (see also figs. 7 and 8, below), suggesting that these are origin classes which are disproportionately likely to be vacated by those born into them.

[21] We included a quadratic term in addition to the linear term for the evolution of the log-odds ratios, but this was not a significant improvement for either sex.

[22] An age effect cannot explain change over periods, since, by definition, an age effect is specific to an age group but constant over periods and cohorts. Thus, to explain change, age effects must change—as they would, for example, in an age by period interaction.

[23] Note that this does not mean that members of a cohort do not change occupations or classes across their careers (they most certainly do, at least up to the age of 30–40), only that these changes are unrelated to their social origins.

[24] We thank an *AJS* reviewer for suggesting this analysis.

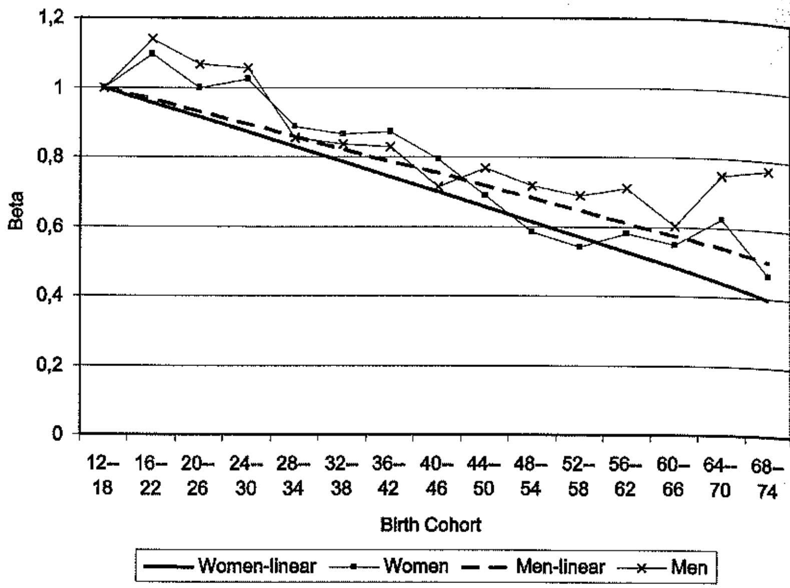

FIG. 3.—β coefficient estimates from model 3; linear β estimated from model 4

Respondents born before 1921 were too old, and those born after 1966 too young, to have a value on $D2$: furthermore, respondents born between these dates may also not have a value for $D2$ depending on how old they were when they were first interviewed. Finally, only half of each year's sample was reinterviewed eight years later. As a result we have panel data on 9,906 respondents (compared with the 63,280 used elsewhere in our analyses).

Because our goal is to check whether the origin-destination association is the same over the life course we form the table of origins by $D1$ by $D2$ for each sex, and distinguish two broadly defined cohorts. In the older cohort both $D1$ and $D2$ occurred after the age of 34: it is made up of those born 1921–44 together with those born 1945–48 who were 35 years old or older at $D1$. In the younger cohort $D1$ occurred before 35. These respondents were born 1949–66 or were born 1945–48 and were 34 years old or younger at $D1$. Forming two cohorts in this way allows us to distinguish those whose life-course mobility may have occurred before the age of occupational maturity (which we take to be 35) and those whose mobility, if any, occurred after this point. The small number of cases forced us to move to a four-class categorization for origins, first destination ($D1$), and second destination ($D2$). We distinguish class I, clas-

TABLE 5

SOCIAL FLUIDITY MODELS FITTED TO EACH COHORT

			DEVIANCES			
			MEN		WOMEN	
COHORT		TABLE	*PO PD OD*	*PO PD OD*β^P	*PO PD OD*	*PO PD OD*β^P
2	1916–22	2	15.42	15.32	25.19	19.02
3	1920–26	3	44.82	41.17	31.75	30.73
4	1924–30	4	64.57	64.24	66.41	65.89
5	1928–34	5	72.81	70.65	102.92	100.36
6	1932–38	6	119.65	112.28	131.04	125.34
7	1936–42	6	105.41	103.97	104.79	99.73
8	1940–46	6	109.11	108.18	137.88	134.43
9	1944–50	6	103.28	98.53	85.02	79.10
10	1948–54	6	119.55	113.17	136.71	129.14
11	1952–58	5	112.66	107.43	108.62	98.80
12	1956–62	4	61.24	61.07	70.57	68.34
13	1960–66	3	50.62	49.69	46.08	45.97
14	1964–70	2	22.27	22.17	22.78	22.20

NOTE.—Degrees of freedom for *PO PD OD* model = (no. of tables − 1) × 25, and degrees of freedom for *PO PD OD*β^P model = (no. of tables − 1) × 24; *df* for the comparison of the two.

ses II and IIIa, classes IVa, IVb, and IVc, and classes IIIb, VI, and VIII.[25] The test is simple: we fit the model $OD1$ $OD2$ $D1D2$ and we compare its goodness of fit with the same model in which the $OD1$ and $OD2$ associations are fixed to be the same (which forces the association between origins and destinations to be the same for both destinations). The test has nine degrees of freedom and in no case is the constrained model a significantly poorer fit: the deviances are 13.25, 13.86, 10.37, and 14.87 for the older and younger cohorts of men and the older and younger cohorts of women, respectively.[26]

Period Change

The absence of change within cohorts as they age does not mean, of course, that social fluidity does not change across periods, and so table 6 uses the same set of five models as table 4, this time applied to test for period change, now ignoring cohorts. Evidence for period change is much weaker than for cohort change, but adding parameters for the main diagonal once

[25] The sample sizes in each of the origin by $D1$ by $D2$ tables are 3,493, 1,503, 3,375, and 1,535 for older and younger men and older and younger women, respectively.

[26] We carried out the same test using the model $OD1$ $OD2$ (i.e., omitting the $D1D2$ term capturing association in life course mobility) with the same results (details available on request from the authors).

TABLE 6

Goodness of Fit of Models for the Origin × Destination × Period Table for Men and Women

			Men (n = 33,281)			Women (n = 29,999)		
No.	Model	df	Deviance	P	BIC	Deviance	P	BIC
1 ...	OD	125	180.3	.001	−1,121	184.6	.001	−1,104
2 ...	$OD\beta^P$	120	167.5	.003	−1,082	167.4	.003	−1,070
3 ...	$OD\beta^P_l$+diag	114	133.3	.105	−1,054	160.0	.003	−1,015
4 ...	$ODl\beta^P$+diag	118	142.8	.060	−1,086	167.8	.002	−1,049
5 ...	$ODl\beta^P$	124	174.4	.002	−1,117	175.1	.002	−1,103

Note.—All models include the terms OP DP: O = origin; D = destination; P = period; diag = 6 parameters fitted to cells on main diagonal of the O–D table; l = 1, . . . , 6.

again leads to a large improvement in fit for men (compare models 2 and 3). On the other hand, models 3 and 4 again fit the data and are a significant improvement over both models 1 and 5.[27] Comparing models 3 and 4, the latter is more parsimonious. Among women, although none of the models fits the data according to the deviance, model 5 would be preferred according to this criterion, once again indicating the lesser importance of inheritance. If, however, we take model 4 as the preferred model for both sexes, we find that the slope of β is −.060 for men and −.050 for women.[28]

Period and Cohort Change

Having established a trend toward increasing social fluidity over both cohorts and periods, we turn, in table 7, to models that allow for both types of change. All the models reported in table 7 fit the OPC and DPC margins: thus we allow the origin distribution and the destination distribution to vary over both cohorts and periods. Our interest is in the OD association, and we test whether, given change over periods, cohort change persists, and vice versa. We take as the point of departure two models of change in social fluidity incorporating both period and cohort change, as before using the unidiff model in its unconstrained (models 2–4) and linear (models 5–7) forms to model change (including time constant parameters

[27] There is no evidence that the diagonal parameter values vary over periods.

[28] If period fluidity were simply a weighted sum of the fluidity in each cohort represented in that period's table, and if all cohorts were the same size, then the period and cohort slopes in our data would be identical (within sampling error). This is because, given a cohort slope −b, the association in an entering cohort in any period is equal to (−b × 10 × the association in the exiting cohort), and as there are 10 cohorts in every period survey, the difference between periods is then −b.

TABLE 7

Goodness of Fit of Models for the Origin × Destination × Cohort × Period Table for Men and Women

			Men (n = 33,281)		Women (n = 29,999)	
No.	Model	df	Deviance	BIC	Deviance	BIC
1 ...	OD	1,475	1,404.28	−13,955	1,493.23	−13,712
2 ...	ODβ^P_l+diag	1,464	1,361.65	−13,883	1,472.42	−13,620
3 ...	ODβ^C_k+diag	1,455	1,317.37	−13,883	1,420.52	−13,579
4 ...	OD$\beta^P_l\beta^C_k$+diag	1,450	1,311.85	−13,787	1,412.11	−13,536
5 ...	ODlβ^P +diag	1,469	1,370.84	−13,925	1,481.12	−13,663
6 ...	ODkβ^C+diag	1,469	1,332.54	−13,964	1,428.55	−13,715
7 ...	ODkβ^Clβ^P+diag	1,468	1,332.18	−13,954	1,427.77	−13,706

Note.—All models include the terms OPC DPC. P-values are not reported: they are all greater than 0.4. O = origin; D = destination; P = period; C = cohort; diag = 6 parameters fitted to cells on main diagonal of the O–D table; k = 1, . . . , 15; l = 1, . . . , 6.

for the main diagonal). These models allow a common pattern of fluidity to vary over either or both the C and P margins simultaneously: in other words, the pattern of social fluidity is multiplied by a cohort-specific and a period-specific β parameter. Ignoring, for ease of notation, the effects of the diagonal parameters, in the case of model 4 we have

$$\ln\theta_{ij|kl} = \beta^C_k\beta^P_l\ln\theta_{ij}, \tag{5}$$

where k indexes cohorts and l periods, while for model 7 we have

$$\ln\theta_{ij|kl} = k\beta^C l\beta^P\ln\theta_{ij},$$

$$\text{for } k = 1, \ldots, 15, \text{ and } l = 1, \ldots, 6. \tag{6}$$

The results show that, given a model which includes cohort change, the addition of change over periods (model 3 compared with model 4, and model 6 compared with model 7) does not significantly improve the fit of the model, but the reverse is not the case (model 2 compared with 4 and model 5 compared with 7). That is to say, when we control for cohort effects, the period effects vanish. Reflecting this, in model 7, which is the direct counterpart to model 4 in the earlier cohort and period analyses, the partial period slope is now estimated as being not significantly different from zero (−0.008 for men and +0.015 for women), whereas the partial cohort slope is almost unchanged from its unconditional value (−.034 for men, compared with −.036 and −.045 for women, compared with −.043). These results, together with the finding, reported in table 5, that fluidity does not vary within a given cohort, and the results of our analysis of the panel element in our data, lead us to conclude that period change in fluidity in the last quarter of the 20th century was the consequence of the replacement of older, less fluid cohorts, with younger, more fluid ones.

Understanding Cohort Change

How are we now to understand the change in social fluidity across cohorts? Perhaps the most plausible mechanism to explain cohort change concerns the transmission of assets during childhood, pointing to the important role of education for social mobility. Increasing social fluidity may then come about in two different ways, which we earlier labeled *equalization* and *compositional effects*. We will look at each of these in turn.

A common explanation of increasing social fluidity is that it is driven by a weakening association between social origin and educational qualifications. Such a weakening did occur in Sweden among cohorts born approximately between the early 1920s and the 1950s (Jonsson and Erikson 2000), and this weakening is likely to have brought about changes in the gross association between origins and destinations of the kind that we have shown above. In our data we find the same result. Table 8 reports three models fitted to the origin × education × cohort table, the first of which is constant association between origins and education (*OE*), the second of which allows for uniform change in the *OE* association over cohorts, and the third of which constrains this change to be linear. As with the *OD* association, the model of linear change provides a good account of the data. In figure 4, the β coefficients from the second of these models show the decline in the association flattening out after the 1950s: this is very similar to the trend shown in figure 3, which provides some prima facie evidence that the change in social fluidity was driven, in some part, by changes in educational inequality. Our estimates, which we presented below, suggest that around half of the total association between

TABLE 8

Goodness of Fit of Models for the Origin × Education × Cohort Table for Men and Women

			Men (n = 33,281)			Women (n = 29,999)		
No.	Model	*df*	Deviance	*P*	BIC	Deviance	*P*	BIC
1 ...	OE	350	442.66	.001	−3,203	406.32	.020	−3,202
2 ...	$OE\beta^C{}_k$	336	386.19	.031	−3,112	365.61	.128	−3,098
3 ...	$OEk\beta^C$	349	405.01	.021	−3,229	385.43	.087	−3,212

Note.—All models include the terms OC EC. O = origin; D = destination; C = cohort; E = education, $k = 1, \ldots, 15$.

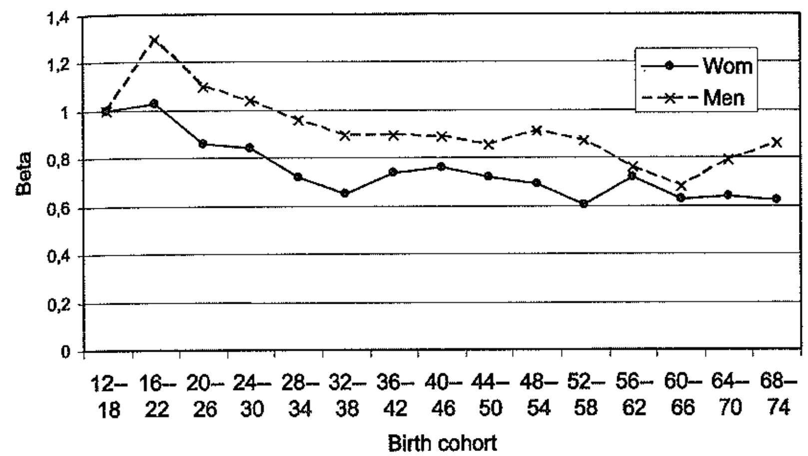

Fig. 4.—β coefficients for uniform change in the OE association over cohorts

origins and destinations is mediated via education, showing that a weakening of this indirect path through a decline in the origin-education association can indeed lead to increases in social fluidity.

Table 9 uses the origin × destination × education × cohort table to address the compositional question. Here there are two sets of models: 2–4 fit completely unrestricted log-linear models, and 5–7 fit unrestricted log-multiplicative uniform difference models. In models 5 and 7 we denote the log-multiplicative evolution of the *OD* association over educational levels as $OD\beta^E_m$, where m indexes educational levels from 1 to 6. In both sets of models we reach the same conclusion: once we allow for variation in the *OD* association between levels of education, there is no significant change in fluidity over cohorts. So, model 2 is not a poorer fit to the data than model 4, but model 3 is, indicating that the omission of the *ODC* term is not statistically significant once the *ODE* term is included in the model, whereas the *ODE* term cannot be omitted even when *ODC* is included. Likewise, model 5 is not a poorer fit than model 7, but model 6 is, indicating that the $OD\beta^E_m$ term is required but that the $OD\beta^C_k$ term is not.

This result does not mean that it is the compositional effect that is wholly responsible for partialling out cohort change, because the model of necessity includes the *OEC* and *EDC* terms: thus the trend in fluidity over cohorts may depend on both the equalization effect—which is already included in the models reported in table 9—and on the compositional effect. The basis of this compositional effect is shown in figure 5, which reports the β coefficients for each educational level, taken from model 5

TABLE 9

GOODNESS OF FIT OF MODELS FOR THE ORIGIN × DESTINATION × EDUCATION × COHORT TABLE FOR MEN AND WOMEN

			MEN (n = 33,281)		WOMEN (n = 29,999)	
No.	MODEL	df	Deviance	BIC	Deviance	BIC
1 ...	OD	2,225	1,928.98	−21,239	1,918.85	−21,018
2 ...	ODE	2,100	1,677.76	−20,189	1,689.20	−19,960
3 ...	ODC	1,875	1,571.07	−17,953	1,545.22	−17,784
4 ...	ODE ODC	1,750	1,335.32	−16,887	1,326.36	−16,714
5 ...	$OD\beta^{E}_{m}$+diag	2,214	1,841.16	−21,213	1,875.45	−20,948
6 ...	$OD\beta^{C}_{k}$+diag	2,205	1,877.20	−21,083	1,874.66	−20,857
7 ...	$OD\beta^{E}_{m}\beta^{C}_{k}$+diag	2,200	1,834.34	−21,074	1,859.45	−20,820

Model Comparisons	Term Tested	df	Deviance	P	Deviance	P
2 versus 4	*ODC*	350	342.44	.603	362.84	.307
3 versus 4	*ODE*	125	235.75	.001	218.86	.001
5 versus 7	$OD\beta^{C}_{k}$	14	6.82	.941	16.00	.313
6 versus 7	$OD\beta^{E}_{m}$	5	42.86	.001	15.21	.010

NOTE.—All models include the terms *OCE DCE*. O = origin; D = destination; P = period; C = cohort; E = education; diag = 6 parameters fitted to cells on main diagonal of the O-D table; k = 1, . . ., 15; m = 1, . . . , 6.

of table 9. This shows a clear pattern, though not a linear one, of weaker association between origins and destinations at higher educational levels. Thus, as successive cohorts have come to have higher levels of education (see table 2) so the gross association between origins and destinations has weakened.

This effect is illustrated in figure 6, which shows the weighted sum of the estimated β values for each educational level in every cohort (where the weights are the relative sizes of the educational categories). In other words, this would be the social fluidity in each cohort if that were simply the weighted sum of the fluidities in each educational category. The picture shown in figure 6 is similar to that in figures 3 and 4: a decline, parallel for both sexes, until the cohorts born in the 1950s, then stability. The temporal coincidence of the equalization and compositional effects derives from the fact that the equalization that affected cohorts born in the first half of the century had the consequence of expanding the middle and upper levels of the educational system; this expansion translated into a compositional effect because labor markets were more meritocratic the higher the level of educational qualifications.

Although the weighted averages of the educational level βs in each cohort point to similar trends to those shown by the cohort βs themselves, the relationship between the unconditional cohort trends and the corresponding trends across educational levels is more complicated than this.

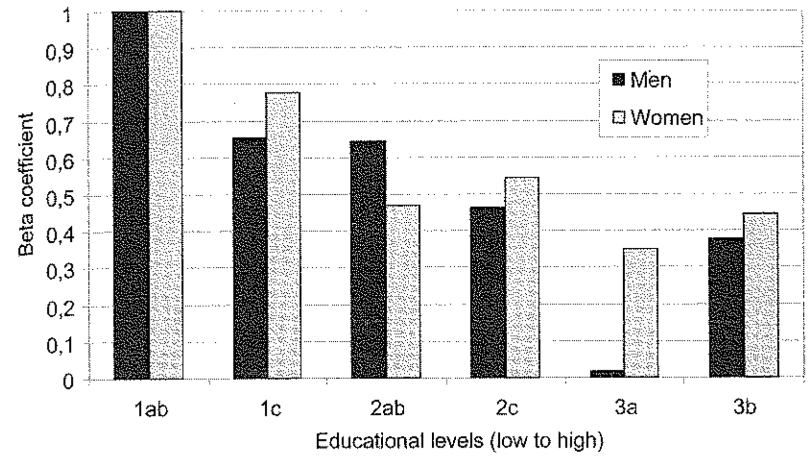

FIG. 5.—β coefficients for each educational level

Drawing on results by Goodman (1972, pp. 1070–75), we know that we can find an unconditional cohort effect, (that is, *ODC*) even when the partial *ODC* terms are zero, provided that the partial *ODE* and *EC* associations are nonzero. From this it follows that the three-way *ODC* term when *ODE* is not in the model depends not only on the educational effect but also on the distribution of educational categories across periods. It is therefore not a simple matter to infer what pattern of unconditional cohort effects is implied by a given set of education effects, and figure 6 should be taken only as illustrative.

Education and Social Fluidity

In the final part of our analysis we try to make a more formal assessment of the effects of educational equalization and compositional change on the trend over cohorts in social fluidity. Essentially this involves determining how much of the association between origins and destinations is mediated via educational attainment and, following from this, how much of the change in fluidity over cohorts comes about through, on the one hand, changes in the effects of origins on educational attainment and of educational attainment on class destinations, and, on the other, the shift of the population into those educational categories in which origins have a weaker effect. Given continuous measures of social position we could do this using path analysis, but with categorical variables this is not possible. We therefore use an approximation, following Breen and Luijkx (2004*a*).

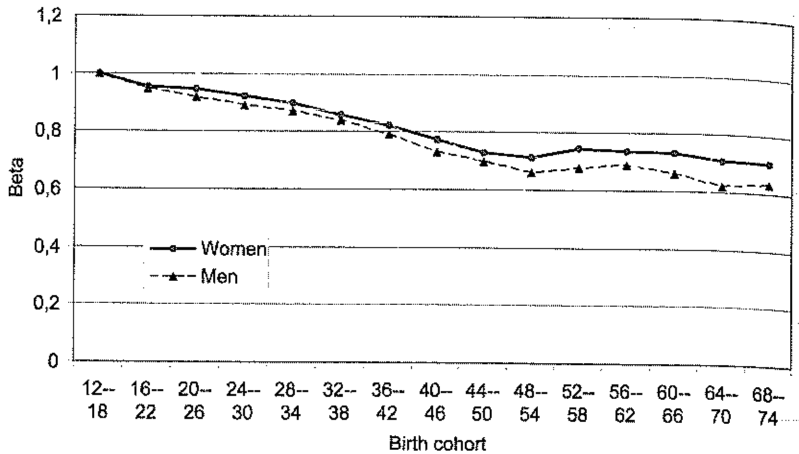

FIG. 6.—Weighted sum of estimated β values for each educational level, every cohort

We must begin with a measure of the evolution of the gross *OD* association over cohorts, and so we could fit the log-multiplicative $OD\beta^C$ model to the three-way origin by destination by cohort table (as reported in the models of table 4). The obvious next step might then be to use the four-way table of origins by destinations by cohorts by education to fit a model which includes the partial effects of education on destination controlling for origins, and the partial effects of origin on destination controlling for education. This would, in fact, be one of the models reported in table 9 and, if the partial *OD* association were fitted using a log-multiplicative specification, it might seem that we could compare the βs from this model (the partial βs) and compare them with the βs from one of the models of table 4 (the gross βs). But such a comparison would be invalid because the pattern of local *OD* association will be different in the two cases: that is, the pattern of *OD* association that evolves log multiplicatively over cohorts depends on whether we control for the effect of education on destinations or not.[29] We attempt to overcome this difficulty by constraining the pattern (though not the strength) of the *OD* association in the partial model to be the same as the estimated gross *OD* association. This allows us to make a comparison of the β parameters from the two models and so measure the relative strength of the association with and without controlling for the effects of education.

[29] This is because the local *OD* association and the β parameters are estimated together in the log-multiplicative model. In the unconditional model they depend on the gross association between origins and destinations, whereas in the conditional model they depend on the association between origins and destinations holding education constant.

But things are not so simple because we can reasonably suppose that the pattern (and not just the strength) of the *OD* association will differ significantly depending on whether education is in the model or not. Educational attainment is a more important asset for mobility to some class destinations rather than others: in particular, entry into self-employment or farming among children born into these classes is a question of inheritance, rather than of educational attainment (Ishida, Müller, and Ridge 1995). With this in mind we use, as our baseline model, model 4 of table 4, which includes the set of parameters applied to the main diagonal of the table (but whose effects are held constant over cohorts and which we henceforth referred to as "diag").

Our model for the unconditional *OD* association can be written $OC\ DC\ ODk\beta^C$ + diag: this allows the *OD* association to evolve linearly over cohorts. The first of our models for the conditional or partial *OD* association is $OEC\ EDC\ X^{OD}k\beta^C$ + diag. X^{OD} here represents the *OD* association which is fixed to be equal to that estimated from the gross model, and this is, once again, constrained to evolve linearly over cohorts. We allow the diagonal parameters to differ between the partial and gross models and we fit the *OEC* and *EDC* margins exactly in order to focus on the difference between the βs from the gross and partial models. Comparing the β^C estimates tells us the extent to which the *OD* association, and its trend over cohorts, weakens once we take into account the association between origins and education and that between education and destination.

However, we can go further than this and fit a second partial model in which we allow the association to vary over both cohorts and educational levels as follows: $OEC\ \ EDC\ \ X^{OD}\beta^E_m k\beta^C$ + diag. Now the β^C parameter tells us the slope of the *OD* association over cohorts when we also allow that association to vary over educational levels. Here the constrained *OD* association, X^{OD}, evolves linearly over cohorts, as before, but varies freely over educational levels.

Table 10 contains the results of these analyses. Model 1 repeats model 4 of table 4. Model 2 is a poorer fit to the data than the counterpart model, which preserves the linear trend but estimates the *OD* association freely (the difference in deviance is 60.4 for men and 88.3 for women; 25df), but nevertheless still provides an adequate fit to the data,[30] as does model 3. Once we control for education (model 2), the strength of the *OD* association—as measured by the value of the log-odds ratios in model 2

[30] The large increase in the deviance for model 2 compared with model 6 of table 8 suggests that the pattern of the OD association in the off-diagonal cells also changes when education is taken into account—but this is something which our model cannot capture.

TABLE 10

GOODNESS OF FIT OF MODELS TO DECOMPOSE THE OD ASSOCIATION

Models	Description	Deviance	df	Relative Log-Odds Ratio*		Proportional Slope	Absolute Slope	
Men:								
1	OC DC OD$k\beta^C$ + diag	331.12	343	1		−.036	−.036	
2	OEC EDC $X^{OD}k\beta^C$ + diag	1,956.59	2,230	.53		−.057	−.030	
3	OEC EDC $X^{OD}\beta_m{}^E k\beta^C$ + diag	1,892.29	2,225	.37	.01	−.034	−.013	−.017
				.43	.52		−.018	−.015
				.51	.68		−.000	−.013
Women:								
1	OC DC OD$k\beta^C$ + diag	368.79	343	1		−.044	−.044	
2	OEC EDC $X^{OD}k\beta^C$ + diag	1,993.71	2,230	.58		−.066	−.038	
3	OEC EDC $X^{OD}\beta_m{}^E k\beta^C$ + diag	1,947.42	2,225	.10	.20	−.049	−.033	−.034
				.51	.43		−.021	−.025
				.70	.68		−.010	−.005

NOTE.—For men, n = 33,281; for women, n = 29,999.
* Set to 1 for model 1.

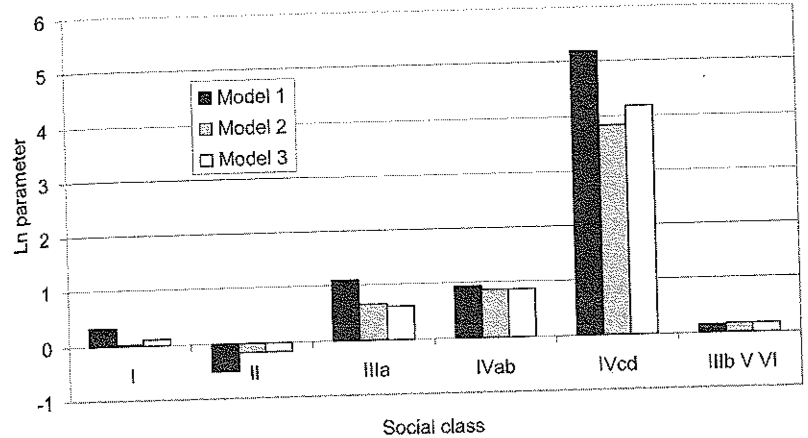

FIG. 7.—Diagonal parameter estimates from models 1–3: men

compared with model 1—is reduced by just less than half. That is, once we control for the path that links origins to destination via education, the direct effect of origins on destination is about halved. The slope of the log-odds ratios over cohorts, expressed as a *proportionate* decline in the association for that model, strengthens in model 2 compared with model 1; however, if we express the slope as the change in the *absolute* value of the log-odds ratios then it is only slightly weaker in model 2 than in model 1.[31] So, although controlling for class inequality in educational attainment accounts for a good deal of the association between origins and destinations, it does not explain much of the trend of change in this association over cohorts. In model 3 we also allow the *OD* association to vary over educational levels, and we report the strength of the origin-destination association at each educational level and also the absolute slope within each educational level. In both cases the figures run from the lowest to the highest educational level, and they should be read by row. Not only does the association vary in strength quite noticeably over educational levels (as we already saw in fig. 5), but, among men, the absolute slopes are quite close to zero. Among women they are close to zero at the higher educational levels but somewhat further from zero at the lower levels. These results suggest that the declining trend in the association between origins and destinations is mainly due to the com-

[31] The proportional slope reports the proportional decline per birth cohort: so, in model 2, for men this is just less than 6%. But this is relative to the initial association, which, in model 2, is weaker than that in model 1. If we then ask, What would this decline be as a share of the original association? (i.e., that in model 1), we find that this 6% proportional reduction equates to a 3% absolute decline.

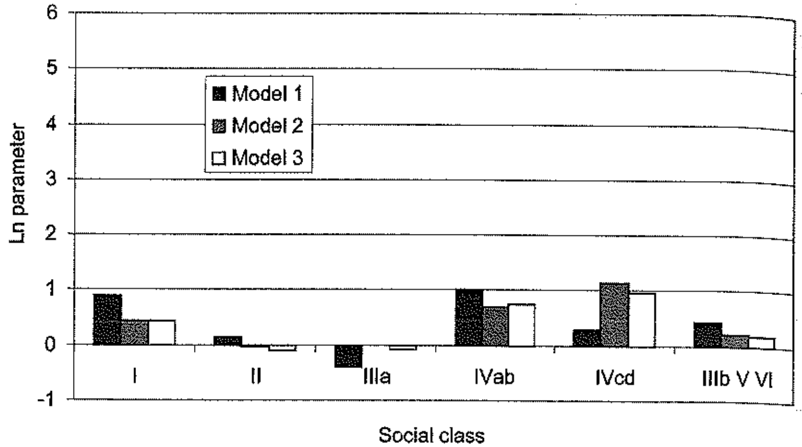

FIG. 8.—Diagonal parameter estimates from models 1–3: women

positional effect of educational expansion rather than to the process of educational equalization. Finally, the diagonal parameter estimates from the three models reported in table 10 are shown in figures 7 and 8. The first bar of the histogram for each class repeats that shown in figure 2. Overall the parameters show remarkably little change, indicating that the tendency for classes to be self-recruiting (where this tendency exists) operates largely independently of the educational mechanism.

CONCLUSION

Previous research on social fluidity has not been successful in accounting for temporal change; nor, indeed, have researchers agreed on the extent to which modern societies are characterized by change at all. Our analysis documents trends toward greater openness in Sweden, and we extend previous research by addressing the question of how this change came about. On the basis of a theoretical model of the intergenerational transmission of class position we argue that change in social fluidity will, under normal circumstances, be driven by cohort replacement rather than by period effects. Following from this, we pay special attention to the way in which changes in children's educational attainment can account for changing fluidity.

The Swedish data are particularly suitable for addressing the question of temporal change: we have access to 24 annual surveys covering the period 1976–99, with cohorts born between 1912 and 1974, and which have comparable classifications of social class origins, educational qualifications, and class destinations (derived from current job). Our first conclusion is that social fluidity increased in Sweden, particularly among those cohorts of men and women born during the first half of the 20th century. We also observed a trend toward greater fluidity in the working-age population during the last quarter of the 20th century, but this period change disappears when we control for differences between birth cohorts. Moreover, we show that social fluidity does not vary across periods within a cohort, and, using the panel element in the data, we find that fluidity does not change across the life course. These results led us to conclude that period change is, in fact, driven by a process of cohort replacement. To the extent that change in social fluidity generally comes about through cohort replacement, this may explain why many sociologists have failed to discern any period change when analyzing two or more surveys that are not so many years apart and that therefore mainly comprise samples from the same cohorts. It may well be the case that true period changes are likely to occur only in specific, and perhaps rather dramatic, circumstances, but historical changes that equalize the opportunities for successive birth cohorts may have a substantial impact that is only visible in a longer time perspective.

Our second conclusion is that the evolution of fluidity over cohorts in Sweden has been driven by educational equalization and by a compositional effect based in the changing educational distributions of successive birth cohorts. Hout and Dohan (1996) have portrayed the U.S. and Swedish "strategy of educational equality" as very different: the first seeks to expand educational opportunities while the second focuses on equality of condition as a means of improving the relative chances of children from disadvantaged origins. While there is little evidence that educational expansion in the 20th century led to a decreasing association between class origins and educational attainment in Sweden (Jonsson and Erikson, in press), our results show that expanding the educational system nevertheless helped to reduce the association between class origins and class destinations. It allowed more children to reach educational levels that led them to labor market segments where meritocratic selection was more prevalent and origin characteristics counted for less. The equalizing of educational chances in Sweden also led to increasing social fluidity. Moreover, although one can have compositional effects without equalization, equalization almost certainly implies expansion of the middle and/or higher levels of education, given that equalization is unlikely to occur through a reduction in educational participation by the middle classes. This means that promoting such equalization is likely to be a very effective strategy because it also advances compositional change. Our results provide perhaps the clearest example, within the social mobility literature, of how educational equalization can drive social fluidity: equalization implies expansion of the middle and/or higher levels of the educational system, and, if social origin has less impact in the labor market for those with such qualifications, compositional change accelerates the trend toward increasing social fluidity. It is of course possible that expansion also might lead to countervailing tendencies, with social origin reasserting itself in the labor market for graduates as their number increases (something that Vallet [2004, p. 142] reports for France), though we did not find any such development in our data.

Our results indicate that the trend toward educational equalization in Sweden—and thus also the trends in the compositional effect and in overall social fluidity—ground to a halt in those cohorts born around midcentury. If the situation had persisted among cohorts born after 1974 (and assuming the continued absence of any independent period influences on fluidity) the trend of equalization in period fluidity in Sweden would have been expected to come to a halt by 2020, as the Swedish workforce came to consist only of cohorts born after the middle of the 20th century. However, recent studies show that the impact of social origin on the transition to upper secondary education diminished further during the 1990s (Gustafsson, Andersson, and Hansen 2001),[32] which may, along with the rapid expansion in the provision of higher education in the same decade, allow the trend toward equalization in Sweden to continue.

Finally, we turn to the relevance of our results for other countries and for questions of policy. Sweden is well known in the mobility literature for its high level of social fluidity (see Breen and Luijkx 2004*b*, pp. 59, 72) and for its long period of gradual equalization. Nevertheless, the educational reforms introduced in Sweden during the 20th century, to which we referred earlier, are rather typical of many developed countries, especially in Europe. This raises two questions: Has the same trend of increasing social fluidity occurred in other countries? and Why is fluidity higher in Sweden than elsewhere? One difficulty in answering the first question is the dearth of analyses that adopt a cohort, rather than a period, perspective: even so, there is now a large body of evidence from period-based studies to suggest a widespread trend toward increasing social fluidity (Breen and Luijkx 2004*a*). Whether this can be attributed to the same causes as the Swedish case is not known, but recent research has demonstrated an equally widespread trend toward a weaker association between class origins and educational attainment (for a comparative analysis see Breen et al. [2005]; various single country studies are referred to in Breen and Jonsson [2005, p. 226]).

Sweden's high level of fluidity might be attributed to two main factors. On the one hand, equality of condition (especially with respect to income and the risk of prolonged unemployment) has been attained to a much greater level there than elsewhere, and this has played an important role in weakening the transmissibility of mobility assets between generations. On the other hand, Swedish employers appear to consider formal merits rather than characteristics related to the family of origin when employing those with higher education, particularly graduates (see fig. 5). This, in turn, may be because the tertiary educational system is relatively homogeneous, being free of fees and not displaying any marked differences in prestige between institutions (at least not for the cohorts we analyze). One implication, noted by Breen and Luijkx (2004*a*, p. 400), is that much more of the link between origins and destinations is mediated via education in Sweden than in other countries, and so reforms of the educational system, as well as the mere expansion of it, are likely to have a greater impact on social fluidity there than elsewhere. Thus, in Sweden, the organization of the educational system has been allied with efforts toward greater equality of condition. We might expect each of these to contribute to increasing openness, but, if the Swedish case is any guide, their combination seems to be particularly effective.

[32] We cannot observe this in our data because the individuals affected by it are too young to have been included in our samples.

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Adolescent First Sex and Subsequent Mental Health[1]

Ann M. Meier
University of Minnesota

The 1996 Welfare Reform Legislation and its reauthorization in 2002 included financial provisions for programs promoting sexual abstinence until marriage. Under this legislation, programs are encouraged to teach that nonmarital sex is likely to have harmful psychological effects. Life course concepts and identity theory suggest that sex may be consequential for the mental health of some adolescents. Using the National Longitudinal Study of Adolescent Health, this article investigates mental health consequences of adolescent sex. The analyses reveal important contingencies of the effect of first sex. Timing relative to age norms, romantic relationship factors, and gender interact to condition the effect of first sex on mental health. While some adolescents experience mental health decrements, the majority of those who had first sex did not. This finding highlights the importance of considering contingencies when investigating the effects of life events on mental health.

Until the mid-1990s, the average age at which young people began having sex had steadily decreased. Now, almost half of American adolescents report that they have had sex by the time they graduate from high school

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