“Intergenerational Occupational Mobility in Britain and the U.S. Since 1850”

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Abstract

The U.S. both tolerates more inequality than Europe and believes its economic mobility is greater than Europe’s. These attitudes and beliefs help account for differences in the magnitude of redistribution through taxation and social welfare spending. In fact, the U.S. and Europe had roughly equal rates of inter-generational occupational mobility in the late twentieth century. We extend this comparison into the late nineteenth century using longitudinal data on 23,000 nationally-representative British and U.S. fathers and sons. The U.S. was substantially more mobile than Britain through 1900, so in the experience of those who created the U.S. welfare state in the 1930s, the U.S. had indeed been “exceptional.” The margin by which U.S. mobility exceeded British mobility was erased by the 1950s, as U.S. mobility fell compared to its nineteenth century levels.

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> [W]e have really everything in common with America nowadays, except, of course, language. Oscar Wilde, *The Canterville Ghost* (1887).

The economies of Britain and the U.S. have had much in common over the two centuries since the American Revolution: their legal traditions and property rights systems, sources of labor, capital, and technology, political ties and alliances in two world wars, and – Wilde's quip notwithstanding – language and culture are the most obvious. One significant respect in which they have differed, however, is the progressivity of their taxation and the scale of their social welfare spending, at least through the late 1970s. Policies in the U.S. reflect a belief that high rates of economic mobility leave little need for substantial redistribution by the state. Public opinion surveys are consistent with these priorities and a belief in high rates of mobility: Americans are less concerned by inequality and are less willing to support redistribution than Europeans regardless of their position in the income distribution. (Alesina, Di Tella, and MacCulloch, 2001)

Since the 1970s, new large, nationally-representative longitudinal datasets for a variety of industrialized countries have made possible systematic cross-country mobility comparisons that call into question the assumptions regarding mobility that seem to underlie U.S. redistributive policies. The U.S. today exhibits no more income mobility or occupational mobility across generations than similarly developed countries (Solon, 2002; Solon, 1999; Erickson and Goldthorpe, 1992), though U.S. policies for the last 75 years have been predicated on American "exceptionalism" to the mobility patterns seen across a broad set of nations. Piketty (1995) provides a model of "dynastic learning" in which two economies can, as a result of differences in mobility in the past, settle upon and retain very different redistributive regimes even after their mobility patterns have converged.[2]

[2] Piketty contends that "The multiplicity of steady states explains at the same time why different countries can remain in different redistributive equilibria, although the underlying structural parameters of mobility are essentially the same. This is particularly likely if a country exhibited for some time in the past a significantly different experience of social mobility before joining the 'common' pattern. The 'canonical'

The question we address is whether we can identify, for Britain and the U.S., those historical differences in mobility, particularly intergenerational occupational mobility.

Commentators throughout the nineteenth century suggested that the U.S. was indeed "exceptional" in the occupational mobility experienced by its population (as well as in its geographic mobility). Using nationally-representative data for Britain and the U.S. that follows 23,000 pairs of fathers and sons from the beginning of the 1850s to the beginning of the 1880s, we offer the first detailed comparisons of the mobility regimes experienced by these two countries in the three generations before they constructed their respective welfare states. In the process, we also offer a new perspective on the very different histories of labor relations and political activity by workers in Britain and the U.S. that past scholars (e.g. Turner in the 1890s; Sombart in the early 1900s; Thernstrom in the 1970s) have attributed to different amounts of economic opportunity and mobility by individual workers. Can we actually observe sufficiently large differences to explain these differences in labor radicalism?

Britain was chosen as the country to which to compare the U.S. experience because of the availability of comparable data (described below). But this is also a particularly illuminating comparison because of the large number of characteristics these two economies have shared since the middle of the nineteenth century when U.S. industrialization got underway. Intergenerational occupational change was adopted as the metric for mobility for reasons of convenience as well: it is the only economic outcome that can be examined throughout the period since 1850. It is in some

application is the United States, whose nineteenth century mobility and class structure differed significantly from that of Europe before the two countries [sic] converged in the twentieth century." (p. 554) As we shall see below, the extent of the difference in mobility between the nineteenth century U.S. and the twentieth century U.S. is itself a subject of some controversy and one upon which we offer new evidence below.

ways superior to income as a measure of mobility, and in some ways inferior.[3] But it is what we have, and has already been the object of a great deal of research in sociology where methods to analyze mobility have evolved substantially since the 1960s.

I. Previous Research on Mobility in Britain and the U.S.

Our primary interest is in (1) assessing the differences in mobility between Britain and the U.S. in the second half of the nineteenth century; (2) comparing that difference to the difference observed by the 1970s; and (3) explicitly evaluating the change in mobility within the U.S. from the second half of the nineteenth century to the second half of the twentieth.[4] There has been until now a lack of appropriate data to undertake any of these tasks (though there has been considerable work comparing twentieth century mobility rates across a set of developed countries, including Britain and the U.S., in the absence of data adequate to task (1), it has not been possible to say how mobility differences among countries have changed over long periods of time). We briefly survey the existing literatures in these areas before proceeding to our own contribution.

The comparison between Britain and the U.S. in the nineteenth century has been marked by the boldest pronouncements and the weakest empirical evidence. Britain has been viewed, since the time of de Tocqueville and Marx, as a considerably more rigid system in which family background plays a much more significant role is determining current prospects than in the U.S.[5] These

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[3] Björklund and Jäntti (1999, pp. 15-19) summarize some of the relative merits of income and occupation for the measurement of intergenerational mobility, and discuss scenarios in which they provide very different results. McMurrer et al. (1997) offer a similar discussion of the relative advantages of different measures of intergenerational mobility.

[4] No explicit comparison for Britain between mobility in the second half of the nineteenth century and in the second half of the twentieth century is made because of data comparability issues discussed below.

[5] In the 1830s, de Tocqueville noted, “Among aristocratic peoples, families remain for centuries in the same condition and often in the same place. . . . Among democratic peoples [e.g. in the U.S.], new families

differences have been attributed to a number of factors – the frontier and the rapid growth of completely new cities in the U.S., the feudal tradition and guild and apprenticeship systems in Britain, and the wide availability of free, public education in the U.S. But there has been no consistent data with which these assertions could be directly tested. There are several studies that have looked at both British nineteenth century mobility and U.S. nineteenth century mobility in isolation.

For nineteenth century Britain, Miles (1993 and 1999) and Mitch (1993) have each used samples of marriage registrations from 1839 to 1914 to measure intergenerational occupational mobility.[6] At the time of registration, both bride and groom as well as bride's father and groom's father were required to list their occupation. From this information, Miles calculates that between 60 and 68 percent of grooms married between 1839 and 1894 were in the same occupational class as their fathers when the grooms married. (Miles, 1999, p. 29). Though his finding are in general quite similar, Mitch finds evidence for slightly more mobility – 61 percent of grooms married between 1869 and 1873 were in the same class as their father, 20 percent were higher, and 19 percent lower. The data used in both studies, however, are less than ideal.[7]

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continually spring from nowhere while others disappear to nowhere and all the rest change their complexion." Three decades later in the 1860s, Marx saw the U.S. as more open and fluid than the older European societies, with their "developed formation of classes." American classes, on the other hand, "have not yet become fixed but continually change and interchange their elements in constant flux." He related "this situation to the immature character of the American working-class movement." He characterized the U.S. as having "a continuous conversion of wage laborers into independent self-sustaining peasants. The position of wages laborer is for a very large part of the American people but a probational state, which they are sure to leave within a longer or a shorter term."

[6] Their samples were somewhat different. They both used marriage registries, but they used different (possibly overlapping) samples of registries.

[7] The marriage registry data include only couples married in Anglican churches, so toward the end of the nineteenth century, these samples are increasingly unrepresentative. By 1914, 42 percent of all marriages took place outside the Anglican church (Vincent, 1989, p. 281). Also, the occupations of the groom and his father are recorded at the time of the groom's marriage, so the father's and son's occupations are observed at

For the nineteenth century U.S., a large number of studies have been completed for specific communities in the U.S. that give us a rough sense of occupational mobility in the past. For example, among males who remained in Boston, from 37 to 40 percent of sons ended up in the same occupational categories as their fathers over the period 1840-89. (Thernstrom, 1973, p. 83) Though this might in itself seem a sufficient basis on which to conclude that the nineteenth century U.S. had greater intergenerational occupational mobility than nineteenth century Britain (total mobility – the fraction of sons found outside their fathers' occupational categories – was twice as great in Boston as in Britain), the data for Boston suffers, like that from Britain, from a number of shortcomings that prevent such simple comparisons.

The principal difficulty with historical estimates for the U.S. is that they were most often constructed by observing a single community over a period of decades. The only individuals whose occupational mobility could be observed were those who remained in the community. It would be surprising if the movers and stayers did not have systematically different patterns of occupational mobility, given the positive and often substantial costs of migration. Occupational mobility measured using marriage records suffers from the same shortcoming as the British data: sons' occupations are examined at different points in their careers than fathers' occupations. The new nineteenth century data used below for the U.S. (like that for Britain) is not limited to individuals who remained in a place for a decade or more and examines sons' and fathers' occupations at similar ages, presenting a more representative picture of mobility than has previously been available. Two additional difficulties apart from the inconsistencies in the collection of the data and biases

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different points in their life cycles, with the son being considerably younger than the father. If it were possible to observe the father's and son's occupations holding age constant, a different picture of intergenerational mobility might emerge. Specifically, we might expect to observe a greater likelihood of mobility as the son gained years and experience in the labor market.

introduced by the source materials are: (1) the possibility that differences between the British and U.S. occupational structures account for much of the difference in total mobility; and (2) the possibility that even in the absence of these differences in occupational distributions, the imprecision of the mobility measure employed would obscure more fundamental differences or similarities in mobility. The measures of mobility provided in our analysis overcome these difficulties.

One study offers a long-run perspective on intergenerational occupational mobility within Britain: Miles (1999) attempts to reconcile his findings of increasing fluidity over the nineteenth and early twentieth centuries with work by Erickson and Goldthorpe (1992), among others, who discern no trend in intergenerational mobility from the 1940s to the 1970s. Differences in the data for the two eras (Miles used marriage registers and Erickson and Goldthorpe relied on survey data with a retrospective question on the occupation of the respondent's father when the respondent was 14 years of age) diminish the reliability of this comparison.

Only two studies have attempted to assess how intergenerational mobility changed between the nineteenth and twentieth centuries in the U.S. In a re-analysis of several city-specific studies from the nineteenth century and together with the Occupational Change in a Generation (OCG) cohorts for the twentieth, Grusky (1987) concluded that there was significant immobility in the nineteenth century, with the non-manual/manual divide particularly difficult to cross, and an increase in intergenerational mobility from the nineteenth century to the twentieth century.

The work by Guest et al. (1989) is closest to the comparison between U.S. mobility in the nineteenth century and twentieth centuries carried out below. Comparing a sample of young males linked from the 1880 U.S. census to the 1900 U.S. census, they find little change from the last two decades of the nineteenth century to the end of the period covered by the second OCG cohort (1973). Their comparison is less than entirely apt, however. Their nineteenth century data excluded most interstate migrants, and the time between the observation of the fathers' and sons' occupations was in all cases greater (by as much as a factor of two) in the nineteenth century data than in the twentieth century data.[8]

The literature comparing twentieth century intergenerational mobility across developed countries is now voluminous.[9] The comparison between Britain and the U.S. undertaken by Kerckhoff et al. (1985), like almost all international comparisons involving these two countries, uses the Oxford Social Mobility Study (1972) for Britain and the second cohort of the OCG (1973) for the U.S. They find "considerably more overall inter-generational and career mobility in the United States, but . . . the major differences between the two societies are due to shifts in the distributions of kinds of occupations." (1985, p. 281). Erickson and Goldthorpe (1992) examine a broader set of countries, and likewise find the U.S. and Britain roughly similar in intergenerational mobility, after accounting for differences in the distributions of occupations across the two countries, as did Grusky and Hauser (1984) in analyzing a set of 16 countries including Britain and the U.S.[10] In income terms, Solon (2002) and Björklund and Jäntti (1999) find similarly high rates of income immobility across generations in Britain and the U.S., with both exhibiting considerably less mobility from fathers to sons than Canada, Finland, and Sweden.

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[8] In their nineteenth century data, the individual's father's occupation was observed in 1880, and the individual's own occupation was observed in 1900, twenty years later. In the two OCG cohorts, the individual's own occupation was observed in the survey year (1962 or 1973), but the father's occupation reported was that for the father when the respondent was 16 years of age. Guest et al. (1989) used males from the OCG who were 25-34 in the survey year, so they have between 9 (for 25 year olds) and 18 years (for 34 year olds) between the report of their father's occupation and the report of their own.

[9] Treiman and Ganzenboom (2000) provide a useful survey of the entire history of comparative research on occupational mobility, both within and across generations.

[10] Contrasting views are found in Wong (1990) who finds greater mobility in Britain than in the U.S., and Yamaguchi (1987) who finds mobility greater in the U.S. than in Britain.

II. The Data

We use a common methodology in constructing nineteenth century data to compare mobility between the U.S. and Britain. For both countries we link a sample of males from the 1850/1851 census to the census taken thirty years later in 1880/1881. Our choice of Britain as a comparison was dictated by the availability of sources making it possible to construct longitudinal data in exactly the same manner as for the U.S. For Britain we use information on approximately 13,000 males linked from the 1851 British census to the 1881 British census, and for the U.S. on nearly 10,000 males linked from the 1850 to the 1880 U.S. Federal Censuses. Details on the matching procedure, representativeness, and sensitivity tests are described in Appendix 1.

The only economic outcome available in the longitudinal data used here is self-reported occupation. We observe the father's occupation in 1850 (U.S.) or 1851 (Britain) and the son's occupation thirty years later. After collapsing hundreds of occupational titles into a reasonable set of categories it becomes possible to construct tables that describe the transitions from fathers' occupational categories to sons' occupational categories. We have used four categories (white collar, farmer, skilled and semi-skilled, and unskilled) to reduce the sparseness of the mobility tables, but where it has been possible to use a larger number of categories, the basic qualitative results reported below are unchanged.[11]

[11] "White Collar" is comprised of professional, technical, and kindred; managers, officials, and proprietors; clerical; and sales. "Farmer" is comprised of only farm owners and farm managers. "Skilled/Semiskilled" is comprised of craftsmen and operatives. "Unskilled" is comprised of service workers and laborers, including farm laborers. These categories are sufficiently broad and the boundaries between them are sufficiently well understood that we believe that movement among them represents a good approximation to the conventional understanding of "intergenerational mobility." Nonetheless, in comparing mobility across countries or over time, a reasonable concern is that these categories are not consistent, and that as important sub-divisions arise within them, ignoring those sub-divisions will lead to an understatement of mobility for the period or country where such distinct, new groupings have become prominent. For example, over the century and a half spanned by our inquiry, the white collar category has changed substantially in the U.S. as the fraction of the labor force in clerical and sales positions has grown. To account

For the twentieth century, we have employed the same data as others who have worked in this area: the Oxford Mobility Study for Britain and the OCG (1973 cohort) for the U.S.[12] In each the respondent's occupation at the time of the survey is taken as the son's occupation, and the occupation that the respondent reported his father to have had when the respondent was age 14 (Britain) or 16 (U.S.) is taken as the father's. To prevent differences in the impact of World War Two and the Great Depression from influencing the results, males age 31-37 (whose fathers' reported occupations would have been in 1949-1955) were used from the British data and males age 33-39 (whose fathers' reported occupations would have been in 1950-1956) were used from the U.S. data.[13] This yields a range of years between fathers' and sons' occupations of 17 to 23 years, and an average of roughly 20. This was done to ensure comparability with the U.S. data from the nineteenth century: though the direct nineteenth century comparison between Britain and the U.S. will use a thirty-year interval between fathers' and sons' occupations (a restriction dictated by the sources available for Britain), the U.S. sources also allowed the creation of two twenty-year samples (one

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for such changes (as well as the greater fraction of factory operatives in nineteenth century Britain and in the modern U.S. compared to the nineteenth century U.S.), we will employ up to six occupational groups where the data make this possible (separating high and low white collar workers, and splitting skilled and semiskilled blue collar workers).

[12] The Oxford Mobility Study (University of Oxford, 1978) for Britain is available at the U.K. Data Archive at the University of Essex as study number 1097. See http://www.dataarchive.ac.uk/. The original 1962 Occupational Change in a Generation study (Blau and Duncan, 1967) and its 1973 replication (Featherman and Hauser, 1975 and 1978) are available from the Inter-University Consortium for Political and Social Research as study number 6162. See http://www.icpsr.umich.edu/.

[13] In the 1973 OCG, sons who were 31-39 (41-49) in the survey year who reported their fathers' occupations when they themselves were 16 years of age would have been referring to the calendar years 1950-56 (1940-46). Similarly, in the 1972 Oxford Mobility Study, sons who were 31-37 (41-47) in the survey year who reported their fathers' occupations when they themselves were 14 years of age would have been referring to the calendar years 1949-55 (1939-55). Comparisons between the samples using males 43-49 (OCG) and 41-47 (Oxford Mobility Study) will be provided, but readers are cautioned against drawing strong conclusions from them as they may reflect differences in the experience of U.S. and British fathers during World War Two more than underlying differences in "normal" levels of intergenerational mobility.

with fathers observed in 1860 and sons in 1880, and one with fathers observed in 1880 and sons in 1900). These will be used for assessing change in mobility over time within the U.S.[14]

III. Measuring and Modeling Intergenerational Occupational Mobility

Intergenerational occupational mobility can be assessed through the analysis of simple two dimensional matrices, with categories for fathers' occupations arrayed across one dimension and categories for sons' occupations arrayed across the other. Comparing mobility across two places or times requires comparison of two matrices. Suppose fathers and sons can be found in either of two jobs.[15] A matrix that summarizes intergenerational mobility in location P has the form $P = \begin{bmatrix} p_{11} & p_{21} \\ p_{12} & p_{22} \end{bmatrix}$ with numbers of fathers in the two occupations (1 or 2) in columns and numbers of sons in these occupations in rows. The entry in the power left (p_{12}) is the number of sons of job 1 fathers who themselves obtained job 2. One simple measure of the overall mobility in P is the fraction of sons who end up in jobs different from those of their fathers: $M_P=(p_{12}+p_{21})/\ (p_{11}+p_{21}+p_{12}+p_{22})$.

Though this measure has the virtue of simplicity as a benchmark, it also has a shortcoming when mobility is compared across two matrices P and Q: it does not distinguish between differences in mobility (1) arising from differences across the matrices in the distributions of fathers' and sons' occupations (differences in what Hauser, 1980, labels "prevalence") and (2) arising from differences

[14] Two additional British sample with twenty year intervals (1861-1881 and 1881-1901) are compared both to U.S. data for 1860-80 and 1880-1900 and to the Oxford Mobility Study in Long and Ferrie (2006). The results presented there support the conclusions below: that in the nineteenth century, intergenerational occupational mobility was considerably more pronounced in the U.S. than in Britain, but that this gap was largely eliminated by the second half of the twentieth century.

[15] No ordering can be imposed on the occupations. When we turn to analysis of the nineteenth century data with four categories (white collar, farmer, skilled/semi-skilled, and unskilled), it is possible to rank unskilled last unambiguously, but it is not clear how to rank the others relative to unskilled. There are no good sources that would allow us to calculate average incomes by occupation. We thus require analysis techniques that rely not on the ordering of occupational categories but only on their labeling.

across the matrices in the association between father's and sons' jobs that may occur even if the distributions of fathers' and sons' occupations were identical in P and Q (differences in what Hauser, 1980, calls "interaction"). Consider $P = \begin{bmatrix} 3 & 1 \\ 2 & 2 \end{bmatrix}$ and $Q = \begin{bmatrix} 2 & 1 \\ 6 & 1 \end{bmatrix}$ for which $M_P = 3/8$ and $M_Q = 7/10$. The marginal frequencies differ, so it is not clear whether the difference in observed mobility M results from this difference or from something more fundamental such as differences between P and Q in the amount of human capital necessary to achieve job 1.

One way to proceed is to adjust one of the matrices so it has the same marginal frequencies as the other. Such a transformation, if achieved by multiplication of rows and columns by arbitrary constants, does not alter the underlying mobility embodied in the matrix. (Mosteller, 1968; Altham and Ferrie, 2007) If we multiply the first row of Q by 2 and then multiply the first column of the resulting matrix by ½, we produce a new matrix $Q' = \begin{bmatrix} 2 & 2 \\ 3 & 1 \end{bmatrix}$ with the same marginal frequencies as in matrix P, with an associated total mobility measure $M_{Q'} = 5/8$. We could then calculate the difference M_P - $M_{Q'}$ and be confident that the difference in mobility does not result from differences in the distributions of occupations between the two locations.

There still may be differences in mobility between P and Q, even after adjusting the marginal frequencies and finding that M_P - $M_{Q'} = 0$, however. The fundamental measure of association between rows and columns in a mobility table is the cross-product ratio, which for P is $p_{11}p_{22}/p_{12}p_{21}$ and can be rearranged to give $(p_{11}/p_{12})/(p_{21}/p_{22})$, the ratio of (1) the odds that sons of job 1 fathers get job 1 rather than job 2 to (2) the odds that sons of job 2 fathers get job 1 rather than job 2. If there is perfect mobility, the cross-product ratio would be one: sons of job 1 fathers would have no advantage in getting job1 relative to sons of job 2 fathers. The more the cross-product ratio exceeds one, the greater the relative advantage of having a job 1 father in getting job 1. The cross-product ratio for P is 3 and for Q is $1/3$ (as it is for Q'), so there is more underlying mobility in Q than in P.

For a table with more than two rows or columns, there are several cross-products ratios, so a summary measure of association should take account of the full set of them. One such measure has been suggested by Altham (1970): the sum of the squares of the differences between the logs of the cross-product ratios in tables P and Q. For two tables which each have r rows and s columns, it measures how far the association between rows and columns in table P departs from the association between rows and columns in table Q:

$$d(P,Q)=\left[\sum_{i=1}^{r}\sum_{j=1}^{s}\sum_{l=1}^{r}\sum_{m=1}^{s}\left|log\left(\frac{p_{ij}p_{lm}p_{im}p_{lj}}{p_{im}p_{lj}q_{ij}q_{lm}}\right)^{2}\right|\right]^{1/2}$$

The metric d(P,Q) tells us the distance between the row-column associations in tables P and Q.[16] A simple likelihood-ratio χ^2 statistic G^2 (Agresti, 2002, p. 140) with (r-1)(s-1) degrees of freedom can then be used to test whether the matrix Θ with elements $\theta_{ij}=\log(p_{ij}/q_{ij})$ is independent; if we can reject the null hypothesis that Θ is independent, we essentially accept the hypothesis that $d(P,Q)\neq 0$ so the degree of association between rows and columns differs between table P and table Q.

The statistic does not tell us which table has the stronger association, but that can be determined by calculating d(P,I) and d(Q,I), which use the same formula as d(P,Q) but replace one table with a matrix of ones. If d(P,Q)>0 and d(P,I)>d(Q,I), we conclude that mobility is greater in

[16] See Altham and Ferrie (2007) for a discussion of the distance measure and test statistic. As it obeys the triangle inequality, so $d(P,Q) + d(Q,I) \geq d(P,I)$, the metric d(P,Q) can be thought of as the distance between the row-column association in table P and the row-column association in table Q, while d(P,I) and d(Q,I) are the distances, respectively, between the row-column associations in tables P and Q and the row-column association in a table in which rows and columns are independent. This property of the Altham statistic – its interpretation as a *distance* measure – makes it possible to visualize how the row-column associations differ across various tables. For a set of N tables, the pair-wise distances among all the tables and the distance from each to a table with independent rows and columns are sufficient to allow us to display the positions of these tables relative to independence in a multidimensional space. The idea is the same as generating a map of cities in the U.S. knowing only the distances between each pair of cities and selecting an arbitrary point of reference.

table Q (i.e. mobility is closer in Q than in P to what we would observe under independence of rows and columns, in which the occupation of a father provides no information in predicting the occupation of his son). It is, of course, possible that in some circumstances d(P,Q) > 0 but d(P,I) $\approx$ d(Q,I), in which case we will say that tables P and Q have row-column associations that are equally distant from the row-column association observed under independence, but that tables P and Q differ in how they differ from independence (i.e. the odds ratios in table P that depart the most from independence are different from those that depart the most from independence in table Q).

Contingency tables are often dominated by elements along the main diagonal (which in the case of mobility captures immobility or occupational inheritance). It will prove useful to calculate an additional version of d(P,Q) that examines only the off-diagonal cells to see whether, conditional on occupational mobility occurring between fathers and sons, the resulting patterns of mobility are similar in P and Q. This new statistic will then test whether P and Q differ in their proximity to "quasi-independence." (Agresti, 2002, p. 426) For square contingency tables with r rows and columns, this additional statistic d^i(P,Q) will have the same properties as d(P,Q), but the likelihood ratio χ^2 statistic G^2 will have $[(r-1)^2-r]$ degrees of freedom.

Because it is a pure function of the odds ratios in tables P and Q, d(P,Q) is invariant to the multiplication of rows or columns in either table by arbitrary constants. As a result, d(P,Q) provides a measure of the difference in row-column association between two tables that abstracts from differences in marginal frequencies. Because $[d(P,Q)]^2$ is a simple sum of the squares of log odds ratio contrasts, it can be conveniently decomposed into its constituent elements: for an r × s table, there will be [r(r-1)/2][s(s-1)/2] odds ratios in d(P,Q) and it will be possible to calculate how much each contributes to $[d(P,Q)]^2$, in the process identifying the locations in P and Q where the differences between them are greatest.

In analyzing how mobility differs between two tables, we will then proceed in three steps:

1. calculate total mobility for each table as the ratio of the sum of the off-diagonal elements to the total number of observations in the table, and find the difference in total mobility between P and Q;

2. adjust one of the tables to have the same marginal frequencies as the other and re-calculate the difference in total mobility to eliminate the influence of differences in the distribution of occupations;

3. calculate d(P,Q), d^i(P,Q), d(P,I), and d(Q,I) and the likelihood ratio χ^2 statistics G^2; if d(P,Q) $\neq$ 0, calculate the full set of log odds ratio contrasts and identify those making the greatest contribution to $[d(P,Q)]^2$.

This differs from common practice in sociology, where the estimation of log-linear models has dominated the empirical analysis of mobility since the 1960s.[17] Log-linear analysis decomposes the influences on the log of each entry in a contingency table into a sum of effects for its row and column and an interaction between the row and column. Controlling for row and column effects eliminates the effect of the distribution of fathers' and sons' occupations on mobility. The remaining interaction between rows and columns captures the strength of the association between rows and columns which in turn measures mobility, though the coefficient on the interaction term has no meaning in itself as it is a component of a highly non-linear system.[18] In comparing mobility in two tables, the underlying question addressed is how well a particular pattern of mobility fits the different layers of the table, through comparisons of likelihood ratios. Attention is generally focused on the statistical significance of the difference in the fit of particular models across layers rather than on the magnitude of differences in row-column association. Simple comparisons of differences in the strength of the row-column association are not generally performed without the imposition of

[17] See Hauser (1980) and Hout (1983).

[18] Goodman (1970) suggests using the standard deviations of the log-linear model parameters in additive form as a measure of the strength of the row-column association in each layer. The approach adopted here instead has some advantages, described below, over this approach.

additional structure. For example, an analysis may have as its maintained hypothesis that all of the odds ratios in P differ in exactly the same degree from all of the odds ratios in Q, or that the odds ratios can be partitioned into sets that differ uniformly across the tables.

The measure of underlying mobility adopted here has several advantages over the more commonly employed measures of mobility derived from log-linear analysis: the measure used here (1) generates a simple, meaningful measure of the *distance* between the row and column association in P and the row and column association in Q that is conceptually straightforward and easy to visualize (see Figure 1 below);[19] (2) can be easily decomposed, allowing us to isolate the specific odds ratios that account for the largest part of difference between the association in P and the association in Q; (3) has a simple associated one-parameter test statistic that allows us to say whether the difference between the row-column association in P and the row-column association in Q is non-zero; and (4) answers a question ("does the row-column association in P differ from that in Q, and if so by how much and in which odds ratios?") that should be methodologically prior to the question addressed by more commonly employed measures of differences in row-column association based on log-linear analysis ("can we find a particular pattern of row-column association that is common to tables P and Q?"). For purposes of comparison, we will nonetheless provide measures based on log-linear analysis for the historical data.

IV. Britain vs. the U.S. in the Twentieth Century

Before turning to the nineteenth century, we assess the difference in mobility between Britain and the U.S. using the tools described in the previous section and males age 31-37 in 1972

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[19] Goodman and Hout (1998) provide a method to visualize differences in row-column association across tables for each log-odds ratio, but do not offer a summary measure for the entire table.

from the Oxford Mobility Study and white, native-born males age 33-39 in 1973 from the Occupational Change in a Generation survey. All cases in which the respondent reported a non-civilian occupation for himself or his father were excluded. Table 1 provides a cross-classification of son's occupation by father's occupation, and Table 2 provides summary measures of mobility for each panel in Table 1 and for differences in mobility between the panels.

According to the simple measure of total mobility M (Table 2, panel 1, column 1), young men in their thirties in 1972-73 were less likely in the U.S. than in Britain to find themselves in the occupations their fathers had in 1949-55. But this difference was largely a result of differences in the occupational structures of the two economies. If total mobility is measured for both countries using either the British (45.3 vs. 48.3) or U.S. (53.7 vs. 56.7) distributions of occupations, the gap in total mobility falls from 11.4 percentage points to 3 percentage points.[20] If Britain had the U.S. occupational distribution but the underlying association between rows and columns actually seen in Britain (panel 1, column 2, row 1), and the U.S. had the British occupational distribution but the underlying association between rows and columns actually seen in the U.S. (panel 1, column 2, row 2), the British (53.7 percent) would have actually had more total mobility than the U.S. (48.3 percent).

In both Britain and the U.S., an underlying association between fathers' and sons' occupations apart from that induced by differences in occupational distributions was present (for both, we can reject the null hypothesis that their association between rows and columns was the

[20] All of the underlying five-way mobility tables employed in the following analyses are contained in Appendix 3. To illustrate, Table A2-5 in Appendix 3 shows the British and U.S. mobility tables from Table 1 that result from applying the other country's marginal frequencies to each country's mobility table, using iterative proportional fitting. The M' entries in Column (2) of Table 2 were generated by calculating the percentage who end up off the main diagonal (i.e. in occupations different than their fathers) in Table A2-5 . For example, when the U.S. marginal frequencies are imposed on the British mobility table, 53.7 percent of British sons are off the main diagonal; when the British marginal frequencies are imposed on the U.S. mobility table, 48.3 percent of U.S. sons are off the main diagonal.

same as we would observe under independence). The difference between them in their degrees of association (Table 2, panel 1, column 7) is small in magnitude (7.9), and we cannot reject at any conventional significance level the null hypothesis that their associations are identical.[21] This is not solely the result of strong similarities in the tendency of sons to inherit their fathers' occupations, as we cannot reject the null hypothesis that association is identical even if we focus only on the off-diagonal elements in each table (panel 1, column 9). These results confirm the findings of Erickson and Goldthorpe (1992) and Kerckhoff et al. (1985) that, after accounting for differences in their occupational distributions, Britain and the U.S. exhibited similar intergenerational occupational mobility in the third quarter of the twentieth century.

The white collar category is quite broad in both countries in the twentieth century, spanning professionals and managers as well as clerical and sales workers. If substantially more mobility occurs within this category in one country than in another, mobility comparisons based on only four categories may be misleading. To remedy this, we divided "white collar" into "high white collar" (professional, technical, and kindred; managers, officials, and proprietors) and "low white collar" (clerical and sales) and calculated new Altham statistics for Britain (P) and the U.S. (Q). The magnitudes of the Altham statistics rose somewhat for both countries (d(P,I)=37.50, d(Q,I)=31.06), as did the magnitude of the difference between them in row-column association (d(P,Q)=17.81), but

[21] The comparison between Britain and the U.S. is substantially different if males age 41-47 (Britain) and 43-49 (U.S.) whose fathers' occupations are reported during World War Two are used instead of 31-37 and 33-39 year old males: the Altham statistics for Britain (d(P,I)=30.02), for the U.S. (d(Q,I)=17.98), and for the difference in row-column association (d(P,Q)=15.18, G^2=41.89, $p<0.001$) reveal a great deal more mobility in the U.S., and a large difference between the row-column associations in the two countries that is not apparent when younger males whose father were observed after World War Two are used. This could reflect either the influence of differences between the two countries in fathers' occupations during the war years (a cohort effect) or greater occupational mobility in the U.S. than in Britain during the additional ten years between fathers' and sons' occupations (a time effect) that the 41-47 & 43-49 year olds' data captures.

it was again not possible to reject the null hypothesis that the true difference was zero (prob[H_o:d(P.Q) = 0] = 0.88).

V. Britain vs. the U.S. in the Nineteenth Century

How different were Britain and the U.S. in intergenerational occupational mobility a century earlier? Table 3 presents the cross-classification of sons' and father's occupations using our new data linking fathers in 1850 (U.S.) or 1851 (Britain) and sons in 1880 (U.S.) or 1881 (Britain). Summary mobility measures again appear in Table 2. The simplest measure of mobility shows the U.S. with a slight advantage (inheritance of the father's occupation was 2.8 percentage points less likely in the U.S.), but substantial differences in occupational distributions obscure much larger differences. If the U.S. had Britain's occupational distribution, the U.S. advantage in total mobility would have been 5.3 percentage points; if Britain had the U.S. distribution, the U.S. advantage would have been 9.9 percentage points. Finally, if Britain and the U.S. had swapped occupational distributions and retained their underlying association between fathers' and sons' occupations, the U.S. advantage would have been 12.4 percentage points.

These simple comparisons suggest that more fundamental measures of association between fathers' and sons' occupations would reveal a weaker association (and greater mobility) in the U.S. The second set of summary mobility measures in Table 2 shows that this was indeed the case: though the association between fathers' and sons' occupations differed from independence in Britain and the U.S., the magnitude of the association was twice as great in Britain (22.7) as in the U.S. (11.9) (compare Table 2, panel 2, columns 3 and 5). We can safely reject the null hypothesis that the difference between them in their associations was actually zero. The point estimate for d(P,Q) was 13.2, indicating a difference in mobility after controlling for occupational distributions that was not only statistically significant but also large in magnitude, compared to d(P,I) and d(Q,I).[22]

Table 4 disaggregates $[d(P,Q)]^2$ into its components, and calculates the contribution of each of the $[d_i(P,Q)]^2$ that account for three quarters of $[d(P,Q)]^2$ to the total. G^2 is also reported for each contrast, as well as the underlying odds ratios from P and Q. For example, the first entry is the relative advantage in entering farming rather than unskilled work from having a farmer father rather than an unskilled father. In Britain, sons of farmers were 49 times more likely to enter farming rather than unskilled work than were the sons of unskilled workers. In the U.S., the ratio was only 4.5 to one, so the advantage of having a farm father rather than an unskilled father in making this move (into farming rather than unskilled work) was 11 times greater in Britain than in the U.S. This odds ratio contrast alone accounts for nearly 13 percent of the difference between the association in P and the association in Q. Of the nine odds ratios that account for 75 percent of the difference in association between P and Q, six display a smaller disadvantage in the U.S. in entering farming rather than another occupation for the sons of non-farmers, indicating that an important source of greater intergenerational mobility in the U.S. than in Britain was an easier path to farm operation from outside agriculture, regardless of the distribution of occupations for fathers and sons. But the importance of farming by no means exhausts the sources of higher mobility in the U.S. For example, in Britain, white collar sons had a 20 to one advantage in entering white collar rather than unskilled jobs compared to the sons of unskilled workers; in the U.S., their advantage was only 4 to one, a fifth of the advantage in making this transition conveyed in Britain by having a white collar father.

[22] If we split white collar into high and low white collar groups, the Altham statistics reveal the same stark differences between Britain (P) and the U.S. (Q): d(P,I)=41.17, d(Q,I)=16.98, d(P,Q)=29.58, G^2=136.98, $p<0.001$. If we use six categories (splitting both high and low white collar and skilled and semiskilled), the difference between Britain (P) and the U.S. (Q) remains: d(P,I)=59.18, d(Q,I)=28.40, d(P,Q)=43.90, G^2=185.85, $p<0.001$.

Not only is overall mobility greater in the U.S., but upward mobility also exceeds that in Britain. Without a comparable scheme of fully ranked occupational categories for both countries, a complete analysis of upward and downward mobility is impossible. However, some conclusions follow from innocuous assumptions. Assuming that unskilled occupations are less desirable than all others, Table 3 indicates that in the U.S. 81.4 percent of all sons of unskilled laborers moved up into other occupations, while only 54.3 percent of unskilled British sons experienced upward mobility; if the British marginal distribution of occupations is imposed on the U.S. mobility table, the U.S. advantage is narrowed but not eliminated (upward mobility in the U.S. falls to 61.8 percent, compared to 53.3 percent in Britain), while if the U.S. marginal distribution of occupations is imposed on the British mobility table, the British disadvantage is narrowed slightly but remains large (upward mobility in Britain rises to 61.2 percent, compared to 81.4 percent in the U.S.). Downward mobility in the U.S. was lower than in Britain: 8.7 percent moved into unskilled labor in the U.S. versus 14.0 percent in Britain, though this difference is reversed if either the British or U.S. marginal distributions are used for both countries. Thus, the U.S. was not only a less static labor market than Britain (as the Altham statistics reveal), but also a labor market with (1) better prospects for upward movement even after accounting for differences between its occupational structure and Britain's, and (2) less downward mobility than in Britain, though downward mobility would have been slightly greater in the U.S. than in Britain if the two countries had the same occupational distributions.

VI. Nineteenth Century vs. Twentieth Century Mobility in the U.S.

The difference in mobility between Britain and the U.S. in the nineteenth century was substantial, both before and after taking account of differences in their distributions of occupations. We have already seen that Britain and the U.S. were indistinguishable in terms of intergenerational occupational mobility in the third quarter of the twentieth century, after taking account of their occupational distributions. How was this convergence in underlying mobility achieved? Did U.S. mobility fall or did British mobility rise to U.S. levels? We cannot directly assess the change over time in British mobility in the absence of nineteenth century longitudinal data that span twenty years, unless we were to include the Great Depression. For the U.S. however, we have samples that span 1860-80 and 1880-1900 that are identical in their construction to the 1850-80 sample we used in the comparison to Britain 1851-81. Males age 33-39 at the end of the 1860-80 and 1880-1900 U.S. samples can be compared to males age 33-39 in the 1973 cohort of the OCG. These samples then both span either exactly 20 years between fathers' and sons occupations (1860 to 1880 and 1880 to 1900) or an average of 20 years between fathers' and sons' occupations (1949-55 to 1973). Table 5 presents the cross-classification of fathers' and sons' occupations for the 1860-80 data, which are compared to the OCG data from the lower panel of Table 1. Summaries of the comparison between them appear in the third set of contrasts in Table 2.

Total mobility shows a 6.1 percentage point advantage for the modern data, but when it is calculated for both tables using common marginal frequencies, the nineteenth century table has higher total mobility, by from one (using the 1860-80 frequencies) to 6.9 percentage points (using the 1973 frequencies). If the marginal frequencies are swapped but the underlying associations are left unchanged, the nineteenth century U.S. had a total mobility rate 1.3 times greater than that in the 1949-73 period. The more fundamental measure of mobility, d(P,Q), also shows greater mobility (i.e. a weaker association between fathers' and sons' occupations) in the nineteenth century than in the twentieth: we can safely reject the null hypothesis that the associations are equal (G^2=46.7 on 9 degrees of freedom, probability(H_0: same association)$<$ 0.0001), and the difference d(P,Q) is large in magnitude.[23] We cannot, however, reject the hypothesis that the associations are identical when the diagonal elements in P and Q are excluded, suggesting that change in the likelihood of direct inheritance of the father's occupational status by the son was the greatest difference between these eras, rather than more subtle change in the structure of association between one generation's occupation and that of the next.[24]

Table 6 decomposes the elements of d(P,Q) into those that account for three quarters of the difference between mobility in the nineteenth century and mobility in the twentieth. The single greatest difference – making up nearly 15 percent of the difference between the association in the nineteenth century and the association in the twentieth – is in the upper left four cells of the contingency table. In the nineteenth century, getting a white collar job rather than a farm job was 11 times more likely for the son of a white collar worker than for the son of a farmer; by the twentieth century, the advantage of white collar sons had grown nearly eight-fold relative to farm sons in getting white collar jobs rather than farm jobs. The second and third contrasts in Table 6 show swings in the odds ratios of similar magnitude from the nineteenth to the twentieth centuries (the advantage of farm sons relative to skilled and semiskilled sons in getting (1) white collar rather than farm jobs, and (2) farm jobs rather than unskilled jobs). Of the seven substantial differences between the nineteenth and twentieth centuries, three provide evidence of greater difficulty entering white collar jobs (for the sons of farmers relative to sons of white collar workers, for the sons of skilled workers relative to sons of farmers, and for sons of unskilled workers relative to sons of farmers).

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[23] This is in marked contrast to the findings of Hauser et al. (1975) who found no trend toward an increase in the association between fathers' and sons' occupations in the U.S. over the twentieth century. Though part of this difference may result from different methodologies, we suspect that most is the result of fundamental changes in the U.S. economy explored in the next section.

[24] When five occupational categories are used rather than four, the greater mobility for the nineteenth century (P) compared to the twentieth (Q) persists: d(P,I)=21.90, d(Q,I)=31.06, d(P,Q)=16.71, probability(H_0: d(P,Q)=0)<0.001.

The difference between nineteenth and twentieth century mobility persists into the last two decades of the nineteenth century. If the 1880-1900 sample is used (P) and compared to the 1973 OCG cohort (Q), substantially more mobility is again observed in the historical data than in the more recent past. Contrast 4 in Table 2 shows these results: total mobility was greater in the past if the nineteenth century occupational distributions are used or if the occupational distributions are swapped and each period retains its actual association between fathers' and sons' occupations. The unadjusted total mobility and total mobility using the twentieth century frequencies, however, favor the more recent data. But the underlying association measured by d(P,I), d(Q,I), and d(P,Q) was substantially greater in the past than more recently. We can safely reject the hypothesis that the association was identical (G^2=36.7 on 9 degrees of freedom, probability(H_0: same association)< 0.0001). Even in the last two decades of the nineteenth century, mobility was greater than in the 1949-73 period, a difference that was both large in substance and statistically significant.[25] Ferrie (2005, pp. 206-208) reports Altham statistics comparing mobility from three intervals in the nineteenth and early twentieth centuries (1860-80, 1880-1900, and 1900-20) to mobility from three intervals in the second half of the twentieth century (the 1973 OCG, the General Social Survey for 1977-90 , and the National Longitudinal Survey of Youth 1979 cohort). All six samples span roughly twenty years from the report of the father's occupation to the report of the son's occupation. After calculating d(P,I) for each table and calculating d(P,Q) for each pair of tables, multidimensional scaling (Davison, 1983) can be employed to locate each table's mobility in a two-dimensional space relative to an arbitrarily located origin representing independence, as in Figure 1.

[25] When five occupational categories are used rather than four, the greater mobility for the nineteenth century (P) compared to the twentieth (Q) again persists: d(P,I)=26.41, d(Q,I)=31.06, d(P,Q)=18.10, probability(H_0: d(P,Q)=0)<0.001.

High mobility in the nineteenth century U.S. was thus not principally a consequence of the enormous turnover in the U.S. labor force occasioned by the death of a substantial fraction of the working-age male population in the Civil War, or of the presence of an expanding agricultural frontier – the frontier was already "closed" by 1890, according to the U.S. Census Office.[26] It is also not the result of some peculiarity of the OCG data used for the twentieth century, as similar results are obtained when other modern surveys are employed.[27]

VII. Economic Sources of Higher 19th Century Mobility in the U.S. Than in Britain and of Declining Mobility in the U.S. During the Twentieth Century

The U.S. was considerably more mobile than Britain in the nineteenth century and roughly similar in mobility in the twentieth. At least some of this convergence occurred because of declining mobility in the U.S. (as opposed to improved mobility in Britain). Unfortunately, the foregoing analysis sheds little light on the sources of either the U.S. advantage in the nineteenth century or its relative decline in mobility from the nineteenth century to the twentieth. Because the metric for the distance in association used here focuses on odds ratios, it is not even possible to say for certain whether the observed differences result from differences in the numerators, in the denominators, or in both.[28]

[26] The Superintendent of the Census reported in 1890 that "Up to and including 1880 the country had a frontier of settlement, but at present the unsettled area has been so broken into by isolated bodies of settlement that there can hardly be said to be a frontier line. In the discussion of its extent, its westward movement, etc., it can not, therefore, any longer have a place in the census reports." (U.S. Census Office, 1891).

[27] In each case, the late nineteenth and early twentieth century displays mobility that is greater in magnitude than the late twentieth century, differences that are in every case highly statistically significant. By contrast, differences within the twentieth century are small in magnitude and not statistically significant.

[28] For example, the third contrast in Table 4 and the first in Table 6 – [(WW)/(WF)]/[(FW)/(FF)] – is the ratio of the odds of white collar sons entering white collar jobs rather than farming to the odds of farm sons entering white collar jobs rather than farming. It is greater in nineteenth century Britain and the

Are the differences we have observed (between the mid-nineteenth century U.S. and either mid-nineteenth century Britain or the mid-twentieth century U.S.) simply a reflection of differences in the size of the farm sector, i.e. so many more farmers in the mid-nineteenth century U.S. and as a result much movement out of farming and more "mobility?" The measure of mobility we have used already adjusts for differences in the size of the occupation groups, however. If the mid-nineteenth century U.S. farm sector is driving the results, it must be more than the difference in the sector's sheer size generating differences with Britain at the same time or the U.S. 100 years later. There must be a selectivity effect as well.

Consider nineteenth century Britain versus the nineteenth century U.S.: Britain has already seen almost all of its flight from agriculture by 1851 (Figure 2), so the sons of farmer fathers are already selected for remaining in farming (all the sons who were more loosely attached to the sector have already left by 1851). At the same time, the sons of non-farm fathers are already selected for remaining outside farming (all the sons eager to enter farming have already done so). In the U.S., this weeding out process has not taken place in the nineteenth century, so the U.S. has more mobility both out of and into farming that gets added onto whatever the underlying amount of mobility would be otherwise.[29] At least some of the high mobility in the nineteenth century U.S. may then result from it being at an earlier stage of development than nineteenth century Britain or the

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twentieth century U.S. than in the nineteenth century U.S. But is this because the nineteenth century U.S. has (1) greater ease for farm sons in attaining white collar jobs in the nineteenth century U.S. (FW ↑) , (2) a weaker attachment to farming among farm sons (FF ↓), (3) easier entry by white collar sons into farming (WF ↑), or (4) a weaker attachment to white collar jobs among sons of white collar workers (WW ↓)? Or does it result from some combination of these?

[29] Alternatively, as a referee has pointed out, we can think not of the survival of sons as farmers but rather of the survival of farms as the mechanism leading to some change in the "quality" of movers out of farming over time: as more farms fail or are sold off in a process of consolidation, more of the less-able farm sons are entering non-farm occupations. Of course, for the story to work for the nineteenth century versus twentieth century U.S. case, there must be no increase in the selectivity of movement out of or into farming even as late as 1900 when farmers as a fraction of the labor force had fallen to 20% from 45% in 1850.

twentieth century U.S., so its farm sector was relatively larger and selective exit from farming and entry into farming were less apparent than in Britain at the same time or in the U.S. a century later. As late as 1850, 45 percent of U.S. workers were still in farming, compared to 4 percent in Britain in 1880 and 7 percent in the U.S. in 1950.

To get at the amount of mobility after taking out the effect of selective mobility out of or into farming, we re-ran the analyses after removing the cell Farm [father]-Farm [son] and the cells White Collar-Farm, Skilled/Semiskilled-Farm, and Unskilled-Farm. This is preferable to leaving out the farm sector altogether, as it still allows us to include the mobility of sons of farmers conditional on their departure from farming. The results are

P=Britain 1851-81, Q=U.S. 1850-80

d(P,I)=12.79 (prob $<$ 0.0001)
d(Q,I)= 8.81 (prob $<$ 0.0001)
d(P,Q)=7.42 (prob $<$ 0.009).

Even if we ignore the Farm-Farm immobility difference and ignore differences in entry into farming, then, the differences in mobility still go in the same direction (the nineteenth century U.S. is markedly more mobile than nineteenth century Britain). For the U.S. over time, the same is true as well, though the remaining magnitudes are smaller:[30]

P=U.S. 1860-80, Q=U.S. 1953-73

d(P,I)=8.00 (prob $<$ 0.0001)
d(Q,I)=8.15 (prob $<$ 0.0001)
d(P,Q)=3.35 (prob $<$ 0.078).

[30] When five occupational categories are used rather than four, the nineteenth century results for Britain (P) and the U.S. (Q) are: d(P,I)=29.95, d(Q,I)=13.09, d(P,Q)=23.10, probability(H_0: d(P,Q)=0)$<$0.001. For the comparison between the 1860-80 U.S. (P) and the 1973 OCG, the results are: d(P,I)=16.97, d(Q,I)=14.41, d(P,Q)=9.26, probability(H_0: d(P,Q)=0)$<$0.03.

Simple differences in the selectivity of exit from or entry into farming, in any case, cannot explain the majority of the contrasts in Tables 4 and 6.[31] Other features of the nineteenth century U.S. economy may help explain its uniquely high rates of mobility. A useful starting point for analyzing the economic causes of differences in mobility across times or places is the formulation of Becker and Tomes (1986) who model intergenerational mobility as an outcome generated by the endowments transmitted directly from parents to children, and by investments made by parents faced with several investment opportunities and possibly constrained by the operation of capital markets from making the efficient level of investment in their children.

As Grawe and Mulligan (2002) demonstrate, this simple model provides some testable implications regarding spatial or temporal differences in earnings mobility.[32] Ignoring capital constraints (generated by the inability of parents to borrow against the future labor earnings of children), intergenerational earnings mobility will be higher when the ease with which ability is transferred to children is reduced. Han and Mulligan (2001, p. 225) show that earnings mobility is also greater when ability displays less variance. Finally, if parents are constrained in the credit market, they will invest less in their children, whose earnings will more closely reflect ability, reducing

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[31] For example, the third contrast in Table 4 does not involve farmers. The fifth contrast in Table 4 (which is also the seventh in Table 6) – [(WF)/(WS)]/[(FF)/(FS)] – can be higher in nineteenth century Britain and the twentieth century U.S. than in the nineteenth century U.S. because of selectivity only if exit from farming by farm sons exceeded entry by white collar sons into farming by a greater margin in the nineteenth century U.S. than in either nineteenth century Britain or the twentieth century U.S. If this was not the case, differential entry into skilled and semi-skilled jobs by white collar and farm sons must also account for some of the greater size of this contrast for the nineteenth century U.S. To see this, re-write the contrast as [(WF)/(FF)]/[(WS)/(FS)].

[32] Though these implications relate to earnings mobility, it is straightforward to map them into occupational mobility. If there are two possible jobs and investment (by parents or the state) both raises (1) the odds that sons of job 1 fathers will get job 1 rather than job 2 and (2) the odds that sons of job 2 fathers will get job 1 rather than job 2, but (2) rises by more than (1), the odds ratio will fall, indicating greater mobility. The only additional assumption necessary for the implications discussed by Grawe and Mulligan (2002) to apply to occupational mobility as well is that all workers qualified for job 1 can obtain job 1.

mobility. Where credit markets function well, or where wealth is greater so fewer parents find the capital constraint binding, mobility will be greater than where credit markets do not function well, or where most parents find themselves constrained by low wealth.

We have no direct evidence on how easily abilities were transmitted from parents to children in the nineteenth century in Britain and the U.S. or in the twentieth century U.S. But we can suppose that the greater heterogeneity in the origins of the U.S. population compared to the British population in the nineteenth century corresponded to greater variance in abilities in the U.S., a force working to undermine the U.S. advantage in occupational mobility relative to Britain at this time. Though we cannot test directly for the role of credit market constraints in generating the advantage enjoyed by the U.S. relative to Britain in the nineteenth century and the decline in relative mobility by the twentieth century, it is possible to see how important such impediments to investment may have been in generating the level of mobility seen within the nineteenth century U.S.

The 1860-80 sample provides information from the 1860 population census on the total wealth owned by the household (the sum of real estate and personal estate). This makes is possible to assess the role of credit constraints by examining whether mobility differs systematically by household wealth, an indicator of the probability that a household is credit constrained. Following Mazumder (2001), the 1860-80 sample was divided in half: high total wealth families (wealth $\geq$ median wealth=\$1,400) and low total wealth families (wealth < median). Intergenerational occupational mobility matrices were then constructed (P=high wealth, Q=low wealth), and the underlying association between fathers' and sons' occupations was calculated along with the difference in association between P and Q. For both types of households, mobility was different from that expected under independence though slightly greater in high wealth households (d(P,I)=12.65, d(Q,I)=13.25, while the G^2 statistics for both are significant at 0.01. Of greater interest is the difference in association between P and Q: d(P,Q)=6.14 (G^2=21.2, prob=0.01), confirming that mobility in the 1860s and 1870s was in fact greater among high wealth households than among low wealth households.

Grawe and Mulligan (2002, p. 51) suggest that "one way to investigate [the role of credit market imperfections] is through analysis of cross-country evidence on whether countries with greater public provision of human capital experience greater intergenerational mobility." The U.S. provided considerably more public education than Britain in the middle of the nineteenth century: 68.1 percent of 5-14 year olds were enrolled in primary school in the U.S. in 1850 compared to only 49.8 percent in England and Wales (Lindert, 2004, p. 92). The U.S. educational system in the second half of the nineteenth century, though less extensive at the secondary and post-secondary levels than European systems was considerably more egalitarian (Goldin, 1999, p. 2). To the extent that intergenerational mobility is greater where fewer parents are constrained, superior mobility in the U.S. may well have been a consequence of its educational system, which provided a public alternative to a private education that was outside the reach of many families.

The importance of free, public education provides a less satisfactory explanation for the trend in mobility over time within the U.S., though: while enrollment rates, graduation rates, and spending have increased dramatically since the nineteenth century (Goldin, 1999, pp. 52-68), intergenerational occupational mobility has nonetheless fallen. Though the educational requirements to advance in occupational status (or to avoid a decline in occupational status) may have risen more rapidly than the aggregate statistics on the provision of education, there is no evidence with which to confront this conjecture.[33]

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[33] Becker and Tomes (1986, p. S31) suggest with some justification that capital constraints construed more generally fell from the nineteenth century to the twentieth in the U.S. But this, too, runs counter to the trend of decreasing mobility from the nineteenth century to the twentieth. The model's prediction that larger

A potentially more promising avenue for explaining both the U.S. advantage in mobility compared to Britain within the nineteenth century and the decline in relative U.S. mobility since the nineteenth century is to consider characteristics of the U.S. economy that correspond to both of these contrasts. The most obvious candidate is residential mobility. Migration can be seen as an investment (Schultz 1961, Becker 1964). These investments made by families can then improve a child's chances for occupational mobility in the same way as a family's investment in the human capital of its children can promote mobility in the Becker and Tomes (1986) model. If a family is credit-constrained and unable to undertake such investments, or unaware of such investment opportunities either because of poor information or because some opportunities did not yet exist when the child was young, the child may be able to make the investment instead, by migrating later in life.

In Britain, 27 percent of sons were observed in different counties in 1851 and 1881, while in the U.S. 62 percent of sons were in different counties in 1850 and 1880. Sons in the U.S. were also more likely to cross a state boundary over these three decades than British sons were to cross a county boundary.[34] Though we lack comparable data on mobility over a span of thirty years for the twentieth century U.S., the National Longitudinal Survey (NLS) cohorts of Older Men and Young Men provide a comparison over ten years, the shortest span we can observe in the nineteenth century linked files. Between 1870 and 1880, 55 percent of young (20-29 years) white, native-born males changed county and 30 percent changed state; between 1971 and 1981, only 42 percent of

family size will be associated with lower investment per child and lower mobility (if fertility is exogenous) is another force working against the finding of relatively greater mobility in the nineteenth century U.S. than in the twentieth: the total fertility rate in the U.S. fell from 5.42 in 1850 to 2.98 in 1950.

[34] The 52 traditional counties of England and Wales are 1,123 $mi.^2$ in area on average; the 3,112 county units in the continental U.S. are 1,003 $mi.^2$ in area on average.

otherwise identical males changed county, while only 22 percent changed state. Among older men (45-59 years) the declines in both inter-county mobility (from 35 percent 1870-80 to 16 percent 1966-76) and inter-state mobility (from 22 percent 1870-80 to 8 percent 1966-76) were even more dramatic.

Though mid-nineteenth century Britain was a considerably more compact economy with an extensive transportation network, residential mobility was greater in the mid-nineteenth century U.S. Though transportation costs fell dramatically over the century from 1870 to 1970 within in the U.S., residential mobility at the county and state levels fell from the 1870s to the 1970s in the U.S. These comparisons imply that the rate of return on geographic mobility must have been greater in the U.S. in the second half of the nineteenth century than in either late nineteenth century Britain or the late twentieth century U.S.[35]

The late nineteenth century U.S. was remarkable in an additional respect: it also probably had a greater distribution in the returns to migration across its physical geography. One force promoting differences in the rate of return to migration across locations was differences in the economic activities being undertaken in different places. Using data on employment by one-digit SIC code sectors, Kim finds that "Regional specialization in the overall economy rose through the early nineteenth century, leveled off between the late nineteenth and the early twentieth centuries, and then fell precipitously through most of the twentieth century." (Kim 1998, p. 667).

These differences in the geographic concentration helped generate large differences in the rates of growth of urban places across the country, as cities and towns arose to meet region-specific

[35] The high rates of return to geographic mobility in the late nineteenth century was not a direct result of the existence of large, internal frontier; however: (1) both intergenerational occupational mobility and geographic mobility were as high through 1910, twenty years after the frontier's demise; and (2) the vast majority of internal migrants never went to the western frontier.

demand. Figure 3 shows the standard deviation in population growth rates for the largest 100 urban places in the U.S. since 1840, with separate tabulations for all places ever in the top 100, places that were in the top 100 for at least 5 decades, and places that were in the top 100 for at least 10 decades. This simple measure of how differently cities were growing falls through the second half of the nineteenth century and remains low through the end of the twentieth century.[36]

At the regional level, the absence of large differences in wages indicates that the U.S. labor market, at least within the North, was well-integrated by the middle of the nineteenth century. (Margo, 2000; Rosenbloom, 1996). There remains the possibility, however, that differences across smaller units of geography than regions may have continued to present opportunities for "locational arbitrage" – migrating from a place with poor prospects for occupational mobility to one with better prospects – that could be exploited as avenues to occupational mobility through the 1930s. Higher levels of regional specialization and the presence of more urban places growing at widely divergent rates in the late nineteenth century U.S. may have provided greater opportunities for such locational arbitrage than in late nineteenth century Britain or the late twentieth century U.S.[37]

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[36] Late nineteenth century Chicago is an example of an urban place that arose to provide services to the Midwest's growing farm sector, grew more rapidly than other U.S. cities, and emerged as a site of extraordinary economic opportunity. From 1850 to 1870, its population grew by a factor of ten, and then grew by nearly a factor of three from 1870 to 1890. Galenson (1991, p. 597) characterized late nineteenth century Chicago as "a place of unusually great economic opportunity."

[37] Though Blau and Duncan (1967, pp. 252-253) find no link between geographic mobility and intergenerational occupational mobility other than that arising from differences in the marginal distributions of occupations across locations, their data (from the original Occupational Changes in a Generation 1962 survey), as well as that from the OCG 1973 replication, come predominantly from a period well after the decline in regional specialization and the homogenization in urban growth rates that are suggested here to underlie a link between geographic and occupational mobility in the late nineteenth and early twentieth century U.S.

VIII. Conclusion

Though the U.S. exhibited no more intergenerational occupational mobility in the late twentieth century than similarly developed countries, a widely-shared belief that the U.S. is a place of unusually easy mobility has consistently guided public policy and shaped debate regarding the appropriate functions of the government in promoting social welfare from the 1930s to the present. Using new longitudinal data for the nineteenth century, we have identified an era when the U.S. mobility experience was indeed exceptional: even after controlling for differences in their occupational structures, the U.S. had substantially more occupational mobility across generations than either Britain in the three decades after 1850 or the modern U.S. Though it remains to be seen exactly why nineteenth century U.S. mobility exceeded that in both nineteenth century Britain and the twentieth century U.S., and when the transition to a lower mobility regime in the U.S. took place, high U.S. intergenerational occupational mobility corresponded to high rates of residential mobility. A fall in U.S. residential mobility after 1910 as economic activity across locations became more homogenous may have reduced the ability of families and individuals to “invest through migration” and foster occupational mobility across generations.

Appendix 1: Linked Census Data

The population censuses of Britain and the U.S. are generally regarded to be the best sources of individual-level, nationally representative data from the nineteenth century for those countries. However, the cross-sectional censuses do not provide the continuity over time needed to study issues of mobility at the level of the individual. Two new sources have made it possible to create the necessary continuity from the British and U.S. historical census records. The Genealogical Society of Utah in conjunction with the Federation of Family History Societies has computerized the individual-level records from the enumerators' books of the 1881 Census of the Population of England, Wales, and Scotland and from the 1880 U.S. Federal Population Census. These data make it possible to search for specific individuals in the 1881 British or 1880 U.S. census. To construct the data for this study, we searched for individuals from two other censuses: the 1851 British and the 1850 U.S. census.

For Britain, we attempted to match all the English and Welsh born males age 25 and below from the computerized two percent sample of the 1851 census compiled principally by Anderson, Collins, and Stott. For the U.S. we attempted to match white males age 25 and below from the 1850 Federal Census one percent public use sample.[38] We employed a common matching technique for the British and U.S. data. Both countries' censuses provide information that either remains consistent between enumerations (name and birthplace) or changes predictably (age) that can be used to identify a given individual in more than one census. The British census has more specific

[38] The 1851 data for Britain are from a 2% Public Use Sample available as Study No. 1316 from the U.K. Data Archive at the University of Essex (http://www.dataarchive.ac.uk). It is a stratified two percent systematic cluster sample from the enumerators' books. For a full description see Anderson (1987). The complete 1881 census for Britain was obtained as Study No. 3643 from the U.K. Data Archive. The 1880 U.S. file was obtained from the North American Population Project (http://www.nappdata.org) and the 1850 U.S. 1% Public Use Sample was obtained from the Integrated Public Use Microdata Series available from the Minnesota Population Center (http://www.ipums.org).

information than the U.S. census on each individual's birthplace (parish in Britain, state in the U.S.). In the 1880 U.S. census, respondents were asked to give the place of birth of their parents as well (state for those whose parents were born in the U.S. and country for those whose parents were born abroad). This question was missing entirely from the nineteenth century British census.

For Britain, in order to be considered a true match for an individual from 1851, an individual from 1881 had to have either the same name or a close phonetic variation thereof (for example, Aitken and Aitkin were considered to be equivalent), a year of birth different by no more than five years, and the same county and parish of birth. For the U.S., the individual must provide the same state of birth for himself (and his parents if they were present in 1850) in 1850 and 1880, and the year of birth could differ by no more than three years. The variation in birth year was allowed in order to account for age misreporting, a fairly common phenomenon in nineteenth century societies which lacked the systematic record keeping and where individuals often had only an approximate idea of their age.[39] None of the matching information could be missing from an individual's record. Also, only unique matches were considered: if an individual from the 1850/51 sample had more than one match in the 1880/81 census, then that individual was dropped.[40]

Applying this matching process to 69,785 English and Welsh males age 25 and under from the 1851 two percent sample yielded 14,191 men observed in Britain both in 1851 and 1881, a success rate of 20%. From a pool of 43,438 U.S. white males age 25 and under in 1850, 9,497 were found in the 1880 U.S. census, a 22 percent success rate. The inability to link every observation from the initial public use sample (1850 for the U.S. and 1851 for Britain) is a function of mortality (and

[39] The smaller margin of age reporting error for the U.S. matching process is in response to the less specific birthplace information. For a discussion of age enumeration in the Victorian census, see Higgs (1986).

[40] The same procedure created the 1860-80, 1880-1900, and 1870-80 linked U.S. samples.

out-migration from Britain) over the following thirty years, under-enumeration in the terminal census (1880 for the U.S. and 1881 for Britain), and the inaccurate recording in either the initial or terminal year by the census takers or by those who performed the census transcriptions of the characteristics on which the linkage is based: name, year of birth, and birthplace (for the individual as well as his parents in the U.S. and for the individual only in Britain).[41]

For the U.S., 69 percent of white, native-born males under age 25 survived from 1850 to 1880 (based on the survival of five-year age cohorts in the IPUMS 1850 and 1880 samples); for Britain, 67 percent of males both survived from 1851 to 1881 and remained in Britain (based on published population-by-age tables in Mitchell, 1962, p. 12). Estimates of under-enumeration for the nineteenth century U.S. range from as high as 22% (Adams and Kasakoff, 1991) to as low as 9% (Hacker, 2000). Though we lack estimates of the extent of mis-reporting for names, birth years, and birth places, if we take the error in each of these to be 5 to 10 percent and assume for simplicity that all of the factors preventing linkage occur independently, we can calculate a set of projected linkage rates ranging from optimistic to pessimistic.[42] For the U.S., the anticipated linkage rate ranges from $(0.69)(0.91)(0.95)^{10}=37.6\%$ (“optimistic) to $(0.69)(0.78)(0.90)^{10}=18.8\%$ (“pessimistic”).[43]

The actual linkage rate for the U.S. is safely within this range, even without taking account of the fraction of individuals who could not be uniquely matched (e.g. they were matched to more than

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[41] Steckel (1991) surveys research on the accuracy of nineteenth century U.S. population censuses.

[42] One of the few studies to report an estimate of mis-reporting for a characteristic contained in the U.S. population census for the nineteenth century is Knights (1971): he reports that 11 percent of those located in Boston in both 1850 and 1860 reported a year of birth (inferred from age at the census) that differed by five or more years between the two censuses. Steckel (1988) found that literacy was inconsistently classified for seven percent of household heads located in both 1850 and 1860.

[43] There are five characteristics that must be reported correctly (name, birth year, own birth place, father's birth place, and mother's birthplace) in each of two censuses, so the proportion with correctly reported characteristics is between $(0.95)^{10}$ (if each is reported with error 5% of the time) and $(0.90)^{10}$ (if each is reported with error 10% of the time).

one individual in the 1880 census, and it was not possible to identify the best match). In 1880, 1.5 percent of white, native-born males shared the same name, birth year, birth place, and parents' birthplaces with at least one other individual, while 80.5% were uniquely identified by this set of characteristics. For the remaining 18%, there were several individuals who had names that were phonetically close and birth years that were within three years, but when an individual from the 1850 pubic use sample was matched to one of these individuals, it was possible in these cases to rank the matches by the proximity of the name and birth year, and choose the "best" match.

We lack estimates of under-enumeration and mis-reporting in the nineteenth century British censuses, though we know that the combined effects of mortality and net migration were slightly higher over the 1851-81 period than in Britain (only 67 percent of males age 25 and under present in 1851 would have still been present in Britain in 1881). At the same time, there was less variety in the distribution of surnames in Britain than in the U.S., so a larger fraction of potential matches (six percent) had to be discarded because a unique match could not be made. Using plausible assumptions for the under-enumeration and mis-reporting and for the probability of multiple matches, the British linkage rate can be shown to lie within the range of expected linkage rates.

The linkage for both the U.S. and Britain excluded those individuals who were linked from the initial census to more than one individual in the terminal census. In some cases, this discards potentially useful information. For example, if an individual whose father's occupation was observed in the 1850 U.S. public use sample was then linked to two individuals in the 1880 census, but both of those individuals had the same 1880 occupation, inclusion of either potential match, or a linear combination of them, would add the same information to a mobility table comparing the occupations of father and sons. We nonetheless excluded such individuals for two reasons: (1) their inclusion would induce a bias in favor of more common father-son occupation pairs; and (2) in some parts of the analysis (e.g. residential mobility, occupational mobility related to family wealth), the particular linkage made is more consequential than in the simple mobility calculations.

For each country, the data come from two nationally representative sources, so as long as the matching process does not skew the sample, the set of matched individuals should also be representative of the two national populations that survived 1850-80 and 1851-81. In order to assess the representativeness of the linked samples, we compared their characteristics to those in the public use samples for the initial year (1850 or 1851) and terminal year (1880 or 1881). Tables A-1 and A-2 present marginal effects from probit regressions in which the dependent variable is 1 for observations from the linked sample and 0 for observations from the public use sample.[44]

In general, the matched samples represent the overall population quite well. Though several characteristics exert a statistically significant influence on the probability of linkage, compared to the predicted probability the magnitude is small in each case.[45] In order to reduce the impact of these already small differences between the linked samples and the general population, we constructed weights to produce linked samples that would duplicate the marginal frequencies of the characteristics in the general population. (Deming and Stephan, 1940) Two sets of weights were generated, one for the initial year and one for the terminal year. In Columns (2) and (4) of Tables A1-1 and A1-2, the weights are imposed on the linked individuals, leaving them statistically

[44] Each linked individual thus enters the regression twice: once in the linked sample and once in the public use sample. This is done to facilitate comparison with the regressions in Columns (2) and (4) of each table in which weights are imposed to make the linked individuals nationally-representative (rather than merely indistinguishable from the unlinked). For the British, 8,655 individuals in the public use sample for 1851 were missing one or more of the characteristics used in the probit analysis and were dropped from the regressions in Columns (1) and (2) of Table A-2. For the comparisons with 1880 and 1881, 25 percent random samples of the complete files for those years were used rather than the complete files.

[45] The large coefficients on the migration history variables in Column (3) of Table A-2 result from the inability to identify the year of arrival in the U.S. for immigrants present in the U.S. in 1880 (the excluded category in the regression).

indistinguishable from the general population. Though we have used the unweighted data throughout this paper, the results are insensitive to the imposition of these weights. The weighting can eliminate the impact of linkage selectivity on observable characteristics; it cannot, however, eliminate the impact of unobservables on the linkage probability.

A final concern is whether the linkage process has resulted in too many "false positives" (individuals who are not in fact the same person in both the initial and terminal years). For the comparison between the U.S. and Britain in the nineteenth century, this is a difficulty if the sample for one country has more false positives than the sample for the other. For example, more false positives in the U.S. than in Britain will generate more "noise" in comparing the occupations of fathers and sons, and lead to a spurious finding of greater occupational mobility in the U.S. than in Britain. For the comparison over time within the U.S., it is also a problem, because the data for the twentieth century were constructed in a way that prevented such incorrect matches (the respondent in the OCG was asked himself to state the occupation of his father when the respondent was 16 years of age). Greater "noise" in the nineteenth century U.S. data would also produce a spurious finding of greater mobility in the nineteenth century U.S. than in the twentieth.

The comparison between mobility in Britain 1851-81 and in the U.S. 1850-80 was performed again, but this time with the samples restricted to those whose surnames matched exactly and whose age was off by no more than one year. Though this will not entirely eliminate false positives, it will reduce their prevalence, so if the difference between Britain and the U.S. in the nineteenth century persists, we can have greater confidence that this finding is not being driven by differences in the prevalence of false positives. The results were quite similar to those shown in the second panel of Table 2:

P=Britain 1851-81, Q=U.S. 1850-80

d(P,I)=24.50 (prob $<$ 0.0001)
d(Q,I)= 14.22 (prob $<$ 0.0001)
d(P,Q)=15.22 (prob $<$ 0.0001)

Though mobility in the U.S. with the more restricted sample is slightly farther from what we would observe if the occupations of fathers and sons were independent, the difference between mobility in the U.S. and in Britain is actually slightly greater. When the 1880-1900 U.S. sample is restricted to those whose surname matched exactly and whose age was off by no more than a year, the nineteenth century U.S. remained substantially more mobile than the twentieth century though the magnitude of the difference is reduced. We again can reject the null hypothesis that the association between the occupations of fathers and sons was identical in these two eras:

P=U.S. 1880-1900, Q=U.S. 1953-73

$d(P,I)=15.90$ (prob < 0.0001)
$d(Q,I)=20.76$ (prob < 0.0001)
$d(P,Q)=6.94$ (prob < 0.005).

Appendix 2: Log-Linear Analysis

Xie (1992) is the standard reference for differences in mobility across tables calculated using conventional log-linear analysis. The "log-multiplicative layer effect" is an estimate, for each "layer" in a contingency table (in the present context, a table is comprised of rows for sons' occupations, columns for fathers' occupations, and layers for countries or time periods), of the amount by which the row-column association in a layer must be multiplied to obtain the average row-column association across the entire table.

In order to test the robustness of our comparative mobility results to alternative log-linear analytical methods, we calculate and compare "Xie statistics" for four of our main mobility comparisons.

1. Britain (1972) vs. U.S. (1973)

Xie (1992, pp. 384-387), using the OCG data for the U.S. and the Oxford Mobility Study data for Britain, finds $\phi_{U.S.}=0.6064$, $\phi_{Britain}=0.6305$, and $\phi_{Japan}=0.4845$ (where the ϕ_i are the log-multiplicative layer effects for the off-diagonal cells) if no ordering is imposed on the occupational categories, so the row-column association is similar in the U.S. and Britain, but 20 to 23 percent lower in Japan. In a comparison of mobility for Britain, France and Sweden (p. 389), he finds $\phi_{Britain}=0.6167$, $\phi_{France}=0.6333$, and $\phi_{Sweden}=0.4676$. If we use the four-way classification of occupations in Table 1, impose no ordering on the categories, and calculate Xie's ϕ_i using all the cells in P and Q, we find $\phi_{Britain}=0.7783$ and $\phi_{U.S.}=0.6279$. Xie provides no test for the statistical significance of the difference $\phi_{U.S.}$- $\phi_{Britain}$, but bootstrapped standard errors yield probability (H_0: $\phi_{Britain} - \phi_{U.S.} = 0$) < 0.001. For the five-way occupational breakdown (which splits white collar into high and low subgroups), the ϕ_i are $\phi_{Britain}=0.8098$ and $\phi_{U.S.}=0.5867$.

2. Britain (1881) vs. U.S. (1880)

The ϕ_i for this comparison are $\phi_{Britain}$=0.8937 and $\phi_{U.S.}$=0.4487, a difference that is statistically significant at any conventional level (bootstrapped standard errors with 1,000 replications). The magnitude of the difference is three times greater in absolute terms than that in the twentieth century (0.4450 compared to 0.1504) and four times greater as a percentage of the U.S. figure (100% compared to 24%). If we split white collar into high and low white collar groups, the ϕ_i for this comparison are $\phi_{Britain}$=0.9092 and $\phi_{U.S.}$=0.4163, which is again substantially greater than the corresponding ϕ_i with five categories for the twentieth century.

3. U.S. (1880) vs. U.S. (1973)

The log-multiplicative layer effect model confirms greater mobility in the 1860-80 period than in the 1973 OCG, whether four or five categories are used. For four categories: $\phi_{1860\text{-}80}$=0.4986, ϕ_{OCG}=0.8668, probability(H_0: $\phi_{1860\text{-}80} - \phi_{OCG}=0) < 0.001$. For five categories, $\phi_{1860\text{-}80}$=0.4932, ϕ_{OCG}=0.8699, probability(H_0: $\phi_{1860\text{-}80} - \phi_{OCG}=0) < 0.001$.

4. U.S. (1900) vs. U.S. (1973)

Again, the log-multiplicative layer effect model confirms greater mobility in the 1880-1900 period than in the 1973 OCG, whether four or five categories are used. For four categories: $\phi_{1880\text{-}1900}$=0.6341, ϕ_{OCG}=0.7733, probability(H_0: $\phi_{1880\text{-}1900} - \phi_{OCG}=0) < 0.05$. For five categories, $\phi_{1880\text{-}1900}$=0.6258, ϕ_{OCG}=0.7800, probability(H_0: $\phi_{1880\text{-}1900} - \phi_{OCG}=0) < 0.001$.

Appendix 3: Expanded Occupational Categories
For online publication only

The following tables provide the raw frequency counts for all of the 5 × 5 and 6 × 6 contingency tables, which can be collapsed down to 4 × 4 tables like those shown in the text by combining high and low white collar or skilled and semi-skilled. The Xie statistic was calculated using the "unidiff" macro in STATA, with a modification to return the ϕ_p, ϕ_q, and ϕ_p-ϕ_q as scalars that could then be bootstrapped (the modified "unidiff" macro is available on request, together with the short program that does the bootstrapping, again using the frequencies in Table 3 that generated the results reported in footnote 27). The entries in Tables 4 and 6 were generated by repeated calculation of the Altham statistic for each odds ratio in Tables P and Q. Finally, this appendix provides a version of Table 1 from the text which imposes each country's marginal frequencies on the other country.

1. Raw 5 × 5 and 6 × 6 contingency tables:

TABLE A2-1 HERE

TABLE A2-2 HERE

TABLE A2-3 HERE

TABLE A2-4 HERE

2. British and U.S. Mobility in the Twentieth Century With Each Country's Marginal Frequencies Replaced By Those From the Other Country

TABLE A2-5 HERE

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TABLE 1–INTERGENERATIONAL OCCUPATIONAL MOBILITY
IN BRITAIN AND THE U.S.,1949-55 TO 1972-73, FREQUENCIES (COLUMN PERCENT)

	Father's Occupation				
Son's Occupation	White Collar	Farmer	Skilled/ Semiskilled	Unskilled	Row Sum
Britain (Table P):					
White Collar	174	11	206	38	429
	(68.2)	(25.6)	(30.7)	(24.5)	
Farmer	2	9	3	1	15
	(0.8)	(20.9)	(0.4)	(0.6)	
Skilled/Semiskilled	71	19	417	102	609
	(27.8)	(44.2)	(62.2)	(65.8)	
Unskilled	8	4	44	14	70
	(3.1)	(9.3)	(6.6)	(9.0)	
Column Sum	255	43	670	155	1123
U.S. (Table Q):					
White Collar	595	144	539	164	1442
	(71.4)	(31.9)	(43.6)	(35.1)	
Farmer	3	61	7	5	76
	(0.4)	(13.5)	(0.6)	(1.1)	
Skilled/Semiskilled	186	193	576	236	1191
	(22.3)	(42.8)	(46.6)	(50.5)	
Unskilled	49	53	115	62	279
	(5.9)	(11.8)	(9.3)	(13.3)	
Column Sum	833	451	1237	467	2988

Notes: Occupation of father when respondent was age 14 (Britain) or age 16 (U.S.), compared to occupation at survey in 1972 (Britain) or 1973 (U.S.), males 31-37 (Britain) and 33-39 (U.S.) in survey year.

TABLE 2–SUMMARY MEASURES OF MOBILITY IN BRITAIN AND THE U.S.

Comparison and Terminal Year	M (1)	M′ (2)	d(P,I) (3)	G^2 (4)	d(Q,I) (5)	G^2 (6)	d(P,Q) (7)	G^2 (8)	d^i(P,Q) (9)	G^2 (10)
1. Britain 1972 (P)	45.3	53.7	24.0	168.4***			7.9	7.5	7.2	2.4
vs. U.S. 1973 (Q)	56.7	48.3			20.8	420.4***				
2. Britain 1881 (P)	42.6	35.5	22.7	777.15***			13.2	98.2***	4.5	6.9
vs. U.S. 1880 (Q)	45.4	47.9			11.9	287.2***				
3. U.S. 1880 (P)	50.6	57.7	12.1	385.4***			10.7	46.7***	2.4	3.2
vs. U.S. 1973 (Q)	56.7	43.7			20.8	420.4***				
4. U.S. 1900 (P)	54.0	54.1	14.6	545.4***			9.1	36.7***	2.4	3.9
vs. U.S. 1973 (Q)	56.7	51.8			20.8	420.4***				

Notes: M is total mobility (percent off the main diagonal), M′ is total mobility using the marginal frequencies from the other table (see Appendix 2), G^2 is the likelihood ratio χ^2 statistic with significance levels *** < 0.01 ** < 0.05 * < 0.10. Degrees of freedom: 9 for columns (4), (6), and (8); 5 for column (10).

TABLE 3–INTERGENERATIONAL OCCUPATIONAL MOBILITY IN BRITAIN AND THE U.S., 1850-51 TO 1880-81, FREQUENCIES (COLUMN PERCENT)

	Father's Occupation				
Son's Occupation	White Collar	Farmer	Skilled/ Semiskilled	Unskilled	Row Sum
Britain (Table P):					
White Collar	103	31	219	63	416
	(36.6)	(11.1)	(13.3)	(7.3)	
Farmer	8	114	39	21	182
	(2.8)	(40.9)	(2.4)	(2.4)	
Skilled/Semiskilled	143	90	1155	386	1774
	(50.0)	(32.3)	(70.2)	(44.6)	
Unskilled	32	44	233	395	704
	(11.2)	(15.8)	(14.2)	(45.7)	
Column Sum	286	279	1646	865	2976
U.S. (Table Q):					
White Collar	55	177	82	30	344
	(38.5)	(12.9)	(22.6)	(23.3)	
Farmer	44	850	92	35	1021
	(30.8)	(62.0)	(25.3)	(27.1)	
Skilled/Semiskilled	33	214	166	40	453
	(23.1)	(15.6)	(45.7)	(31.0)	
Unskilled	11	129	23	24	187
	(7.7)	(9.4)	(6.3)	(18.6)	
Column Sum	143	1370	363	129	2005

Notes: Occupation of father in 1851 (Britain) or 1850 (U.S.) when son was age 13-19, compared to occupation of son in 1881 (Britain) or 1880 (U.S.), males 41-49 in 1881 (Britain) or 1880 (U.S.).

TABLE 4–COMPONENTS OF d(P,I), d(Q,I), AND d(P,Q) FOR BRITAIN 1851-81 (P) VS. U.S. 1850-80 (Q)

Contrast	d(P,I)	Odds Ratio	G^2	d(Q,I)	Odds Ratio	G^2	d(P,Q)	G^2	Pct. of Total	Cumulative Percent
[(FF)/(FU)]/[(UF)/(UU)]	7.77	48.73	272.95***	3.02	4.52	25.30***	4.76	37.94***	12.90	12.90
[(FF)/(FU)]/[(SF)/(SU)]	5.48	15.48	149.30***	1.00	1.65	3.65*	4.48	45.17***	11.40	24.30
[(WW)/(WF)]/[(FW)/(FF)]	7.71	47.35	146.35***	3.58	6.00	65.49***	4.13	22.94***	9.70	34.10
[(WF)/(WS)]/[(FF)/(FS)]	6.24	22.64	114.29***	2.18	2.98	18.80***	4.06	23.81***	9.40	43.50
[(WF)/(WU)]/[(FF)/(FU)]	4.68	10.36	36.75***	1.00	1.65	1.86	3.68	12.39***	7.70	51.20
[(FF)/(FS)]/[(SF)/(SS)]	7.25	37.51	342.03***	3.94	7.17	181.68***	3.31	41.65***	6.20	57.40
[(FF)/(FS)]/[(UF)/(US)]	6.30	23.28	199.97***	3.03	4.54	37.08***	3.27	21.57***	6.10	63.50
[(WW)/(WU)]/[(UW)/(UU)]	6.01	20.18	188.45***	2.77	4.00	11.21***	3.24	9.94***	6.00	69.50
[(FW)/(FF)]/[(SW)/(SF)]	6.06	20.65	165.45***	2.91	4.28	67.65***	3.15	26.31***	5.60	75.10

Notes: First element of each pair is father's occupation, second is son's. W: White Collar, S: Skilled, F: Farmer, U: Unskilled.

TABLE 5–INTERGENERATIONAL OCCUPATIONAL MOBILITY IN THE U.S., 1860-80, FREQUENCIES (COLUMN PERCENT)

	Father's Occupation				
Son's Occupation	White Collar	Farmer	Skilled/ Semiskilled	Unskilled	Row Sum
U.S. 1880 (Table P):					
White Collar	115	233	115	39	502
	(46.0)	(13.8)	(25.2)	(16.5)	
Farmer	43	949	103	60	1155
	(17.2)	(56.2)	(22.5)	(25.3)	
Skilled/Semiskilled	59	286	173	75	593
	(23.6)	(16.9)	(37.9)	(31.6)	
Unskilled	33	220	66	63	382
	(13.2)	(13.0)	(14.4)	(26.6)	
Column Sum	250	1688	457	237	2632

Notes: Occupation of father in 1860 when son was age 13-19, compared to occupation of son in 1880, males 31-39 in 1880.

TABLE 6–COMPONENTS OF d(P,I), d(Q,I), AND d(P,Q) FOR U.S. 1860-80 (P) VS. U.S. 1973 (Q)

Contrast	d(P,I)	Odds Ratio	G^2	d(Q,I)	Odds Ratio	G^2	d(P,Q)	G^2	Pct. of Total	Cumulative Percent
[(WW)/(WF)]/[(FW)/(FF)]	4.78	10.89	176.51***	8.86	84.02	159.17***	4.09	16.29***	14.7	14.7
[(FW)/(FF)]/[(SW)/(SF)]	3.03	4.55	95.11***	6.97	32.62	131.80***	3.94	27.39***	13.7	28.3
[(FF)/(FU)]/[(SF)/(SU)]	2.03	2.76	31.77***	5.88	18.91	72.34***	3.85	21.52***	13.0	41.4
[(WF)/(WU)]/[(FF)/(FU)]	2.39	3.31	22.33***	5.87	18.80	40.93***	3.47	8.93***	10.6	52.0
[(FW)/(FF)]/[(UW)/(UF)]	1.95	2.65	18.42***	5.26	13.89	53.91***	3.32	12.80***	9.7	61.6
[(FF)/(FS)]/[(SF)/(SS)]	3.44	5.57	154.21***	6.52	26.01	115.85***	3.08	16.58***	8.4	70.0
[(WF)/(WS)]/[(FF)/(FS)]	3.03	4.55	51.26***	5.95	19.60	55.06***	2.92	7.15***	7.5	77.5

Notes: First element of each pair is father's occupation, second is son's. W: White Collar, S: Skilled, F: Farmer, U: Unskilled.

TABLE A1-1. PROBIT MARGINAL EFFECTS ON LINKAGE (1=LINKED SAMPLE, 0=PUBLIC USE SAMPLE), BRITAIN

Variable	1851, No Weights $\partial P/\partial X$	1851, Weights $\partial P/\partial X$	1881, No Weights $\partial P/\partial X$	1881, Weights $\partial P/\partial X$
Age 15-25 in 1851	0.0069	0.0015	-0.0015	0.0000
	(1.48)	(0.30)	(21.32)***	(0.44)
Residence:				
London	0.0335	0.0067	0.0011	-0.0002
	(4.46)***	(0.87)	(6.41)***	(1.12)
Midlands-East	0.0354	0.0078	0.0012	-0.0002
	(4.97)***	(1.06)	(7.21)***	(0.94)
North	0.0405	0.0091	0.0004	-0.0002
	(5.66)***	(1.23)	(2.85)***	(1.15)
South	0.0305	0.0056	0.0022	-0.0002
	(4.17)***	(0.75)	(11.58)***	(0.90)
Migration History:				
Birth County=Residence	0.0698	0.0014	0.0007	0.0000
	(13.64)***	(0.22)	(9.66)***	(0.52)
Occupation[a]:				
Farmer	0.0148	0.0030	0.0014	0.0000
	(3.90)***	(0.78)	(8.65)***	(0.01)
Craftsman	0.0332	0.0065	0.0005	-0.0001
	(5.52)***	(1.09)	(6.00)***	(1.03)
Laborer	0.0131	0.0023	0.0000	-0.0001
	(3.69)***	(0.62)	(0.44)	(0.81)
Attended School	0.0167	0.0026		
	(4.81)***	(0.73)		
Employed Outside Home	0.0096	0.0013		
	(2.09)**	(0.28)		
Married			0.0000	0.0001
			(0.42)	(0.42)
Head			0.0014	0.0000
			(13.01)***	(0.23)
Observations	75,321	75,321	934,852	934,852
Pseudo-R^2	0.0043	0.0001	0.0065	0.0000
Predicted Probability	0.1830	0.1886	0.0038	0.0041

Notes: Absolute value of z statistics in parentheses. * significant at 10%; ** 5%; *** 1%. Omitted categories: "Age 0-14 in 1851," "Wales," "Other." 1881 uses a 25% sample of the unlinked. [a] Father's occupation in 1851, Son's occupation in 1881.

TABLE A1-2–PROBIT MARGINAL EFFECTS ON LINKAGE (1=LINKED SAMPLE, 0=PUBLIC USE SAMPLE), U.S.

Variable	1850, No Weights $\partial P/\partial X$	1850, Weights $\partial P/\partial X$	1880, No Weights $(\partial P/\partial X)$x100	1880, Weights $(\partial P/\partial X)$x100
Age 15-25 in 1850	0.0017	0.0013	0.0314	0.0000
	(0.41)	(0.31)	(14.38)***	(0.01)
Residence:				
Midwest	-0.0478	-0.0008	-0.0216	0.0001
	(11.26)***	(0.19)	(8.40)***	(0.02)
South[a]	-0.0519	-0.0011	-0.0235	0.0002
	(12.13)***	(0.25)	(9.17)***	(0.04)
West			-0.0398	0.0003
			(7.83)***	(0.02)
Population > 2,500	-0.0013	-0.0028		
	(0.25)	(0.52)		
Migration History:				
Interstate Mover	-0.0124	-0.0001	0.3214	-0.0001
	(1.34)	(0.01)	(43.35)***	(0.02)
Birthstate=Residence	0.0090	0.0001	0.3123	-0.0002
	(1.10)	(0.01)	(50.16)***	(0.03)
Family Size	-0.0003	-0.0003	-0.0004	-0.0004
	(0.44)	(0.35)	(1.22)	(0.50)
Occupation[b]:				
Farmer	0.0138	-0.0015	0.0126	-0.0003
	(2.07)**	(0.22)	(4.16)***	(0.04)
Skilled	0.0003	-0.0011	-0.0014	-0.0001
	(0.05)	(0.14)	(0.38)	(0.01)
Semi-Skilled	0.0086	-0.0008	-0.0012	-0.0000
	(0.93)	(0.08)	(0.28)	(0.00)
Laborer	0.0128	-0.0019	-0.0102	0.0001
	(1.52)	(0.22)	(2.85)***	(0.01)
Other	-0.0106	-0.0011	-0.0059	-0.0001
	(0.88)	(0.09)	(0.98)	(0.01)
Household Real Estate:				
0 < Real Estate < $1,500	0.0104	0.0024		
	(2.37)**	(0.54)		
Real Estate ≥$1,500	0.0255	-0.0025		
	(5.62)***	(0.56)		
Father Literate	-0.0067	-0.0058		
	(1.09)	(0.93)		
Attended School	0.0065	0.0006		
	(1.82)*	(0.17)		
Household Head			-0.0014	0.0024
			(0.35)	(0.26)
Married			0.0094	-0.0022
			(2.65)***	(0.25)
Observations	52,935	52,935	1,766,147	1,766,147
Pseudo-R^2	0.0071	0.0001	0.0356	0.0000
Predicted Probability	0.1794	0.1794	0.0014	0.0014

Notes: Absolute value of z statistics in parentheses. * significant at * 10%; ** 5%; *** 1%. Omitted categories: “Age 0-14 in 1850,” “Northeast,”, “Population ≤ 2,500”, “Foreign-Born,” “White Collar,” “Household Real Estate=0,” “Father Illiterate,” “Not Attending School,” “Non-Head,” and “Unmarried.” 1880 uses a 25% sample of the unlinked.
[a] Includes “West” in 1850. [b] Father’s occupation in 1850, Son’s occupation in 1880.

TABLE A2-1–INTERGENERATIONAL OCCUPATIONAL MOBILITY IN BRITAIN AND THE U.S.,1949-55 TO 1972-73, FREQUENCIES (FIVE-WAY CATEGORIZATION)

	Father's Occupation				
Son's Occupation	High White Collar	Low White Collar	Farmer	Skilled/ Semiskilled	Unskilled
Britain (Table P):					
H. White Collar	76	51	8	136	20
L. White Collar	19	28	3	70	18
Farmer	1	1	9	3	1
Skilled/Semiskilled	19	52	19	417	102
Unskilled	6	2	4	44	14
U.S. (Table Q):					
H. White Collar	349	134	108	404	116
L. White Collar	70	43	36	135	47
Farmer	2	1	61	7	5
Skilled/Semiskilled	116	69	193	576	236
Unskilled	29	20	53	115	62

Notes: Occupation of father when respondent was age 14 (Britain) or age 16 (U.S.), compared to occupation at survey in 1972 (Britain) or 1973 (U.S.), males 31-37 (Britain) and 33-39 (U.S.) in survey year.

TABLE A2-2–INTERGENERATIONAL OCCUPATIONAL MOBILITY
IN THE U.S., 1860-80 AND 1949-55 TO 1973, FREQUENCIES (FIVE-WAY CATEGORIZATION)

	Father's Occupation				
Son's Occupation	High White Collar	Low White Collar	Farmer	Skilled/ Semiskilled	Unskilled
U.S. 1850-80 (Table P):					
H. White Collar	65	10	174	69	26
L. White Collar	33	7	59	46	13
Farmer	37	6	949	103	60
Skilled/Semiskilled	52	7	285	173	75
Unskilled	30	3	220	66	63
U.S. 1949-55 to 1973 (Table Q):					
H. White Collar	349	134	108	404	116
L. White Collar	70	43	36	135	47
Farmer	2	1	61	7	5
Skilled/Semiskilled	116	69	193	576	236
Unskilled	29	20	53	115	62

Notes: Occupation of father when respondent was age 16, compared to occupation in 1880 or 1973, males 31-39 in terminal year.

TABLE A2-3–INTERGENERATIONAL OCCUPATIONAL MOBILITY
IN BRITAIN AND THE U.S., 1850-51 TO 1880-81, FREQUENCIES (FIVE-WAY CATEGORIZATION)

	Father's Occupation				
Son's Occupation	High White Collar	Low White Collar	Farmer	Skilled/ Semiskilled	Unskilled
Britain 1851-81 (Table P):					
H. White Collar	20	19	4	59	9
L. White Collar	15	49	27	160	54
Farmer	1	7	114	39	21
Skilled/Semiskilled	19	124	90	1,155	386
Unskilled	6	26	44	233	395
U.S. 1850-80 (Table Q):					
H. White Collar	9	15	75	24	17
L. White Collar	14	17	102	58	19
Farmer	20	24	850	92	57
Skilled/Semiskilled	12	21	214	166	45
Unskilled	8	13	165	38	43

Notes: Occupation of father when respondent was age 16, compared to occupation in 1880 or 1881, males 41-49 in terminal year.

TABLE A2-4–INTERGENERATIONAL OCCUPATIONAL MOBILITY IN BRITAIN AND THE U.S., 1850-51 TO 1880-81, FREQUENCIES (SIX-WAY CATEGORIZATION)

	Father's Occupation					
Son's Occupation	High White Collar	Low White Collar	Farmer	Skilled	Semiskilled	Unskilled
Britain 1851-81 (Table P):						
H. White Collar	20	19	4	54	5	9
L. White Collar	15	49	27	139	21	54
Farmer	1	7	114	34	5	21
Skilled	19	112	68	878	92	294
Semiskilled	1	12	22	78	107	92
Unskilled	6	26	44	202	31	395
U.S. 1850-80 (Table Q):						
H. White Collar	9	15	75	18	6	17
L. White Collar	14	17	102	38	20	19
Farmer	20	24	850	75	17	57
Skilled	8	12	140	73	28	34
Semiskilled	4	9	74	38	27	11
Unskilled	8	13	165	22	16	43

Notes: Occupation of father when respondent was age 16, compared to occupation in 1880 or 1881, males 41-49 in terminal year.

TABLE A2-5–INTERGENERATIONAL OCCUPATIONAL MOBILITY
IN BRITAIN AND THE U.S.,1949-55 TO 1972-73, WITH EACH COUNTRY'S MARGINAL FREQUENCIES REPLACED WITH THOSE FROM THE OTHER COUNTRY, FREQUENCIES (COLUMN PERCENT)

	Father's Occupation				
Son's Occupation	White Collar	Farmer	Skilled/ Semiskilled	Unskilled	Row Sum
Britain (Table P, With U.S. Marginal Frequencies From Table 1):					
White Collar	637.0	155.9	497.1	152.0	1442.0
	(76.5)	(34.6)	(40.2)	(32.5)	
Farmer	3.8	66.3	3.8	2.1	76.0
	(0.5)	(14.7)	(0.3)	(0.4)	
Skilled/Semiskilled	159.3	165.0	616.7	250.0	1191.0
	(19.1)	(36.6)	(49.9)	(53.5)	
Unskilled	32.9	63.7	119.4	63.0	279.0
	(4.0)	(14.1)	(9.7)	(13.5)	
Column Sum	833.0	451.0	1237.0	467.0	
U.S. (Table Q, With British Marginal Frequencies From Table 1):					
White Collar	157.9	10.0	220.8	40.3	429.0
	(61.9)	(23.2)	(33.0)	(26.0)	
Farmer	1.3	7.0	4.7	2.0	15.0
	(0.5)	(16.2)	(0.7)	(1.3)	
Skilled/Semiskilled	84.3	22.8	402.8	99.1	609.0
	(33.0)	(53.1)	(60.1)	(63.9)	
Unskilled	11.5	3.3	41.7	13.5	70.0
	(4.5)	(7.6)	(6.2)	(8.7)	
Column Sum	255.0	43.0	670.0	155.0	

Notes: Occupation of father when respondent was age 14 (Britain) or age 16 (U.S.), compared to occupation at survey in 1972 (Britain) or 1973 (U.S.), males 31-37 (Britain) and 33-39 (U.S.) in survey year. The entries were generated by imposing the marginal frequencies from the other country on each country's mobility pattern in Table 1. This was done using the algorithm described in Altham and Ferrie (2007).

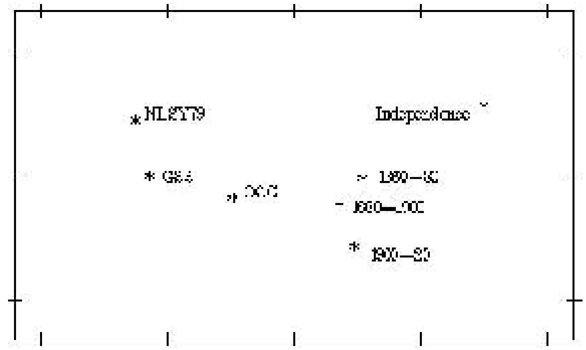

FIGURE 1. INTERGENERATIONAL OCCUPATIONAL MOBILITY IN THE U.S. IN SIX SAMPLES (MULTIDIMENSIONAL SCALING SCORES)

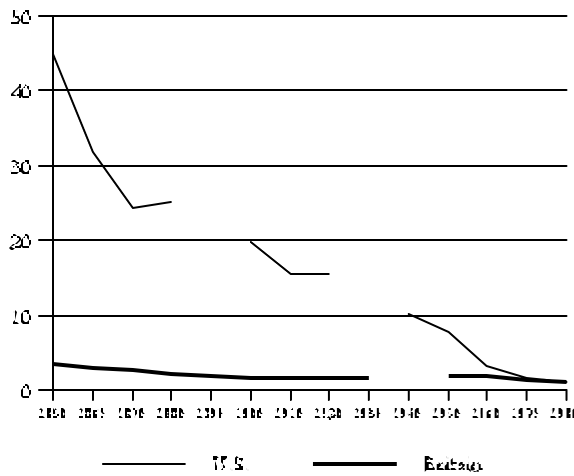

FIGURE 2. PERCENTAGE OF LABOR FORCE IN AGRICULTURE

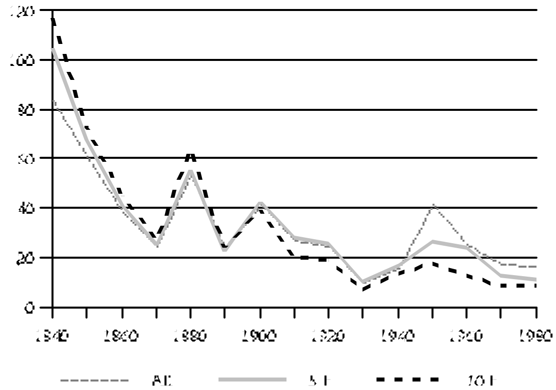

FIGURE 3 STANDARD DEVIATION IN CITY POPULATION GROWTH RATES, 100 LARGEST U.S. CITIES