Tightening Up: Declining Class Mobility during Russia's Market Transition

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American Sociological Review 2004 69: 677
DOI: 10.1177/000312240406900504

Tightening Up: Declining Class Mobility during Russia's Market Transition

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This study analyzes intergenerational occupational mobility in late Soviet and post-Soviet Russia using data from six surveys. Belying claims that class differences did not matter in the Soviet Union, the authors find that social origin did affect occupational opportunity during Russia's Soviet period. But the transition from state socialism to a market economy tightened the link between origins and destinations. Men and women were equally constrained by their social origin, even though they faced significantly different opportunity structures in both periods. As the economic transformation took hold, fewer Russians experienced upward mobility and more were downwardly mobile. Political and economic transition, not the demographic replacement of retiring cohorts by younger ones, strengthened the association between origins and destinations. Career mobility during the 1990s took the form of a regression toward origins, as workers who had the most upward mobility during the Soviet era lost the most in the transition to markets, abetting the reproduction of the class structure across generations as they fell.

The rapid transitions to economies based on markets that most state socialist countries undertook during the 1990s offer unique opportunities to observe how institutional changes affect social stratification. But research on postsocialist stratification has focused almost exclusively on *intragenerational* processes, such as earnings determinants, elite formation, and labor-market transitions (e.g., Nee 1989, 1996; Rona-Tas 1994; Bian and Logan 1996; Xie and Hannum 1996; Zhou, Tuma, and Moen 1997; Gerber and Hout 1998; Cao and Nee 2000; Gerber 2000b, 2001a, 2001b, 2002; Zhou 2000; Walder 2002). We open a new direction in the study of postsocialist stratification by examining how market transition affects *intergenerational* inequalities.

We examine the association between class origins and destinations in late Soviet and post-Soviet Russia using modern tools of mobility research. To set up the context for our empirical analysis we address lingering questions regarding the significance of class differences in Soviet society and the appropriateness of applying the Erikson-Goldthorpe class schema to the USSR. We also argue for treating managers and professionals as separate class categories, a departure from many applications of the Erikson-Goldthorpe schema. We then consider theoretical grounds for expecting the link between origins and destinations to tighten up in Russia following market reforms. Our empirical analyses use 10,264 valid observations from six nationally representative surveys of Russian adults. We describe the observed mobility flows based on our preferred specification of class categories, check for gender differences in the ori-

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gin-destination association, and apply four different mobility models.

Most important, we assess whether Russian mobility changed significantly after the collapse of the Soviet system. Our data are uniquely suited for this task, since roughly half our observations come from the pretransition era (1988–92)[1] and the other half come from the posttransition era (1998–2000). Our analysis of change has three components. First, we compare gross rates of different types of mobility in the pre- and the posttransition period. Second, we perform statistical tests for change in the net origin-destination (OD) association. These tests reveal a statistically significant and substantial increase in association, regardless of which type of mobility model is used. Third, we determine whether this strengthening of the effect of origins on destinations in Russia results from cohort replacement or a period effect. We find that institutional change induced a pattern of (intragenerational) occupational moves that disproportionately demoted people who had been most upwardly mobile under Communism, resulting in a regression toward origins. This led to a tightening up of the OD association in posttransition Russia, a development virtually without precedent.

Mobility researchers have long sought to understand whether and how national institutional arrangements shape intergenerational inequalities. They have usually attempted to do so by comparing mobility patterns in different countries at similar times (Featherman, Jones, and Hauser 1975; Grusky and Hauser 1984; Erikson and Goldthorpe 1987, 1992b; Ishida, Müller, and Ridge 1995). But small sample size and collinearity among explanatory factors produce uncertainty as to whether cross-national variations in institutions, in level of development, or in culture account for similarities or differences in mobility. Analyses that track a single country over time (e.g., Grusky and DiPrete 1990; Jonsson and Mills 1993) encounter similar problems when changes along these dimensions are incremental and correlated. Market transition in Russia, however, altered so many fundamental economic institutions so rapidly that we can confidently ascribe changes in social mobility during the 1990s to this source rather than to cultural change or industrialization, making Russia an especially informative case for mobility researchers.

UNDERSTANDING MOBILITY IN RUSSIA

Classes in Soviet-Era Russia?

Soviet leaders trumpeted the open character of Soviet society, and Soviet sociologists repeatedly asserted that origin-based differences in occupational position were disappearing even as they produced rudimentary empirical evidence that such differences had not yet disappeared (e.g., Shubkin 1965; Rutkevich 1977; Rutkevich and Fillipov 1978). Ironically, Western observers who adhered to the totalitarian model of Soviet politics and society echoed the view that class distinctions had no meaning in the Soviet Union; they believed that all group differences were erased by the powerful, terroristic state apparatus (Feldmesser 1960). Nonetheless, most specialists on the Soviet Union would now agree that both Soviet leaders and totalitarian theorists erred when they proclaimed that Soviet society was classless (e.g., Connor 1991). More recently, some Western sociologists argue that social class distinctions have lost significance in modern societies (Clark and Lipset 1991; Giddens 1994; Beck 2000). Many stratification researchers, however, criticize this "end of class" thesis, contending that class continues to shape opportunities in developed capitalist countries (e.g., Goldthorpe and Marshall 1992; Hout, Brooks, and Manza 1995; Goldthorpe 2002). Still, in light of this debate and earlier, Soviet-era assertions of classlessness, researchers must take seriously the claim that classes do not matter in the Soviet Union. We address this assumption at length in our data analysis. But first we discuss what role class distinctions may have played in Soviet society.

Soviet politicians and sociologists based their claim that classes did not exist in the USSR on the complete absence of private ownership of the means of production there.[2] But although pri-

[1] Technically, of course, 1992 postdates the fall of communism. But the data were collected early in 1992 (February), so they reflect the stratification process just as the transformation was launched—before its effects could have been felt.

[2] More precisely, the official position held that Soviet society consisted of two classes—the prole-

vate ownership of the means of production is a key factor shaping class position, Weberian, neo-Marxist, and cultural capital approaches to class analysis have all drawn attention to other important criteria, including distinctions of skill, discursive ability, credentials, authority, and employment contracts. All of these distinctions existed in Soviet-era Russia. Not surprisingly, many theorists have characterized the class structure of the USSR and other state socialist nations as based on one or several of these dimensions (Timasheff 1944; Djilas 1957; Gouldner 1979; Konrad and Szelenyi 1979). These theories all point to two main factors as the defining features of class position in state socialist societies: bureaucratic authority, often equated with membership in the ruling Communist Party; and expertise. For example, in Wright's (1985) neo-Marxist framework, "organization assets" play a greater role as the source of class advantages in state socialist societies than in capitalist societies because control over the means of production was coordinated through an administrative hierarchy; and "skill assets" provide important advantages as well. Although couched in different terminology, Walder's "dual elite" model for China also stipulates that political loyalty/authority and education-based expertise represent alternative bases for membership in a Chinese elite consisting of party officials and professionals (Walder 1995; Walder, Li, and Treiman 2000).

These studies analyze class from an intragenerational perspective, focusing on how class location affects material standing, prestige, and power. While they do not directly address the intergenerational transmission of class position, they establish that class position affects life chances in state socialist societies.[3] Were this not the case, origin-based inequality in access to certain class positions would not matter.

The effect of class position on life chances in state socialist societies provides us with theoretical grounds to expect that occupational class origins shape destinations in Soviet-era and post-Soviet Russia. We need only assume that parents try and, to some degree, manage to pass their advantages on to their offspring. So long as inequalities are systematically linked to occupational classes based on expertise, credentials, authority, and status, parents in privileged classes will do all they can to improve their childrens' access to privileged class positions. Thus, if class position affects life chances in Soviet-era and post-Soviet Russia, we should observe some association between class origins and class destinations in our data from the Soviet period.

Several analyses of class structure in the Soviet bloc countries have explicitly examined the intergenerational transmission of class position. Parkin (1972, 1979) applied a neo-Weberian framework emphasizing social closure processes as the basis of class position. He pointed to pronounced rates of upward intergenerational mobility into the ranks of experts as a factor that tended to defuse social conflicts and prevent the crystallization of a coherent class culture. True enough, concerted campaigns to create a "socialist intelligentsia" sparked high mobility from the working class to specialist occupations in the 1920s and early 1930s (Fitzpatrick 1979). But the Soviet regime abandoned these policies by the end of the 1930s (Timasheff 1944).

From that time onward, the Soviet state sought to manipulate mobility only indirectly—through educational expansion and, during the Khrushchev era, university admissions policies intended to favor those with worker origins (see Gerber and Hout 1995). Soviet industrial growth spurred massive structural mobility from the peasantry to the working class (Connor 1991), but this process was essentially spent by the 1950s. Western observers detected by the 1970s an increasing heredity of class position in Soviet society, based on a smattering of highly localized survey data reported by Soviet sociologists (Matthews 1972; Yanowitch 1977; Dobson 1980; Lapidus 1983). Key mechanisms for the reproduction of elite class position were education and Communist Party membership. Cohort-based studies using more representative data found throughout the post-World War II era suggest that the effects of family background on educational attainment—an essential

tariat and the peasantry—and a "stratum" (*sloi*) of intelligentsia, but that collective ownership removed any antagonism or systematic inequality among these groups. See Shkaratan (1996).

[3] Szelenyi (1988) examines the third-generation inheritance of agricultural entrepreneurship in Hungary. But the link between this study and his earlier work on the intellectual class is not clear, and the strict proscription of self-employment in Soviet-era Russia for most of seven decades makes it hard to apply the findings from Hungary.

contributor to the association between origins and destinations—were stable or increasing (Gerber and Hout 1995; Gerber 2000a). Communist Party membership, a prerequisite for access to most positions of authority, was substantially more accessible to the offspring of Communist Party members (Gerber 2000b, 2001b). These effects suggest that the postwar era saw the growth of increasingly hereditary classes of salaried professionals and managers.

While the consensus built that advantages were passed from fathers to sons in Soviet Russia, the precise degree of class inheritance in late Soviet-era Russia has remained a matter of speculation. Evidence based on national samples from the 1970s shows that origins and destinations were correlated in the Eastern European countries of the Soviet bloc (e.g., Erikson and Goldthorpe 1992b). However, no national data emerged regarding the level of social mobility in Russia until the dawn of the post-Soviet era. Without data, researchers could not distinguish structural from exchange mobility or apply modern techniques for the measurement of mobility patterns.

During the transition, Marshall, Sydorenko, and Roberts (1997) produced the first and only (to our knowledge) nationally representative study of Soviet-era mobility patterns in Russia using data collected in the fall of 1991. They reported substantial structural mobility out of the agricultural classes, moderate shrinking of the industrial working class (reminiscent of the British pattern), and an unusually large professional and managerial class, which they attributed to the proliferation of bureaucratic positions in the Soviet administrative apparatus. They found that men and women experienced a common pattern of association between origins and destinations, even though they worked in different occupations due to the gender typing of jobs in the Soviet era. They also compared Russian and British mobility and found no difference between Russian and British men and that the association between class origins and destinations was weaker for Russian than for British women. Modest sample size (1,150 cases) precluded more detailed analysis. It also tempers our confidence in these authors' findings of no difference between Russian and British men. With barely 600 observations, their test of cross-national differences lacks statistical power.

In light of Soviet assertions of classlessness and current sociological claims of the "end of class," we cannot presume in advance that class origins affected destinations in Soviet-era Russia. Even though what little we know about Russian mobility in the Soviet period suggests that there was substantial intergenerational inheritance of class position, the data are too sketchy to reach definitive conclusions. Therefore, our first objective is to measure the pattern and strength of the origin-destination association, which will give us the first comprehensive picture of the degree of intergenerational mobility in the world's first and longest-lasting state socialist society.

Choosing a Class Schema

Parkin's (1972) conception of the Soviet class structure, with its distinction of grades within nonmanual and manual classes, recalls the widely used class schema of Erikson and Goldthorpe (1992b). In its most basic form, the Erikson-Goldthorpe schema is based on four asymmetric distinctions that shape the rewards, status, and working conditions associated with different jobs: (1) owner-proprietors versus hired employees (property relations); (2) among hired employees, the difference between salaried, service employment, and hourly contract labor (employment relationship); (3) among hourly workers, the nonmanual-manual divide (type of work); and (4) among manual workers, agricultural versus all other (sector). More elaborate versions of the schema incorporate further distinctions among salaried employees (managers versus professionals), gradational distinctions among them and also among contractual nonmanual workers, and sector and size distinctions among proprietors (e.g., Hout and Hauser 1992). With the exception of property relations, the remaining criteria for class distinctions could well apply to the occupational structure of Soviet-era Russia. Thus, although the Erikson-Goldthorpe schema was developed in reference to advanced capitalist societies, we have adopted it for our analysis of intergenerational class mobility in Russia. There is sound precedent for this: Variations on the Erikson-Goldthorpe schema have been fruitfully employed in analyses of intergenerational class mobility in Eastern European countries (Ganzeboom, Luijkx, and Treiman 1989; Ganzeboom, DeGraaf, and

Robert 1990; Erikson and Goldthorpe 1992b; Wong 1992; Ishida, Müller, and Ridge 1995).

The Erikson-Goldthorpe schema is clearly superior to the broad distinctions characteristic of the more theoretically oriented accounts of state socialist class structure because its finer grain can capture more of the patterns in mobility as it is experienced. But the advantages of the Erikson-Goldthorpe schema are potentially diminished by the practice (common to all the studies just cited) of aggregating categories in advance, without theoretical or empirical justification.[4] In particular, we would expect that distinctive patterns of association obtain for professionals and managers, two classes that are usually collapsed into a single "salariat" or "service class."

To justify the practice of treating professionals and managers as the two components of a single class Erikson and Goldthorpe stress that both groups share a similar "service relationship" with their employers. Employers, be they individual owners of firms or top executives of large organizations, cannot easily monitor or control the work of managers and professionals; yet that work is particularly important to the success of the organization. Employers thus seek to induce cooperation, commitment, and loyalty by offering professionals and managers long-term guarantees that align their personal success with organizational success (Goldthorpe 2000:18).

The service relationship between professionals and employers, however, differs from that between managers and employers. Professionals apply highly specialized expertise in narrow tasks that require it; managers exercise broad authority delegated by the employer. Professionals do things that employers cannot do because they lack the necessary skill; managers do things that the employers could in principle do themselves but instead hire others to do, whether by choice or, in the case of collectively owned corporations and public sector organizations, by necessity. This difference between the service relationships of professionals and mangers justifies testing whether they have distinctive mobility patterns—not just in Russia, but elsewhere too. Moreover, Konrad and Szelenyi (1979), Parkin (1972, 1979), and Wright (1985) all emphasize the distinction between organizational authority and expertise as bases for different class positions in state socialist societies.

Differences in recruitment patterns provide further general justification in the Weberian tradition of class analysis. For the most part, professionals are recruited on the basis of their credentials, while managers get to their positions by demonstrating loyalty and performing well in responsible positions, either in their employers' firms or elsewhere. This suggests to us that the intergenerational reproduction of credential-based professions will be less direct, and mediated by universities, while the intergenerational reproduction of managers will have a larger network component. We expect this difference in recruitment to be important in Russia. Indeed, analysis of the school-to-work transition demonstrates the important role educational credentials have played in shaping access to professional jobs in both Soviet and post-Soviet eras (Gerber 2003). During the Soviet era, Communist Party connections were important for advancement through managerial ranks (Gerber 2000b, 2001b). Since 1991, "crony capitalism" and embedded favoritism have been integral to the transition to private ownership and control (Blasi, Kroumova, and Kruse 1997). Both practices draw attention to the role of social capital in managerial recruitment and placement. In light of these considerations, we believe managers should be distinguished from professionals within the upper and lower "salariat" classes (Erikson-Goldthorpe classes I and II).[5] We test the utility of this disaggrega-

[4] Hout and Hauser (1992), in their re-analysis of the CASMIN data (which include data from Hungary and Poland), found that using the full Erikson-Goldthorpe schema revealed more stratification than the reduced versions did. Replying to Hout and Hauser, Erikson and Goldthorpe (1992a) endorsed the idea of using their fully elaborated schema but expressed concern that differences among nations in the details of their classification systems would make fine distinctions unreliable for cross-national comparisons.

[5] Studies of earnings and of educational attainment in post-Soviet Russia have found that professionals and managers are distinct in these respects (Gerber and Hout 1995, 1998; Gerber 2000a). Hout, Brooks, and Manza (1995) gained empirical leverage in their class analysis of US presidential elections by distinguishing professionals from managers.

tion empirically in a preliminary step in our analysis.

We apply the Erikson-Goldthorpe schema to both Soviet-era and post-Soviet Russia not only because it is the conventional schema used in mobility studies (though that would be a sufficient rationale for doing so), but also because we believe there are good theoretical reasons why the schema is suitable, despite the emphasis many observers place on the differences between state socialist and capitalist systems. As our discussion of the need for loyalty mechanisms and the reliance on credentials and social capital for recruitment in Soviet-era and post-Soviet Russia implies, roughly the same considerations can influence the decision-making of employers in both state socialist and modern capitalist societies. Researchers should not draw distinctions in the incentives facing capitalist and state socialist organizations too sharply. While it may be true that loyalty to the state played a larger role in the allocation of rewards under state socialism, the socialist state also placed a premium on productivity. Even if their organizations were not faced with bankruptcy, directors of state socialist firms could face censure and dismissal for underperformance. They thus had every reason to seek out more productive employees and to align the goals of lower managers and professionals with those of the organization. In this sense, their situation hardly differed from that of managers within capitalist systems who work in large corporations or public organizations. In short, distinctions between the considerations of state socialist and capitalist employers are distinctions of degree rather than kind.

CHANGE OVER TIME?

Market Transition and Stratification Processes

The dramatic transformation of the Russian economy after 1991 included many elements likely to affect workers' careers. The "shock therapy" of price, trade, and currency liberalization followed by the privatization of many state-owned enterprises produced spiraling inflation, recession, unemployment, and unprecedented inequalities (Blasi, Kroumova, and Kruse 1997; Gerber and Hout 1998). Even workers who kept their jobs faced endemic wage arrears and involuntary furloughs (Desai and Idson 2000; Earle and Sabirianova 2002). Displaced workers coped by engaging in barter and primitive production of either food or handicrafts or both (Burawoy and Krotov 1992; Burawoy 1997). As structural and institutional changes rippled through the economy, job mobility increased, bringing with it occupational mobility (Gerber 2002).

Even as the economy shrank, new opportunities arose. Entrepreneurs found ways to get modern consumer goods to market. The sharp devaluation of the ruble in August 1998, import substitution, high oil and gas prices on the world market, improved tax collection, and controls over capital flight have reversed the long crisis in Russia's economy. Foreign investment, production indicators, and growth rates have all risen since 1999. Wage arrears have been paid up at most enterprises; incomes and spending have rebounded.

From the vantage point of 2004, it appears that Russia's market transition has survived its crisis period. Clearly, there have been winners and losers; inequality is very high by historical and international standards (Gerber and Hout 1998). A lively literature examines the impact of market transition on inequality of opportunity and/or outcome in Russia, China, and other former socialist societies (Nee 1989, 1996; Rona-Tas 1994; Bian and Logan 1996; Xie and Hannum 1996; Zhou, Tuma, and Moen 1997; Gerber and Hout 1998; Cao and Nee 2000; Gerber 2000b, 2001a, 2001b, 2002; Zhou 2000; Walder 2002). With rare exceptions (e.g., Gerber 2000a), this literature has focused almost exclusively on intragenerational processes, such as earnings determinants, elite formation, and job mobility.

Why have researchers ignored the potential impact of market transition on *intergenerational* stratification? First, the concerns of the literature on postsocialist stratification have been largely shaped by the debate over market transition theory (Nee 1989, 1996). This theory deals solely with intragenerational bases of inequality and does not predict rapid changes in the intergenerational transmission of status.

Second, the bases for making predictions are not self-evident. The very nature of reform-based structural dislocations and institutional changes suggests they can rapidly affect outcomes that depend on the labor market or political markets. But intergenerational mobility evolves over time through long processes of education, acculturation, and an extended sequence of transitions through jobs (Hout 2003). Thus, changes in the origin-destination association during a decade of market transition are likely to be subtle.

Third, empirical studies of change over time in mobility indicate that where change does occur, it is incremental (Ganzeboom, Luijkx, and Treiman 1989; Vallet 1999; Breen and Jonsson 2003) and may involve cohort replacement as well as period effects (Hout 1984, 1988). Even nineteenth-century France, which also experienced rapid institutional change, exhibited little movement in origin-destination association, even though structural mobility was substantial (Sewell 1982).

Finally, other stratification processes in Russia have exhibited a surprising degree of stability. The benefits of education and Communist Party membership, the effects of family background on the (conditional) odds of completing secondary and entering tertiary education, and the association between education and first occupational class have changed little, if at all (Gerber and Hout 1998; Gerber 2000b, 2000b, 2001b, 2003; but see Brainerd 1998). Why expect mobility patterns to change rapidly when changes in these areas have been incremental?

In fact, we hypothesize that market transition produces a tightening up of the mobility regime (*increased* OD association) because it increases class-based intragenerational inequality and leads to intragenerational job mobility involving regression toward origins. Imperfect as it was at realizing its goal, the Soviet government sought to minimize, if not eliminate, class-based inequalities. As a result, during the Soviet era officials kept income differentials among classes low.[6] This made class mobility a game with relatively low stakes. In Sweden, relatively low wage differentials reduced parents' drive to give their children advantages (Erikson and Jonsson 1998); it may well have worked that way in Russia, too. As overall wage inequalities spiraled following the Soviet collapse, the Russian government relinquished central control over wages and retreated from regulating the labor market. Coupled with the rise of a new advantaged class—proprietors—the government's withdrawal from the labor market should have resulted in an increase in class differences in earnings as well as a simultaneous increase in overall earnings inequality. Thus, market transition would increase the premium on current class position as a determinant of life chances, which would lead to a tightening up of the intergenerational mobility regime.

The main reason Soviet wage differences were so low was the official premium for skilled manual work. In the late 1980s the earnings of upper blue-collar workers exceeded those of managers and some professionals. As early as 1995, proprietors were emerging as a rich class; and managers (and, to a lesser extent, professionals) were gaining on upper blue-collar workers (Gerber and Hout 1998). We suspect that a wage hierarchy typical of market economies has emerged since then, as privatization and market institutions have taken firmer hold in Russian society. Recent studies suggest that Russian professionals and managers have less risk of job loss (Gerber 2002) and wage arrears (Desai and Idson 2000; Earle and Sabirianova 2002) than manual workers. Observing these developments, Russians, particularly those with elite origins, might easily conclude that class has become a more significant determinant of life chances. Thus, the simultaneous growth of inequality and of its basis in occupational class intensifies the competition for access to favorable class positions.

Because downward mobility has had more serious consequences after the market transition, higher-origin Russians who were downwardly mobile prior to the transition can be expected to use every resource at their disposal to move back up the class hierarchy by changing jobs. In doing so, they displace lower-origin Russians

[6] Class-based wage inequalities reached new lows during the 1970s and 1980s, when the average wages of skilled manual workers equaled or even exceeded those of many professionals. Of course, as one anonymous reviewer reminds us, there were other, hidden inequalities in Soviet-era Russia. Party leaders, top managers, and some professionals enjoyed greater access to scarce goods, quality housing, and other privileges not reflected in income differentials (Matthews 1972, 1978; Szelenyi 1983). But in the USSR these privileges applied only to a very narrow group of elites, not to broad occupational classes.

who had been upwardly mobile prior to the transition. If, as we expect, the competition favors those from advantaged origins, the result will be a regression toward origins: Russians who were upwardly mobile in the Soviet era would return disproportionately to their lower-origin classes, and vice versa. If intragenerational mobility tends to follow this regression-toward-origins pattern, the intergenerational mobility regime should tighten up because destinations will more closely resemble origins.

A study of job mobility patterns in Russia from 1991 to 1998 presents findings consistent with the regression-toward-origins hypothesis (Gerber 2002): College-educated Russians experienced lower rates of job loss and job mobility (as defined by average occupational earnings), but higher conditional rates of upward and lower conditional rates of downward mobility. The opposite pattern obtained for those with less than secondary schooling. Other published results support our hypothesis that higher origins are an advantage in the competition for advantaged class positions: The offspring of professionals are more likely to enter the most privileged class in terms of earnings (proprietors, with or without employees), the rise of this class being clearly a result of market transition (Gerber 2001a). But these studies do not directly test either the tightening-up hypothesis or the regression-toward-origins explanation.

Period Effect versus Cohort Replacement

Shock therapy, the deregulation of wages, economic crisis, and the elimination of many other state protections increased competition for occupational advantages. Competition, in turn, induced an intragenerational mobility pattern that reduced differences between peoples' current occupations and their social origins. That is our argument. But cohort replacement could also be important. Over the decade of the 1990s, the first post-Soviet cohort entered the labor force while the cohort educated during World War II and immediately after (i.e., people born 1925–1940) retired. If the association between origins and destinations is stronger for the cohort moving in than for the one exiting, then all else being equal, the origin-destination association for the workforce as a whole will rise even if the institutional factors we have stressed were not important for mobility. To be sure of the validity of our interpretation of the change in OD association as a period effect, our analysis therefore separates the period effects (which contain institutional factors pertinent to all cohorts) from cohort effects (which are important for replacement).

DATA AND METHODS

Our data come from three pretransition and three posttransition surveys that we pooled to form a cumulative data file. We exclude respondents who at the time occupation was measured were under 25 years old or over retirement age (55 years for women, 60 years for men). We also exclude cases with missing data on current occupation, age, or gender (Table 1). A reasonable date marking the start of Russia's transition is January 1992, when the newly independent Russian government introduced the sweeping reforms that came to be described as "shock therapy." The three pretransition surveys are the Russian respondents in Treiman and Szelenyi's Social Stratification in Eastern Europe (SSEE) survey (Treiman 1994), the Comparative Class Structure and Consciousness Project (CCSCP) (Hout, Wright, and Sanchez-Jankowski 1992), and the 1992 International Social Survey Programme (ISSP). The SSEE was fielded in 1993, too late to be unambiguously pretransition and too early to show much in the way of a transition effect. The survey did, however, ask respondents what their occupation had been in 1988. We use this measure of the respondents' occupation, not their current occupation, to supplement our other pretransition observations. The 1992 ISSP was fielded in February in Russia. Although this is technically one month into the transition, it is unlikely that any major changes in the occupational structure had occurred by then, so we feel comfortable treating the 1992 ISSP data as pretransition. The data from the posttransition period come from the 1998 Survey of Employment, Income, and Attitudes in Russia (SEIAR) (Gerber 1999); the Russian respondents from the 1999 ISSP; and the 2000 Survey on Education and Stratification in Russia (Gerber 2000c).

Each survey employed standard multistage sampling procedures. Details regarding sampling, fieldwork, quality control, and response

Table 1. Data Sources

Survey	Principal Investigator(s)	Data Producer	Date Respondent's Occupation Measured	Valid N
Pretransition Data				
Social Stratification in Eastern Europe (SSEE)[a]	Treiman and Szelenyi	All-Russian Center for the Study of Public Opinion (VTsIOM)	End of 1988	2,928
Comparative Class Structure and Consciousness Project (CCSCP)	Wright, Hout, and Sanchez-Jankowski	Institute of Sociology, Russian Academy of Sciences	February 1991	1,400
International Social Survey Programme (ISSP), 1992		VTsIOM	February 1992	1,061
Subtotal				5,389
Posttransition Data				
Survey of Employment, Income, and Attitudes in Russia (SEIAR)	Gerber	VTsIOM	January–March 1998	2,202
International Social Survey Programme (ISSP), 1999		VTsIOM	February 1999	586
Survey on Education and Stratification in Russia, 2000	Gerber	VTsIOM	September–November 2000	2,087
Subtotal				4,875
Total				10,264

[a] Treated as pretransition data even though the survey was conducted in 1993 because the respondents' 1988 occupation, not their occupation at the time of the survey, is used.

rates are available in the documentation accompanying the original data sets. All surveys except for the CSSCP are nationally representative samples (with the exclusion of remotely populated regions of northern and eastern Russia and the war-torn republics of the Caucasus) of Russians aged 16 and over conducted by the All-Russian Center for the Study of Public Opinion (VTsIOM).[7]

The CCSCP project differs in several ways. First, it sampled those 18 and older from European Russia, thereby excluding respondents east of the Ural Mountains. The CCSCP also used a slightly different origin question (and has correspondingly less missing data). It asked respondents about the occupation of the "main earner" in their household when they were growing up. The other surveys asked specifically about the father's occupation when respondents were either 14 or 16. In addition, the CCSCP used an occupational classification based roughly on US Census categories, while all the other surveys used the 1988 International Standard Classification of Occupations

[7] As an anonymous reviewer noted, it would be appropriate to correct for the differences in sampling designs applied by the various surveys. This could be done by applying specific design weights to each survey. The use of design weights would reduce our effective sample size to reflect the impact of clustering. Unfortunately, we do not have sufficient information about the details of primary or secondary sampling units (PSUs, SSUs) needed to make informed estimates of appropriate weights. In any case, we reestimated all our models after applying an overall design weight of .70. This value, used by the Current Population Survey for most variables, is probably too low, because VTsIOM refreshes its sample more often, rotating SSUs each wave and all except self-selecting PSUs each year. This design factor affects none of our findings. We prefer to analyze the raw cell counts, given that the choice of a particular design factor would be arbitrary. But we determined that the design factor would have to be extremely low (.41) to negate our key findings.

(ISCO88). We converted the CCSCP categories to ISCO88 categories prior to converting the ISCO88 categories to our 11-class (plus "missing") extension of the Erikson-Goldthorpe schema. Finally, the CCSCP survey was carried out by the Institute of Sociology of the Russian Academy of Sciences, which may have used different procedures than VTsIOM, the organization that carried out the other five surveys. These disparities may introduce some biases of an unknown direction and magnitude. More generally, random survey-to-survey fluctuations may account for apparent differences over time. In order to rule out this artifactual interpretation of change from the pre- to the posttransition period (P), we estimate our models on the full year-by-origins-by-destinations (YOD) table rather than the collapsed period-by-origins-by-destinations (POD) table. This permits us to test formally for year-to-year (survey-to-survey) fluctuations in the OD association within each period and across both periods. We use equality constraints on the years within each period to test for stability within and change across period.

The grist of a mobility analysis is the specific set of categories that define the distinctions deemed to be significant. We arrived at our set of origin and destination categories following a series of preliminary statistical tests described in the Appendix. To show how our preferred categories relate to the full set of Erikson-Goldthorpe categories, we display both sets in Table 2. Note that our specification distinguishes managers from upper and lower professionals (for reasons given earlier). Empirically, this separation is justified because managers and professionals are not isomorphic in Goodman's (1981) sense and therefore should not be combined: doing so actually *decreases the observed OD association by 6 percent* (see the Appendix).

Large tables of the sort we work with here are valuable for all the information they contain. But comparing each cell for men and women or for each period is a very inefficient method of deciding whether gender or period affects mobility. So we also use a one-degree-of-freedom test independently developed by Erikson and Goldthorpe (1992b) and Xie (1992). We refer to it by the term Erikson and Goldthorpe used: the uniform difference, or unidiff for short. The unidiff is defined in terms of how a log-odds-ratio from one table ($\ln\theta_{ij|k}$) relates to that same log-odds-ratio in another ($\ln\theta_{ij|k}'$).[8] Suppose that the log-odds-ratios in the k' table differ from the corresponding log-odds-ratio in the k' table by some proportion so that each log-odds-ratio in the k' table could be obtained by

[8] Where i (= 1, . . ., I–1) indexes origins, j (= 1, . . ., J–1) indexes destinations, and k and k' represent two different mobility tables, e.g., the one for men and the one for women or the one for pretransition and the one for posttransition.

Table 2. The Erikson-Goldthorpe Class Schema: Full and Preferred Versions

	Full Class Schema:		Preferred Class Schema:
I(P).	Professionals (upper)	I(P).	Professionals (upper)
I(M).	Managers (upper)	II(P).	Professionals (lower)
II(P).	Professionals (lower)	I(M)/ II(M).	Managers (upper and lower)
II(M).	Managers (lower)		
IIIa.	Routine nonmanual (upper)	IIIa.	Routine nonmanual (upper)
IIIb.	Routine nonmanual (lower)	IIIb.	Routine nonmanual (lower)
IVab.	Proprietors (with or w/o employees)[a]	IVab.	Proprietors (with or w/o employees)
IVc.	Farmers[a]		[not observed]
V.	Supervisors of manual workers	V/VI.	Manual supervisors/skilled manual
VI.	Skilled manual workers		
VIIa.	Unskilled manual, nonagricultural	VIIa.	Unskilled manual, nonagricultural
VIIb.	Unskilled manual, agricultural	VIIb.	Unskilled manual, agricultural
X.	Missing[b]	X.	Missing[b]

[a] Class IVab is not an origin category because self-employment was illegal during the Soviet period. We observed no cases in IVc in either the origin or destination distribution.

[b] "Missing" is not a destination category.

multiplying the corresponding log-odds-ratio in the k table by a constant:

$$\ln\theta_{ij|k'} = \phi \ln\theta_{ij|k} \quad (1)$$

where ϕ is the "unidiff" constant (equal to the proportional difference between $\ln\theta_{ij/k'}$ and $\ln\theta_{ij/k}$ plus 1). Testing whether the association between origins and destinations differs for k and k' then boils down to testing whether ϕ is significantly different from 1. The appeal of this approach is that the unidiff model uses just one more degree of freedom than the corresponding model of no three-way interaction among origins, destinations, and the third variable (represented by the distinction between k and k'). This statistical efficiency gives unidiff far more statistical power than the usual model of three-way interaction, which uses $(I–1) \times (J–1) – 1$ more degrees of freedom.

We use both the BIC statistic (Raftery 1995) and the likelihood-ratio test (which in this case is just the difference in L^2 for nested models) to compare the fit of alternative models. Generally, we regard BIC as definitive when we test a large number of models using a relatively large sample, a strategy that increases the risk that substantively trivial parameters will appear to be significant in a likelihood-ratio test. However, Wong's (1994a) Monte Carlo study showed that with medium to large sample sizes, differences in BIC of fewer than 5 points should be viewed as indeterminate. We therefore rely on the likelihood-ratio test whenever the difference in BIC falls below 5.

ANALYSIS PLAN

Our analysis has five steps. First, we describe men's and women's mobility in each period. Second, we assess the significance of gender differences in the association between origins and destinations within each period.[9] If we find no gender differences we collapse the male and female tables, which simplifies further analyses. In the same step, we test for year-to-year fluctuations in association within each period to see whether our treatment of the two groups of years as distinctive periods is empirically justified. The results of this second step set up the third: our assessment of change from pre- to posttransition in the origin-destination association. Fourth, we test whether the change we observe between periods reflects institutional changes associated with the transition out of state socialism or the demography of cohort replacement. Finally, we test our regression-to-origins hypothesis using data on the intragenerational mobility of people from different origins.

RESULTS

Gross Mobility in the Soviet and Post-Soviet Eras

For descriptive purposes, we define four types of mobility by distinguishing upward from downward mobility and long from short moves. To make these distinctions we consulted the category scores described below.[10] We defined individuals as upwardly mobile if their current class has a score higher than their origin class has; downwardly mobile if their current class has a lower score; and immobile if they currently work in their origin class or in a different class that has the same score that their origin has. A person's mobility is long-range if the origin and destination classes have scores that differ by more than the standard deviation of the scores (a difference of .33, given our identifying restrictions) and short-range if they differ by less.

Mobility was the norm in Russia both before and after the economic transition (see Table 3). In the late Soviet period, 76 percent of Russian men and 85 percent of Russian women were in a class different from the one they grew up in. Since the transition, Russian men and women experienced significantly more downward mobility (both long- and short-range) and less long-range upward mobility. Short-range upward mobility also grew slightly for women, who experienced a sharper drop in long-range

[9] Marshall, Sydorenko, and Roberts (1997) found no gender difference in origin-destination association at the end of the Soviet period, but their sample was small and a difference may have emerged subsequently, because Russia's labor market transition has affected men and women differently (Gerber and Hout 1998; Gerber 2002).

[10] We show how the mobility types correspond to the cells of the mobility table in Table A3, which also provides the score for each class.

Table 3. Mobility Rates by Gender and Period

	Mobility Type				
Gender/Period	Long-Range Upward	Short-Range Upward	Immobile	Short-Range Downward	Long-Range Downward
Men, %					
Pretransition (N = 2,527)	29.4	28.9	24.1	11.2	6.3
Posttransition (N = 2,239)	24.2	28.9	25.5	13.7	7.7
Change	–5.2	0.0	1.4	2.5	1.4
Women, %					
Pretransition (N = 2,862)	47.2	27.0	14.5	8.6	2.8
Posttransition (N = 2,636)	40.0	29.5	14.6	12.1	3.9
Change	–7.2	2.5	0.1	3.5	1.1

upward mobility than men had. Upward moves still outnumbered downward moves, but the ratio of upward to downward moves decreased from 3.3 to 2.5 for men and 6.5 to 4.4 for women. All of these shifts amount to no net change in mobility overall, but Russians' mobility prospects clearly declined as the transition took hold.

Downward mobility surged because many desirable jobs disappeared in the transition. Market pressures and the demise of the state sector caused the top classes to shrink; the proportion of men and women in classes I and II fell by 7 or 8 percentage points (for details, see Table A2 in the Appendix). Growth came in unskilled jobs in production and services (classes VIIa and IIIb, respectively) and self-employment. The rising opportunity for self-employment was the only good news, and this was largely a boon for men. With these shifts in the occupational structure, the upward bias in structural mobility was significantly weaker in the posttransition period, sparking the increase in downward mobility. As we explain below, rising class barriers accentuated the difficulties faced by Russians from modest and lower-class origins.

Gender and Year-to-Year Differences in Mobility Patterns

Studies of other countries have often found that, though the destinations of men and women are distinct, the association between origins and destinations does not differ by gender. Table 4 presents our test of that generalization for Russia, as well as a test of our assumption that essentially the same mobility patterns obtained

Table 4. Testing for Differences in the Association between Origin and Destination within Time Period

Model	L^2	*df*	*p* vs. [1]	BIC	Model *p*	D
A. Pretransition						
1. {GYO}{GYD}{OD}[a]	335.63	320		–2,414	.263	.075
2. [1] + {YOD}	188.28	192	.116	–1,461	.562	.054
3. [1] + unidiff {YOD}	329.12	318	.039	–2,403	.322	.074
4. [1] + {GOD}	258.44	256	.125	–1,941	.445	.061
5. [1] + unidiff {GOD}	335.29	319	.557	–2,406	.254	.075
B. Posttransition						
1. {GYO}{GYD}{OD}[a]	371.16	320		–2,346	.026	.075
2. [1] + {YOD}	240.49	192	.418	–1,390	.010	.067
3. [1] + unidiff {YOD}	369.90	318	.532	–2,331	.024	.085
4. [1] + {GOD}	292.66	256	.105	–1,881	.057	.074
5. [1] + unidiff {GOD}	370.96	319	.651	–2,338	.024	.086

Note: G = gender; Y = year/survey; O = origin; D = destination.
[a] Preferred models.

within each period.[11] For this set of tests, we examine the gender-by-year-by-origin-by-destination (GYOD) tables for each of the two periods. Because our interest is in variation by sex and by year in the OD association, we fit the GYO and GYD marginals in all models. We start with a model that constrains the OD association to be invariant by both gender and year. We then see if relaxing either (or both) of those constraints produces a better-fitting model; if so, we must conclude that the OD association varies by gender or year within that particular period.

The results demonstrate that within each period the association between origins and destinations was the same across years and for men and women (Table 4). In every case, the BIC statistic unambiguously indicates that model 1 offers the best fit. For the pretransition period, the unidiff specification of year-to-year change (model 3) leads to a significant improvement of fit according to likelihood-ratio test, but the improvement clearly is not sufficient according to BIC. Moreover, model 1 already fits the data acceptably well using the conventional criterion. Thus, we are confident in our conclusion based on the BIC statistic that the OD association does not vary by year within either period. Among the posttransition surveys, there is no inconsistency between alternative measures of model fit. Although model 1 does not fit the data at $p > .05$, we think this has more to do with the sparseness of the data than with anything of substance; of the 486 cells 67 have frequencies of zero and 234 have frequencies below 5.[12] We thus can proceed further on the assumptions that the origin-destination association does not vary by gender in Russia and that the years 1988–92 and 1998–2000 represent cohesive periods within which the origin-destination association did not vary.

How Destination Depends on Origin in Russia and How It Changed

To determine the best specification of how destination depends on origin in Russia and to assess whether that association has changed due to market transition, we analyze the 9×9×6 origin-by-destination-by-year table, fitting the OY and DY marginals for all models. We begin with the "core model" from the Comparative Analysis of Social Mobility in Industrial Societies (CASMIN) project, which has been a benchmark for mobility analysts since Erikson and Goldthorpe (1987, 1992b) introduced it. We then develop three versions of an alternative, "hybrid" model that combines a vertical conception of structure of opportunity and mobility with two horizontal inheritance parameters.

For each specification, we begin with a baseline model that assumes constant association over time. We then test four different patterns of change: a unidiff pattern of change from year to year (where the annual change in all parameters representing the OD association is proportional to a multiplicative scaling factor, ϕ_y, associated with each particular year), a unidiff pattern of change across period (equivalent to constant ϕ_y parameters within each period but different values in the two periods), full heterogeneity of association across year (all association parameters freed to vary from year to year), and full heterogeneity of association across period (all association parameters constrained to homogeneity across year within period, but free to vary across period). The unidiff models treat the pattern of association as con-

[11] We combine these steps for economy of presentation, not for any substantive or methodological reason.

[12] Haberman (1977) derives an important result about sparse tables like ours. Sparseness inflates L^2 (and, to a lesser extent, χ^2). Fortunately, the upward bias in L^2 is constant across all models for the same data, so comparisons of two nested models are unbiased because the bias in both is removed by subtraction. Clogg and Eliason (1987) show that if adding .5 to every cell in the table yields a model that fits the data, the lack of fit can be attributed to sparseness. This applies here: When we add .5 to every cell in the posttransition GYOD table, the L^2 for model 1 is 307.07 ($p = .689$). We do not report the results for models estimated on tables with .5 added to each cell because the only consequence of this move is to improve the overall fit of all models—none of our substantive conclusions is affected. However, adding .5 biases all parameter estimates toward zero. Data so altered appear more consistent with independence than the raw data do (Clogg and Eliason 1987). For example, in our data, adding .5 to each cell lowers the baseline L^2 by 6 percent. In this instance, the costs of adding .5 outweigh the benefits. Adding a smaller quantity, such as .1, does not improve fit above the conventional .05 level.

stant but free the magnitude of association to vary over time, while the fully heterogeneous models allow both the pattern and the magnitude of association to vary. The latter are thus less parsimonious (i.e., they consume more degrees of freedom). We test for both year-to-year and period changes in order to assure ourselves that the changes we detect between the pre- and posttransition periods genuinely reflect the impact of market transition as opposed to a more gradual, linear process of change over time, a pattern of trendless fluctuation, or survey-to-survey sampling variations that might be mistaken for a period effect. Although the results in Table 4 suggest stability in the OD association within period, they are based on only one specification of the OD association.

The CASMIN core model does not fit the data by standard criteria, but the negative BIC value indicates that it is preferred over the saturated model (Table 5, model A1). The best-fitting specification of change over time, for the core model as well as for the other four models, is unidiff association across *period* (model A3). Keeping in mind that, given our large sample, a change in BIC of fewer than five points is indeterminate (Wong 1994a), we find that model 3 fits no better or worse than the baseline. Thus, we turn to the likelihood-ratio test, which definitively favors the period change model.[13] The three other specifications of change over time fit worse than the baseline model using the BIC criterion. Therefore, model 3 is our preferred model for group A, as it is in each of the other three groups of models. Each

[13] In fact, with our sample size the equivalence of BIC implies statistically significant improvement of the model, since ln(10,264) = 9.24. For a one degree of freedom test, a reduction in L^2 of 3.84 (the threshold for significance at $p < .05$) produces an increase in BIC of 5.4. An increase in BIC of 4 points (the threshold for BIC's indeterminacy) requires a reduction of L^2 of at least 5.24.

Table 5. Fit Statistics, Selected Models for the Origin-by-Destination-by-Year Table

Number and Description	L^2	*df*	*p* vs. [1]	BIC	Model *p*	D
A. Core Social Fluidity Model						
1. Baseline (no change)	512.8	377	—	–2,969	<.001	.072
2. Unidiff association by year	500.6	372	.032	–2,935	<.001	.071
3. Unidiff association by period	505.0	376	.005	–2,968	<.001	.071
4. Heterogeneous association by year	466.6	342	.098	–2,692	<.001	.067
5. Heterogeneous association by period	496.9	370	.026	–2,921	<.001	.070
B. Unconstrained Quasi-RC (No Equality Constraints on RC Scores)						
1. Baseline (no change)	455.9	367	—	–2,934	.001	.068
2. Unidiff association by year	442.6	362	.021	–2,901	.002	.067
3. Unidiff association by period	446.3	366	.002	–2,934	.003	.068
4. Heterogeneous association by year	435.3	352	.149	–2,816	.002	.066
5. Heterogeneous association by period	444.9	364	.012	–2,917	.002	.068
C. Quasi-RC with Equal RC Scores Plus Two Additional Equality Constraints						
1. Baseline (no change)	473.9	375	—	–2,990	<.001	.070
2. Unidiff association by year	459.8	370	.015	–2,958	.001	.069
3. Unidiff association by period[a]	464.5	374	.002	–2,990	.001	.069
4. Heterogeneous association by year	451.6	360	.101	–2,873	.001	.068
5. Heterogeneous association by period	463.0	372	.013	–2,973	.001	.069
D. Homogenous Quasi-RC with Hout and Hauser Scores						
1. Baseline (no change)	552.2	381	—	–2,967	<.001	.078
2. Unidiff association by year	538.4	376	.017	–2,934	<.001	.078
3. Unidiff association by period	543.7	380	.004	–2,966	<.001	.077
4. Heterogeneous association by year	529.3	366	.088	–2,851	<.001	.076
5. Heterogeneous association by period	541.9	378	.017	–2,949	<.001	.077

Note: {OY} and {DY} are fitted in all models. N = 10,264.
[a] Preferred model.

of the core model's seven parameters was stable from 1988–92 and from 1998–2000, but they all changed by the same proportion between these two periods.

Despite its widespread use by researchers, the CASMIN core model has a number of undesirable properties (Hout and Hauser 1992). Most serious is that the hierarchy terms in the core model present a rather undifferentiated picture of how destinations depend on origins; the model limits origin effects to a single increment (per origin level) to the odds on moving up the social hierarchy. Substantively, the model implies that hierarchy effects work only by providing a safeguard against downward mobility. Even more radical, the diagonal cells are the only nonzero entries in the table of log-odds-ratios implied by the model (see Hout and Hauser 1992).

A more flexible alternative that better reflects most understandings of the relationship between origins and destinations is the so-called RC-II model of Goodman (1979). This model posits an unobserved latent variable made manifest by the categories of the class schema. With appropriate identifying restrictions, one can estimate simultaneously the scores that rank each class and an association parameter representing the slope of the line relating origin scores to the log-odds of being in the higher-scoring destination of a pair of destinations one point apart on the scale (see Goodman 1979; Hout 1983; Wong 1992, 1994b).

The RC-II model is appropriate for modeling the association of any two ordinal variables. Mobility tables are distinctive because of the correspondence between row and column categories. We adopt two standard adjustments that mobility researchers use to take this correspondence into account. First, we give special treatment to diagonal cells (i.e., cells in which i and j refer to the same class or occupational category); these tend to have higher counts in every empirical study we know of.[14] To capture this inheritance, researchers augment RC-II models and other unsaturated models with a dummy variable that distinguishes diagonal cells from other cells. We use such a variable, denoted D_{ij}, which equals 1 if $i < 9$ and $i = j$, zero otherwise. Often occupational inheritance is higher for agricultural workers than for other classes. We also, therefore, include a dummy variable, D_{88}, for $i = j = 8$.

With the addition these two parameters representing association along the diagonal, we obtain our first hybrid model for the OD association, an unconstrained quasi-RC (QRC) model.[15] Adding year to the model and specifying change over year in unidiff form:

$$\ln F_{ijy} = \lambda_0 + \lambda_i^{\mathrm{O}} + \lambda_j^{\mathrm{D}} + \lambda_y^{\mathrm{Y}} + \lambda_{iy}^{\mathrm{OY}} + \lambda_{jy}^{\mathrm{DY}} + (\theta\mu_i\, v_j + \delta_1 D_{ij} + \delta_2 D_{88})\phi_y, \tag{2}$$

which is identified by the following restrictions:

$$\Sigma_i\lambda_i^{\mathrm{O}} = \Sigma_j\lambda_j^{\mathrm{D}} = \Sigma_y\lambda_y^{\mathrm{Y}} = \Sigma_i\Sigma_y\lambda_{iy}^{\mathrm{OY}} = \Sigma_j\Sigma_y\lambda_{jy}^{\mathrm{DY}} = \Sigma_i\mu_i = \Sigma_j v_j = 0,\ \Sigma_i\mu_i^2 = \Sigma_j v_j^2 = 1, \text{ and } \phi_1 = 1.$$

The θ parameter is like a logit regression coefficient, so if $\theta = 0$ then destinations are independent of origins. The ϕ_y parameter is the unidiff parameter: as can be seen, it requires the change in the three association parameters, θ, δ_1, and δ_2, to be proportional. Constraining all ϕ_y to equal one yields the baseline model of no change in OD association over time. Imposing the constraints, $\phi_1 = \phi_2 = \phi_3 = 1$ and $\phi_4 = \phi_5 = \phi_6$, yields a model of unidiff change across *period*. We specify heterogeneous association across year by dropping ϕ_1 from (1) and subscripting θ, δ_1, and δ_1 with y—that is, estimating separate association parameters for each year. We further impose the appropriate time constraints on these estimates—for example, $\theta_1 = \theta_2 = \theta_3$ and $\theta_4 = \theta_5 = \theta_6$ —to specify heterogeneous association across period. Estimates of the row and column scores from our preferred specification of (1) (model B3 in Table 5) are presented in Figure 1, part A.

[14] Row and column categories are usually arrayed in the same order so that $i = j$ when the categories correspond. Our study is unusual in that our "missing" origin has no corresponding destination and our "self-employed" destination has no corresponding origin. We put these two exceptional categories at the end; thus, only if $i < 9$ does $i = j$ imply that origin and destination classes are the same.

[15] As Wong (2001) notes in his discussion of several families of related models, this specification is called PARAFAC RCL(1) in the psychometric literature.

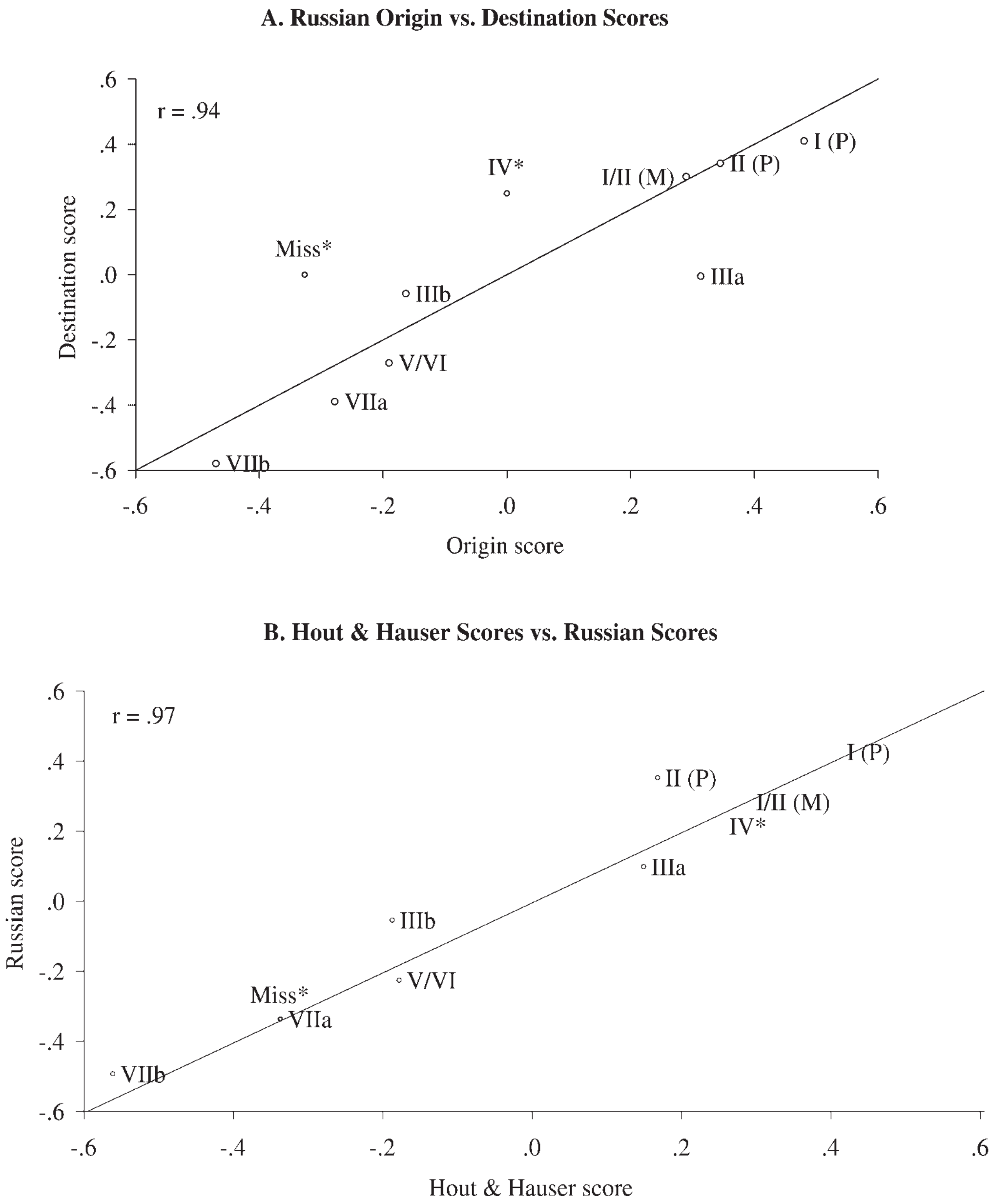

Figure 1. Comparisons between Origin and Destination Scores Estimated from Russian Data and between Homogeneous Row and Column Scores for Russia and the Scores Used by Hout and Hauser (1992)

Researchers who use scores often constrain origin and destination scores to be the same in the interest of parsimony and ease of interpretation (that is, $\mu_i = \nu_j$ for $i = j$). The Russian case is complicated by the extra, "missing" origin and by the lack of a self-employed origin. We achieve a partial symmetry model by applying equality constraints to all equivalent row and column scores—that is, $\mu_i = \nu_j$ for $i < 9$ and $i = j$. In addition, we constrain the scores of our two asymmetric classes (missing origins and class IV destination) to equal those of the classes with the closest scores (see Figure 1A), VIIa and I/II(M), respectively—that is, $\mu_9 = \mu_7$ and $\nu_9 = \nu_3$. This set of equality constraints yields a homogenous quasi-RC (QRC) model, which uses eight fewer degrees of freedom than the unconstrained QRC model.[16]

The unconstrained QRC model is much less parsimonious than the core social fluidity model; thus, the BIC statistic favors the latter (compare the equivalent models in panels A and B of Table 5.)[17] However, the homogenous QRC model achieves nearly the same parsimony as the core model, and its fit is clearly superior using the BIC criterion (compare panels A and C of Table 5). Moreover, comparing the BIC statistics of equivalent models in panels B and C reveals that our equality constraints fit the data quite well. For both homogenous and unconstrained versions of the QRC model, the best-fitting specification allows unidiff change across period in the OD association: The BIC statistics cannot distinguish between models 1 and 3, but the change-in-L^2 tests clearly favor 3. Thus, model C3 is our preferred model of mobility in contemporary Russia.

Although the homogenous quasi-RC model fits the Russian data better than the core model, the wide application of the core model imposes a higher standard of preference on any model that would supplant it. While the structure of the homogenous QRC model is very general, it is more closely tailored to Russia than the core social fluidity model because its scores are estimated on the Russian data. A fairer test would be to assess the homogenous QRC model using scores derived from prior applications to other countries. We therefore inserted the scores Hout and Hauser (1992) reported for the CASMIN data. The Hout and Hauser scores resemble the scores estimated from the Russian data (see Figure 1, part B); their correlation is .97.

By necessity, the imported scores do not fit as well as the scores estimated from the data because the latter are optimized to the data. To compensate, the model with imported scores uses six fewer degrees of freedom. The difference of roughly 80 points in L^2 is surprisingly large, given the high correlation between the two sets of scores. Nonetheless, for the baseline models and the models with unidiff specifications of change over time, the BIC criterion offers no grounds for choice between the core model and the homogenous QRC model with imported scores: the difference in BIC of only 2 points indicates that the two models fit equally well. The core model performs much more poorly using the less parsimonious specifications of change over time, but these specifications generally do not fit well. Neither model fits as well as the homogenous QRC model with scores estimated from the data, which therefore remains our preferred model.

The superior fit of the unidiff-by-period specification of change over time for each model of the OD association demonstrates that the Russian mobility regime did change between the pre- and posttransition periods in a unidiff fashion: All the parameters representing the association between origins and destinations changed in the same direction and in the same proportion.[18] Table 6 shows the maximum likelihood estimates of θ, δ_1, and δ_2—the three structural parameters from the homogenous QRC model that capture the association between origins and destinations—for the Soviet and posttransition eras in Russia. We estimated standard errors for these parameters using the jackknife method (Clogg and Shihadeh 1994:36). Each estimate, including the unidiff parameter, is highly significant, adding further evidence that the OD association increased after Russia's market transition.

Each posttransition coefficient is *26 percent larger* than the corresponding parameter for the Soviet era. The positive uniform association parameter indicates that people from higher-status origins have advantages in the labor market. The positive, if modest, diagonal inheritance parameter implies that people from a given origin have a modestly better chance than their contemporaries of attaining a destination occupation in their origin class. The pattern works

[16] Taking into account the identifying restrictions on scores, 14 scores are estimated from the data using the unconstrained version of (1) and 6 are estimated using the constrained version.

[17] Of course, the core model is not nested in either hybrid model, so the likelihood-ratio test cannot be used to compare it to them.

[18] The fact that model C3 does not fit the data using the conventional criterion might raise concerns that the unidiff-by-period specification does not tell the whole story. However, we once again attribute the poor overall fit to the sparseness of our table (see note 12). When we add .5 to each cell, the model L^2 is 417.6 (p = .059).

Table 6. Parameter Estimates, Preferred Model (C3 from Table 5)

	Estimate	SE[a]	*p*	Exp (estimate)
Pretransition				
Diagonal inheritance	.15	.03	<.001	1.16
Farm inheritance	.71	.10	<.001	2.03
Uniform association	.98	.07	<.001	2.65
Posttransition[b]				
Diagonal inheritance	.18			1.20
Farm inheritance	.88			2.41
Uniform association	1.21			3.36
"Unidiff" period effect (f)	1.26	.10	<.001	

Note: Unidiff = uniform difference.

[a] Standard errors are calculated using the jacknife method (Clogg and Shihadeh 1994, p. 36).

[b] These estimates are calculated from the pretransition estimates and the unidiff period effect, so no standard errors are given.

to the advantage of those with privileged origins, because for them class inheritance maintains a privileged class position, while the opposite is the case for those from disadvantaged origins. The specific inheritance of agricultural destinations by persons with agricultural origins further hurts the prospects of those with agricultural origins, decreasing the odds that people raised in that environment will be upwardly mobile out of it. Thus, following Russia's market transition the advantages of Russians with "higher" class origins increased by 26 percent—no small amount.

Explaining the Tightening Up of Mobility: Cohort Replacement or Institutional Change?

Earlier we developed two alternative explanations for a tightening up of the occupational structure with respect to mobility in Russia. First, we hypothesized that institutional changes associated with market transition led to an intragenerational mobility pattern of regression toward origins. Second, we considered the possibility that the tightening up results not from market transition but from a long process of secular change that affects intergenerational mobility via cohort replacement. Because we found a substantial decrease in intergenerational mobility between the late Soviet and post-Soviet periods, we now assess these alternative explanations as best as we can with our data.

First, we directly test the cohort-replacement explanation using an approach similar to that introduced by Breen and Jonnson (2001). We define three cohorts (C): (1) an entering cohort of persons who turned 25 after 1992 (and thus are part of the posttransition, but not the pretransition, sample); (2) a continuing cohort of persons who were between 25 and retirement age for both pre- and posttransition periods; and (3) an exiting cohort of people who reached retirement age after 1992 but before 1998 (and thus are excluded from the posttransition sample). We then test a series of models for the OD association in the CYOD table using our preferred homogenous QRC specification of the OD association. If cohort placement is the main force behind the increased magnitude of the θ, δ_1, and δ_2 parameters from the pre- to the posttransition eras, then permitting the OD association to vary across cohorts should significantly improve the fit of the model and should also render variation across period superfluous. However, if either the COD interaction does not improve the fit of the model or the POD interaction remains significant net of the COD interaction, then we must conclude that the increase in association results (at least in part, if both COD and POD are significant) from a period effect—that is, from the institutional changes associated with market transition.

The data support the view that a period effect, not cohort replacement, accounts for the tightening up of the mobility regime in Russia (Table 7). Models 2 and 4, which specify variation in the OD association by year and by cohort, respectively, fit the data worse than the baseline (no change) model, according to the BIC statistic. Consistent with the pattern evident in Table 5, BIC does not discriminate between the

Table 7. Fit Statistics, Selected Models for the Origin-by-Destination-by-Year Table

Number and Description	L^2	*df*	*p* vs. [1]	BIC	Model *p*	D
Homogenous QRC, Change across Period and Cohort						
1. Baseline (no change)	848.87	767		–6,235	.021	.089
2. Unidiff across Y	833.50	762	.009	–6,205	.036	.088
3. Unidiff acoss P[a]	838.46	766	.001	–6,237	.035	.089
4. Unidiff across C	842.31	765	.038	–6,224	.027	.089
5. Unidiff across YC	832.00	750	.463	–6,095	.020	.088
6. Unidiff across PC	835.99	762	.025	–6,202	.032	.089
Contrast Models 6 vs. 3	2.47	4	.650			

Note: N = 10,264.

[a] Preferred model.

baseline model and the model allowing unidiff change in the OD association over P (model 3), but the likelihood-ratio test clearly favors the latter. Moreover, permitting simultaneous variation of the OD association across P and C fails to improve fit over model 3 by either BIC or by the likelihood-ratio test. Thus, all the evidence in Table 7 favors the interpretation of the tightening up of mobility in Russia as a period effect.

We hypothesized that regression toward origins in intragenerational mobility is the mechanism mediating the impact of the institutional changes associated with market transition on intergenerational mobility. We tested whether intragenerational mobility indeed conformed to this pattern, using the SEIAR data, which provide information on respondents' jobs in December 1990 as well their parents' jobs and their own jobs in early 1998. The regression-toward-origins hypothesis implies that Russians who were upwardly mobile relative to their origins at the end of 1990 will be more exposed to downward mobility between 1990 and 1998, while those who were downwardly mobile in 1990 will be more likely to experience upward mobility during the transition era.

Using our definitions of upward and downward mobility (see Table 3 and the related discussion), a regression-toward-origins pattern is evident among the 2,018 SEIAR respondents who were employed in both December 1990 and early 1998 (Table 8). Of those who experienced long-range upward mobility by 1990, 11.2 percent experienced long-range downward mobility from 1990 to 1998, which dwarfs the percentages for those who did not experience long-range upward mobility prior to 1990. Similarly, those who were downwardly mobile by 1990 experienced substantially more upward mobility from 1990 to 1998. This occurred during a period of diminished opportunities; downward mobility exceeded upward mobility by 30 percent. In mature economies, panel studies usually show more upward than downward mobility as the panels age, so this 30 percent increase in downward mobility quantifies in

Table 8. Intragenerational by Intergenerational Upward and Downward Mobility

	Intragenerational Mobility (1990 to 1998)					
	Long-Range Upward	Short-Range Upward	Immobile	Short-Range Downward	Long-Range Downward	Total
Intergenerational Mobility (Origins to 1990), %						
Long-range upward	0.2	7.9	69.4	11.5	11.2	33.3
Short-range upward	3.6	6.0	75.0	12.8	2.6	28.9
Immobile	7.8	9.3	70.3	9.5	3.2	20.4
Short-range downward	6.5	13.8	65.2	11.3	3.2	12.2
Long-range downward	12.5	14.4	68.3	4.8	0.0	5.2
Total, %	4.1	8.7	70.6	11.1	5.5	

Source: SEIAR respondents with valid father's occupation, 1990 occupation, and current (1998) occupation. N = 2,018.

yet another way the unprecedented collapse of the Russian economy in the first phase of the transition out of state socialism.

CONCLUSION

A decade of economic crisis in Russia reduced the number of desirable jobs and displaced millions of workers. When the displaced workers found jobs again, these jobs mostly were in occupations closer to the workers' social origins than the ones they had before. This was no hardship for the minority from relatively privileged origins; but most Russians found their new position less desirable than the job they had under the old regime. In this sense, market transition in Russia has made social origins more relevant for how occupational opportunities get apportioned.

While our data show how this happened, we must speculate somewhat as to why it happened. The new Russian state abandoned both the rhetoric and the actions that the Soviet state had taken to promote opportunity for people with working-class and peasant origins. Meanwhile, competition for high-status occupations intensified because there were fewer of them and the pay gap between them and other occupations increased. In such circumstances we can expect elites to work all the harder to pass on their advantages to their adult children. Social capital, cultural capital, or human capital—or a combination of all three—may ultimately explain why social origins became such an important arbiter of opportunity during Russia's market transition. To assess the roles of these respective factors, we need more detailed evidence on origins.

Our study also speaks to debates about how stratification worked in the last days of the Soviet era. Communists' claims about equality of opportunity fail to hold up; we confirmed with national data what local studies had found—that is, occupational destinations still depended on social origins at the end of the Soviet era. However, restating this claim in relative terms—that origin-based inequalities were lower under the Soviet system than they might have been otherwise—appears to have merit. The mobility patterns that emerged following the collapse of Soviet institutions and policies indicate that the Soviet system was effective in opening opportunities for people from lower backgrounds, even if it failed to distribute opportunity equally.

To the mobility literature we thus contribute a rare case. Few studies of other countries report an increase in the association between origins and destinations over time. Where change occurs, it tends to be slow and in the direction of *looser* association.[19] This makes our findings for Russia all the more striking. The strengthening of the effects of origins on destinations in less than a decade is not only statistically significant and substantively important; it is also unusual. A 26 percent increase represents a potent channeling of opportunity in post-Soviet Russia toward those who were in the most advantaged origin classes. The available evidence supports our view that this growth in inequality resulted from market transition, not cohort replacement. Increased competition and decreased state intervention in the labor market produced an intragenerational mobility pattern of regression toward origins that benefited people with higher-status backgrounds. Because upward mobility was more common than downward prior to 1990, this pattern implies that the first decade of post-Communist transition has produced more net downward than upward mobility. Those who were downwardly mobile between the end of 1991 and the beginning of 2001 were those who had been most upwardly mobile prior to 1991. Thus, the relationship between origins and destinations is not a fixed, permanent feature of societies, but rests on the foundations of politics and employment regimes that organize and constrain market forces (Fligstein 2001).

Our results show that change in stratification processes can come in unexpected places. Because the market transition debate has focused on intragenerational processes, participants in the debate have missed an important example of change in stratification. We hope specialists on China and Eastern Europe will follow our lead and look for changes over time in intergenerational stratification there. If similar patterns obtain, it might suggest that state socialism did have an effect on equality of opportu-

[19] The only similar case we know of is Bolivia (Kelley and Klein 1977), although Peter Robert notes (in personal communication) that the OD association increased in Hungary after 1989.

nity. At this preliminary stage, however, we refrain from overgeneralizing from our single case study.

Finally, two of our findings address issues of general interest to mobility researchers. First, we have demonstrated the usefulness of distinguishing between professionals and managers. In both pre- and posttransition Russia, professional and managerial origins had different implications for occupational destinations. Ignoring these differences would have led us to underestimate the overall OD association by 6 percent. Second, this case study joins the growing literature that shows the efficacy of so-called vertical models of mobility. We found empirical support in Russia for the idea that classes are differentiated as points on a continuum of opportunity and status and that the chances of a desired occupational destination increase or decrease incrementally as one moves along the scale of origins. Alternative approaches that treat different origins as qualitatively different and not reducible to a quantitative distinction have produced some interesting and useful results. But our study shows that the quantitative distinctions among classes account for most of the systematic variation in class mobility.

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APPENDIX.

Optimizing the Erikson-Goldthorpe Categories for Russia

The Russian surveys we use consistently coded occupational origins and destinations in sufficient detail for us to identify 12 class categories, listed in Table A1.[20] Most of these are standard distinctions in research based on the Erikson-Goldthorpe schema. For our study, however, we deviate in three ways. (1) We drop the distinction between employers and proprietors without employees (classes IVa and IVb) because employers are very rare in Russia (and consequently we observe few in our data). (2) We add a category of people whose origin is "missing" and we treat that as an origin class, because we suspect most of these respondents grew up without a father present. Excluding them could bias our results, as they amount to almost 25 percent of cases and their destinations are concentrated among manual occupations.[21]

[20] The baseline table is 11 × 11 even though we have 12 classes because class IV does not occur as an origin (self-employment was illegal during the Soviet era, and all of our cases have Soviet-era origins) and because we treat "missing" as an origin but not a destination.

[21] A test of the difference between those with missing origins and all other Russians in occupational destinations is the model of independence for a 2 × 9 table in which the rows are origin missing or present and the columns are the 9 destinations we ultimate-

Appendix Table A1. Testing Alternative Specifications of Origin and Destination Categories

Model Specification	L^2	df	Contrast	*p*	% of baseline L^2
1. Full (11 × 11)	931.53	100			100
2. Combine V with VI	905.23	81	2 vs. 1	.122	97
3. [2] + combine I(M), I(P), II(M), and II(P)	829.51	36	3 vs. 2	.003	92
4. [2] + combine I(M) with I(P) and II(M) with II(P)	847.84	49	4 vs. 2	.004	94
5. [2] + Combine I(M) with I(P)	863.83	64	5 vs. 2	.001	95
6. [2] + Combine II(M) with II(P)	886.39	64	6 vs. 2	.338	98
7. [2] + Combine I(M) with II(M)[a]	890.94	64	7 vs. 2	.646	98
8. [2] + Combine I(P) with II(P)	870.41	64	8 vs. 2	.007	96

Note: Data show Russians 25 years old to retirement age in all six surveys.

[a] Preferred model.

(3) As discussed above, we distinguish between professionals and managers within each of the Erikson-Goldthorpe classes I and II.

We test the merit of collapsing upper manual workers—that is, class V with class VI—and professionals with managers—classes I(M) with I(P) and II(M) with II(P)—using the homogeneity criteria introduced by Goodman (1981). Goodman notes that two row categories can be combined if the column variable is independent of the distinction between them—that is, if a chi-squared test fails to reject the null hypothesis for the 2 × J table composed of the two rows in question and all the columns. Similarly two columns can be combined if the row variable is independent of the distinction between them. If pairs of rows and columns refer to the same two classes, then the classes can be combined if the rows and columns in question can be combined. A straightforward and efficient test of whether two classes can be combined is the difference between the likelihood-ratio chi-square obtained for the test of independence in the $I \times J$ table and the likelihood-ratio chi-square obtained for the test of independence in the $(I-1) \times (J-1)$ table that results from combining the two categories in question.[22]

In support of our claim that class matters in Russia, we note that the association between class origins and destinations as defined by this schema is statistically significant, regardless of which version of the Erikson-Goldthorpe schema we use (see Table 2). We first test whether to combine classes V and VI (model 2), because doing so is less controversial. If combining the two classes significantly diminishes the amount of the association in the 11 × 11 table, then they are not isomorphic and should not be combined. If the model L^2 does not fall significantly, then the classes are isomorphic and we can combine them without losing information about social mobility. The results favor combining V and VI: The loss of fit is not significant and we lose only 3 percent of the total association between origins and destinations—an insignificant amount even in a sample of over 10,000 cases. Thus model 2 is the baseline for further testing.

Next we test several ways of aggregating professionals and managers within classes I and II (models 3 – 8). First, we test the combination of all four into a single "salariat" class (model 3), as well as the standard aggregation of managers and professionals within classes I and II (model 4). Both combinations significantly reduce the OD association compared to the (new) 10 × 10 baseline model (2). The coarsest classification sacrifices 8 percent of the association and the conventional aggregation sacrifices 6 percent. But perhaps some pairs of classes are isomorphic with respect to mobility. We see no theoretical reason to combine I(P) with II(M) or I(M) with II(P) so we ignore those pairings and look at the other four possible paired combinations. Upper-level managers and professionals have significantly different mobility patterns (model 5 versus model 2): Ignoring the difference between them would lead us to miss 5 percent of the baseline association. But the corresponding test (6 versus 2) fails to reject the isomorphism of lower-level managers and professionals. Further tests fail to reject homogeneity between upper- and lower-level managers (7 versus 2), but clearly reject the isomorphism of upper and lower professionals (8 versus 2).

In sum, the mobility pattern of upper-level professionals is distinct from those of both upper-level managers and lower-level professionals. The mobility pattern of lower-level managers, however, resembles patterns found among upper-level managers and lower-level professionals. It makes little theoretical sense to combine lower-level managers with both upper professionals and lower managers, and in any case this model produces a significant loss of association (L^2 = 863.42). We opt to combine them with upper managers rather than lower

ly used in this analysis. It shows that the destinations of those who have missing origins are significantly different from the destinations of others: $L^2 = 46.51$; *d.f.* = 8; $p < .05$.

[22] Goodman (1981) also describes a "structural" criterion for combining categories. If parameters referring specifically to categories (for example, row or column scores from an RC-II model) have the same values, then the categories in question can be combined. In circumstances where the structure of the preferred model is not known in advance, the homogeneity criterion is preferable: Because it is based on independence it yields the same results every time (for the same data), regardless of what ends up being the preferred model. Because we do not know the preferred model in advance, we apply the more agnostic and flexible homogeneity criterion.

professionals because model 7 fits slightly better than model 8 (though they are not nested) and because we have outlined above a theoretical justification for this move. Cell counts for the resulting 9 × 9 table, by period and gender, can be found in Table A2. Despite some ambiguity regarding the distinction between lower-level managers and lower-level professionals, over-

Appendix Table A2. Counts for Origin by Destination by Gender and Period: Russia,1988–2000

	Destination									
Origin	I/II(M)	I(P)	II(P)	IIIa	IIIb	IV	V/VI	VIIa	VIIb	Total
A. Men 1988–1992, n										
I/II(M)	25	14	40	8	5	11	29	22	7	161
I(P)	12	17	38	7	4	6	17	10	3	114
II(P)	17	17	37	3	3	7	31	14	2	131
IIIa	25	17	21	11	4	4	15	11	2	110
IIIb	3	0	6	4	0	2	6	7	0	28
V/VI	45	32	76	24	8	24	205	106	20	540
VIIa	30	22	51	15	2	10	133	102	17	382
VIIb	25	19	38	16	4	8	98	80	51	339
Missing	69	51	108	24	13	16	230	161	50	722
Total N	251	189	415	112	43	88	764	513	152	2,527
%	10	7	16	4	2	3	30	20	6	
B. Women 1988–1992, n										
I/II(M)	27	38	72	30	4	4	11	11	5	202
I(P)	9	27	32	14	7	3	6	8	0	106
II(P)	8	34	42	24	7	3	6	7	4	135
IIIa	7	22	52	24	11	2	11	2	1	132
IIIb	1	1	11	14	2	0	2	5	1	37
V/VI	31	97	163	119	39	9	69	57	18	602
VIIa	15	60	105	93	28	6	56	62	15	440
VIIb	21	43	61	71	36	4	46	59	42	383
Missing	51	85	187	145	60	7	124	119	47	825
Total N	170	407	725	534	194	38	331	330	133	2,862
%	6	14	25	19	7	1	12	12	5	
C. Men 1998–2000, n										
I/II(M)	23	8	28	4	10	14	34	21	1	143
I(P)	11	19	17	3	5	12	17	4	0	88
II(P)	15	15	47	12	10	19	28	27	0	173
IIIa	12	6	23	5	3	6	15	14	2	86
IIIb	0	2	3	2	3	1	10	7	0	28
V/VI	40	21	57	20	33	43	214	140	14	582
VIIa	27	14	36	18	18	30	127	122	10	402
VIIb	23	7	29	6	6	12	89	99	41	312
Missing	18	20	50	18	27	44	135	96	17	425
Total N	169	112	290	88	115	181	669	530	85	2,239
%	8	5	13	4	5	8	30	24	4	
B. Women 1998–2000, n										
I/II(M)	16	25	69	42	16	6	9	9	1	193
I(P)	2	21	37	10	8	3	4	5	0	90
II(P)	14	54	82	46	21	10	9	9	3	248
IIIa	7	13	31	21	11	6	9	7	2	107
IIIb	4	0	17	9	9	2	5	8	0	54
V/VI	27	70	152	128	88	16	66	91	14	652
VIIa	15	50	78	90	66	20	43	66	10	438
VIIb	14	17	52	79	45	7	57	55	35	361
Missing	11	42	93	114	71	19	55	69	19	493
Total N	110	292	611	539	335	89	257	319	84	2,636
%	4	11	23	20	13	3	10	12	3	

all our tests confirm our argument that managers and professionals represent distinct classes as far as mobility is concerned (Table A3). Were we to ignore this distinction, as is common in mobility studies, we would miss 6 percent of the association between origins and destinations in Russia.

Appendix Table A3. Scores and Definitions of Types of Mobility

			Destination Class								
			1	2	3	4	5	6	7	8	9
			I(P)	II(P)	I/II(M)	IIIa	IIIb	V/VI	VIIa	VIIb	IV
Number	Origin Class	Scores	.43	.35	.28	.10	–.05	–.23	–.34	–.49	.28
1	I(P)	.43[a]	.00[c]	–.07[d]	–.14[d]	–.33[d]	–.48[e]	–.65[e]	–.76[e]	–.92[e]	–.14[d]
2	II(P)	.35[a]	.07[b]	.00[c]	–.07[d]	–.25[d]	–.41[e]	–.58[e]	–.69[e]	–.85[e]	–.07[d]
3	I/II(M)	.28[a]	.14[b]	.07[b]	.00[c]	–.19[d]	–.34[e]	–.51[e]	–.62[e]	–.78[e]	.00[c]
4	IIIa	.10[a]	.33[b]	.25[b]	.19[b]	.00[c]	–.15[d]	–.32[d]	–.43[e]	–.59[e]	.19[b]
5	IIIb	–.05[a]	.48[a]	.41[a]	.34[a]	.15[b]	.00[c]	–.17[d]	–.28[d]	–.44[e]	.34[a]
6	V/VI	–.23[a]	.65[a]	.58[a]	.51[a]	.32[b]	.17[b]	.00[c]	–.11[d]	–.27[d]	.51[a]
7	VIIa	–.34[a]	.76[a]	.69[a]	.62[a]	.43[a]	.28[b]	.11[b]	.00[c]	–.16[d]	.62[a]
8	VIIb	–.49[a]	.92[a]	.85[a]	.78[a]	.59[a]	.44[a]	.27[b]	.16[b]	.00[c]	.78[a]
9	X	–.34[a]	.76[a]	.69[a]	.62[a]	.43[a]	.28[b]	.11[b]	.00[c]	–.16[d]	.62[a]

Note: Cell entries contain differences between origin and destination scores. Long-range mobility is defined as a difference greater than one-standard deviation (.33).
i = origin class; j = destination class.
[a] Long-range upward mobility.
[b] Short-range upward mobility.
[c] Immobility.
[d] Short-range downward mobility.
[e] Long-range downward mobility.

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