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Social class and workers' rent, 1983–2001[☆]

Stephen L. Morgan[a,*], Zun Tang[b]

[a] Department of Sociology, Cornell University, 358 Uris Hall, Ithaca, NY 14853, United States
[b] Department of Sociology, Cornell University, 323 Uris Hall, Ithaca, NY 14853, United States

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Abstract

Sociologists and economists continue to seek explanations for the growth of earnings inequality since the late 1970s. In this article, we draw upon the structural tradition of labor market analysis in sociology in order evaluate the conjecture that selective rent destruction is a source of the recent increase in earnings inequality. In empirical analysis of the Outgoing Rotation Groups of the Current Population Surveys from 1983 to 2001, we demonstrate that (1) the earnings of workers at the bottom of the class distribution have declined relative to the earnings of those at the top and (2) the variance of wage premia associated with employment in alternative industries has declined relatively more for those at the bottom of the class distribution. Adopting the position from both the sociology and labor economics literatures that these industry wage premia are reasonable measures of industry rents, we conclude that the results support the rent destruction conjecture and, by implication, that structural models of labor markets can explain some of the increase in earnings inequality.

Keywords: Earnings; Class; Education; Occupation; Rent

Earnings inequality is more pronounced now than at any time since World War II. Reviewing the sociological literature on increases in inequality since the 1970s, Morris and Western (1999:623) "challenge the field of sociology to reconstruct its research agenda on stratification and inequality." The sociological literature on earnings inequality in the United States is now expanding, with attempts to build upon several complementary explanations: technological and associated organizational change (e.g., Fernandez, 2001), the emergence of nonstandard employment relations (e.g., Kalleberg, 2003), the consequences of new post-Fordist production regimes (e.g., DiPrete, Goux, & Maurin, 2002), and institutional changes that may have altered labor's share of income (e.g., Wallace, Leicht, & Raffalovich, 1999).[1]

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* Corresponding author. Tel.: +1 6072550706; fax: +1 6072558473.
E-mail addresses: slm45@cornell.edu (S.L. Morgan), zt22@cornell.edu (Z. Tang).

[1] A similar movement toward complementarity of possible explanations can be found within economics. Berman, Bound, and Griliches (1994), Berman, Bound, and Machin (1998) are credited with first fully developing the skill-biased technological change explanation in economics, which achieved a brief consensus. Strong criticism of this explanation emerged in the late 1990s, in part because variation in the rate of growth in inequality is only weakly related to variation in the growth of skill-biased technology (Bernstein & Mishel, 2001). Card and DiNardo (2002) suggested that the entire explanation may be a tautology, as the extant evidence is mostly indirect and drawn from movements in the college-to-high-school wage ratio that cannot be attributed to changes in the supply of college educated workers (see Eq. (2) in Card & DiNardo, 2002). Because of these misgivings, review articles such as Katz and Autor (1999) now place technological change within a general supply-demand-and-institution framework, which allows for the incorporation of exogenous effects produced

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More generally, in cross-national research, work has begun to focus on how institutional variation in the transition to post-industrialism may explain changes in levels of income inequality.[2] Here, welfare-states scholars have begun to develop a tyranny-of-neoliberalism explanation for both the growth of labor market inequality and the downsizing of the welfare state (see Esping-Andersen, 1999; Hicks, 1999; Kenworthy, 2004). Summing up the general argument, Bourdieu (2003[2001]:29) writes of an emergent "neo-liberal utopia" and declares: "Thus has come into being ... a mode of production that entails a mode of domination based on the *institution of insecurity*." The "deregulated financial market fosters ... a casualization of labor that cows workers into submission."[3]

As part of this resurgence of sociological interest in the growth of inequality, we will evaluate the conjecture that recent increases in earnings inequality in the United States are partly the result of the selective destruction of workers' rent. Attributed in sociology to Sørensen (1996, 2000), this conjecture is best regarded as one component of a more general explanation focused on forms of class-biased structural change. In evaluating only this single conjecture, the empirical agenda of our article is somewhat narrow. However, our more general agenda is to help further re-invigorate the structural tradition of labor market analysis in sociology, in response to the call of Morris and Western (1999) and as a complement to the work cited earlier. Accordingly, we will first clarify the rent destruction conjecture which we will directly evaluate, explaining its origins in the new structuralist mode of labor market analysis which achieved maturity just as these trends in inequality were beginning to unfold.

1. A renewal of the structural tradition of labor market analysis in sociology

As reviewed by Kalleberg and Sørensen (1979; see also Berg & Kalleberg, 2001; Farkas & England, 1994), early labor market analysis in sociology was heavily influenced by the segmented labor market perspective developed by economists. Although some critics of the resulting "new structuralism" saw this reliance on economics as an inherent weakness of the sociological approach (see Smith, 1990), it is undeniable that sociological analyses of labor market outcomes led to important empirical work on the variation of wages in advanced industrial society, as conditioned by mobility opportunities within labor markets and other macro-structural factors (see DiPrete, 1993; Eliason, 1995; Hachen, 1992; Hodson, 1984; Kalleberg, Wallace, & Althauser, 1981; Raffalovich, Leicht, & Wallace (1992); Sakamoto & Chen, 1991; Wallace, Griffin, & Rubin, 1989; Wallace & Kalleberg, 1982). The underlying segmented labor markets perspective proved less powerful in general than its originators in economics had predicted, but the new structuralists helped to demonstrate these limitations through careful empirical work.

Morris and Western (1999) noted correctly that, as of the mid-1990s, little work in the new structuralist tradition had been brought to bear upon alternative explanations for the growth in inequality. Such a claim would now be false, as trends in inequality are clearly back on the research agenda of some of these same researchers (see DiPrete et al., 2002; Kalleberg, 2003). And, the prior reliance on segmented labor market models has been replaced by a more subtle and more explicitly sociological focus on the ecology of organizational and market variation. Structures of inequality are now seen as arising from the labor management decisions made by firms, necessarily under the constraints of the environments within which they operate (see Kalleberg, Reynolds, & Marsden, 2003, which builds upon Bridges & Villemez, 1994; Kalleberg, Knoke, Marsden, & Spaeth, 1996). The distribution of earnings across individuals is then seen as a complex function of these management decisions, as they interact with the relative power that firms and workers bring to wage negotiations.

In this article, we build directly on the new structuralist literature that sought to identify advantages obtained from favorable positions within labor markets, as represented imprecisely but still somewhat reliably by the industrial structure. Stinchcombe (1979; see Beck, Horan, & Tolbert, 1978; Hodson, 1984) represents an early example of this type of analysis, wherein an attempt is made to determine the size and type of wage advantages available to workers employed in alternative industries.

Although we draw upon this tradition of analysis in sociology, in our analysis we will not fall prey to a segmentation-mirage. We will avoid an overly reductionist coding of industry locations as either primary/secondary, core/periphery, or monopoly/competitive. Rather, we will frame our analysis of changes inequality by assuming that: (1) structural

by demographic shifts, deindustrialization, and declines in union power.

[2] Economists have also assessed the importance of cross-national differences in institutional arrangements (e.g., Blau & Kahn, 1996).

[3] For Bourdieu, globalization is part and parcel of a neo-liberal utopia, which he claims is a scholastic conceit invented by economists but backed and then implemented by powerful political actors (see Bourdieu, 1998:94–105).

earnings advantages exist even though they may be unobservable because they are defined with reference to counterfactuals; (2) these advantages are autonomous in the sense that they can be considered unattached to particular individuals; and (3) employment in alternative industries reflects some but not all of the relative structural advantage captured by particular individuals. By direct implication of this third assumption, we assume that there are other structural dimensions that reflect earnings advantages as well, and these are not necessarily related to the employment distribution across industries. In order to link this conception of structural advantage to the empirics of changes in inequality, we will draw on the mechanism of rent destruction that we discuss in the next section.

2. Wages, rent, and changes in labor market inequality

In an attempt to reinvigorate structural analysis of inequality in sociology, Sørensen (1996, 2000) embraced the classical economic concept of rent.[4] First, Sørensen (1996) argued that labor market analysts should explain three different quantities: (1) Y^{a}, actual wages paid in the labor market; (2) Y^{c}, wages that would be paid under perfect competition; (3) rent, defined simply as: $r^{c} = Y^{a} - Y^{c}$. Then, Sørensen (2000) argued that social class analysts should explain patterns of inequality by accounting for rights to rent-generating assets, conceptualized broadly (and some would argue too abstractly; see Wright, 2000) as structurally advantageous positions.

As part of this broad agenda, Sørensen asserted, without citing much evidence, that rent destruction is a plausible explanation for some of the recent increase in inequality in the United States, writing:

> There is, however, substantial recent evidence that shows that capital has become very effective at eliminating the advantages of the working class in terms of rents obtained in the labor market. Eliminating these advantages has contributed to the increase in inequality. (Sørensen, 2000:1550)

Sørensen then noted the evidence from labor economics of increases in within-group inequality (specifically, for groups defined by both education and occupation) as well as greater returns on unmeasured skills. From these developments in the labor market, he concluded that "structural locations seemed less relevant for explaining the variation in earnings" (Sørensen, 2000:1552).

Before developing specific implications of this rent destruction conjecture, and then reporting on the empirical analysis that they motivate, in the next section we discuss the concept of rent with reference to both consumer product markets and labor markets. We have two specific goals: (1) to clear up some of the confusion over the reception of Sørensen (1996, 2000) and (2) to more directly highlight the connections between the rent destruction conjecture and the structural tradition of labor market analysis.

2.1. Rents in product and labor markets

Without too much loss of precision, rent can be thought of as pure profit, usually of the sort that accrues to those who exercise monopoly power. As Sørensen (1996) recounts, rent-focused analyses of market outcomes rely on a slightly contrarian perspective on how real-world markets function (see Rowley, Tollison, & Tullock, 1988). The central claim of the approach is that rents are captured by those who control productive assets for which excess demand exceeds their fixed supply. In classical economics, rent on land accrues to those who own it because nature constrains the supply of land. More recently, scholars have developed the case that rents can be created in any market if appropriate instrumental action successfully constrains the supply of a productive asset. In the rent-seeking literature, corporate actors manipulate political economy processes to secure regulations and licenses to protect themselves from competitors (see Tullock, 1989). The resulting limitation on supply drives up the price of goods in the market, thereby generating rents for corporate actors who have enacted effective rent-seeking campaigns.

For the labor market, the concept of rent is somewhat more difficult to deploy, as first one must identify the market within which the rents originate and then the distribution of rent across those who have a claim to it. To appreciate the complexities, consider the following two types of rent that workers can capture.

[4] The concept of rent has been utilized by Marxist scholars for more than 80 years (e.g., Tawney, 1920, Chapter 5). Moreover, Marx wrote on rent in Volume III of *Capital*. Sørensen, however, saw his work as a reinterpretation of early work in the sociology of labor markets (e.g., Sørensen and Kalleberg, 1981) using the literature in organizational behavior, labor economics, and personnel economics. And, since rent as an economic concept predates Marx in the work of Ricardo, it seems fair to follow Sørensen's lead and not discuss the neo-Marxist literature on rent that unfolded over the twentieth century. If, however, one were to seek to determine the points of connection, one could start with: (1) the loyalty rent perspective developed to formalize managerial advantages (see Wright, 1997:21 and the prior work that he cites) and (2) the contested exchange model of Bowles and Gintis (1993), grounded in earlier work on Marxian definitions of economic rent (Bowles, 1985; Schor and Bowles, 1987).

First, there are rents that emerge in markets for the final goods produced by firms, such as those for a monopolist who is protected from competitors by a license or regulation (or a quasi-monopolist who wields pricing power because capital requirements represent an effective barrier against the entry of competitors). In a typical publicly held firm, any such rents must be divided among managers, production workers, and shareholders (see Fligstein, 2001; see also Dore, 2000). In this case, workers may receive a share of a rent, but that rent is created by a restriction on the supply of the good the firm is producing. Labor market processes are not irrelevant to the size of this type of rent, as, for example, the bargaining power of workers is then an inverse function of the availability of alternative laborers willing to work for a lower wage. Nonetheless, the fundamental source of this rent is not the labor market itself.

Second, there are rents that emerge directly in the labor market. Drawing on the work of Marshall, Sørensen argued that some individuals obtain rents because they possess rare abilities for which there is unmet demand. Here, individuals receive rents on their abilities because they have received free gifts of nature that others value highly.

In this article, we will focus on the first type of rent, as such rents are well defined with reference to meaningful counterfactual conditions. It is, at least to us, unclear what a perfectly competitive labor market actually is, in comparison to what we typically observe. And hence, it seems impossible to definitively determine whether, for example, individuals with unique abilities are earning rent or instead simply a competitively determined price for offering their abilities to an employer.

2.2. Workers' negotiated rent

Although we will discuss rent more broadly in the concluding section of this article, in the interim we will confine our empirical analysis to a worker's share of firm rent. We will label this rent as *workers' negotiated rent*, which we define as the absolute difference between the wage a worker is paid and the counterfactual best wage that he or she could earn performing qualitatively similar duties at any employer. In the counterfactual condition, all potential employers are assumed to be operating in perfectly competitive markets for their final goods.[5]

Why refer to such rents as negotiated? When firms earn rent producing goods for exchange in consumer markets, owners and their employees must bargain over the division of rents. In unionized firms, the bargaining process is explicit and institutionalized. In non-unionized firms, the bargaining may be latent, with loyalty wages and employee ownership programs reducing the perceived benefits of unionization. Whatever the theoretical characterization of the mechanism, and whether or not mediating institutions such as labor unions are central, the crucial idea is that firms that command rents in the markets for their final goods pay their workers more, on average, than firms without rent to share. Since it is reasonable to assume that rent-sharing firms do not pay higher wages entirely based on altruism, we consider these wage premia to be negotiated and label them as such.

Although this focus on negotiated rent is a modest innovation for the rent-based framework outlined by Sørensen (1996, 2000), it is consistent with a dominant theme of much of the new structuralist labor market literature. From Kalleberg et al. (1981) through Wallace et al. (1999), the new structuralist tradition has analyzed how wage rates vary with the power of workers relative to employers. In focusing interest on workers' negotiated share of firm's rent from product market revenues, we bring the rent-based framework closer to this structural tradition of analysis.

3. Testable implications of a decline in workers' negotiated rent

Two presuppositions make possible an empirical evaluation of the rent destruction conjecture. First, individuals who work for rent-generating firms earn wages that are higher, on average, than the wages they would command, on average, if all skill-equivalent workers were randomly redistributed across all skill-demand-equivalent positions in the labor market. Second, individuals who work for rent-generating firms earn wages that are at least as high as those they would receive from the same firm if all firms that

[5] Two related definitions can be found in the literature. According to one definition, a worker's rent is the absolute difference between the wage he or she is paid and the *next best wage* that he or she could earn performing qualitatively similar duties at *an alternative employer*. According to another definition, a worker's rent is the absolute difference between the wage he or she is paid and the *least acceptable wage* for which he or she would work performing qualitatively similar duties for *any employer*. These definitions differ in the specification of the counterfactual competitive wage, and they are only slightly more specific than Sørensen's Y^c. The first definition is the neoclassical definition, inspired by Pareto, and which is infrequently but still sometimes invoked in labor economics (see Lazear, 1995; Lazear and Freeman, 1996). The second definition is the classical definition, inspired by Ricardo's analysis of land rent.

could produce qualitatively similar goods were doing so for sale in counterfactual perfectly competitive markets.

To make the connections with Sørensen's simple definition that workers' rent is $r^c = Y^a - Y^c$, consider the following more specific notation. The rent framework we adopt asserts that average actual wages among those working for firms with positive rents to share are greater than analog average counterfactual wages under hypothetical perfectly competitive conditions:

$$E[Y^a|\mathbf{A}, R = 1] > E[Y^c|\mathbf{A}, R = 1] \tag{1}$$

where $\mathbf{A}$ is a vector of observable attributes of individuals (education, social class, labor force experience, and basic demographics characteristics), R is an indicator variable equal to 1 for all individuals who are employed by firms that command rents in the markets for their final goods, and the counterfactual perfectly competitive conditions which define the counterfactual wage Y^c refer to the conditions in the markets for the firms' final goods. With this definition of the fundamental assumption of our rent-based framework, the average amount of rent received as wages by workers with attributes $\mathbf{A}$ equal to $\mathbf{a}$ is then

$$\begin{aligned} E[r^c|\mathbf{A} = \mathbf{a}] &= \pi_a E[(Y^a - Y^c)|\mathbf{A} = \mathbf{a}, R = 1] \\ &= \pi_a \{E[Y^a|\mathbf{A} = \mathbf{a}, R = 1] - E[Y^c|\mathbf{A} = \mathbf{a}, R = 1]\} \end{aligned} \tag{2}$$

where π_a is the proportion of individuals with attributes $\mathbf{a}$ who are employed by firms that command positive rents.

With this notation, the challenges to the empirical analysis of rents received by workers are laid bear, even after limiting the scope of analysis by stipulating that the counterfactual conditions refer only to the market for firms' final goods. To determine the average amount of rent received by different types of workers (i.e., those in one class versus another), one must know (1) what proportion of these workers are employed by firms with rent to share and (2) what these workers' average wages would be if their employers no longer had rents to share. As a result, analysis is only possible under strict assumptions, and (as with all empirical research) all interpretations are conditional on acceptance of these assumptions.

We will not attempt to recover all of the ingredients of Eq. (2) from our data, and thus we will not provide estimates of average rents received by workers of different types. Instead, we will only model relative advantages for workers within social classes, conditioning on observable attributes, in order to assess whether or not within-class relative advantages and disadvantages have disappeared among those whose average wages have declined the most. Before outlining why these are the crucial results that are needed to evaluate the rent destruction conjecture, we first summarize in the next section our strategy for recovering relative advantages using the industrial structure of employment.

3.1. Identifying rents with industry of employment

One strategy for modeling structural advantages of this type is to estimate the effects of industry of employment on individuals' wages. Stinchcombe (1979), for example, wrote in an early contribution to the new structuralist literature: "To provide a general classification of labor market structures... we need to take into account the degree to which there is a monopoly in the product market which allows wages to remain above the competitive level" (Stinchcombe, 1979:218).[6] In our article, we follow this strategy as well, although we draw model specifications directly from the industry wage differentials literature of labor economics (Katz & Summers, 1989; Krueger & Summers, 1988). With origins in the old institutional economics (see Slichter, 1950), for this literature wage premia paid to observationally equivalent workers in alternative industries are viewed as rent payments. In comparison to much of the new structuralism literature, the tradition in labor economics is to leave the industrial structure "as is," resisting the temptation to lump industries together to form primary and secondary labor market sectors. This restraint is preferable because there is no clear correspondence between the product markets within which alternative industries operate and the labor market that they share, no matter how "balka-

[6] Stinchcombe then analyzed labor market outcomes across a seven-category industrial classification, deriving a subtle set of conclusions validating his definition of labor market structures. It is remarkable how much Stinchcombe's article presages Sørensen's framework, and indeed our own empirical analysis. Stinchcombe (1979:241–42) reaches three basic conclusions: (1) "the presence of monopoly or 'protection' in the commodity and labor markets affects the degree to which a firm or an industry can develop a status system in which the wages of some people are substantially above the wage for which the reserve army of the unemployed would be willing to work;" (2) "there are differences in the wage structure of industries and the kinds of career development of incomes;" and (3) some "workers do not get nearly as large wage premiums for their experience, and wages above the Ricardian competitive level occur among experienced workers in those kinds of labor markets protected by the kinds of barriers we outlined in our classification of industries." For other references from the sociological literature consistent with this Stinchcombe-Sørensen view, see Farkas, England, & Barton (1994:94) and Tilly (1998:238).

nized" (see Kerr, 1977) the labor market may appear to be.

A crucial assumption that enables empirical analysis in this tradition of labor market analysis is that the average actual wages of those employed by firms without positive rents to share are no larger than the counterfactual average wages that would be paid to those currently employed by firms with positive rents if those firms no longer had rents to share:

$$E[Y^{\mathrm{a}}|\mathbf{A}, R = 1] > E[Y^{\mathrm{c}}|\mathbf{A}, R = 1] \geq E[Y^{\mathrm{a}}|\mathbf{A}, R = 0]. \tag{1a}$$

Under this implicit assumption, an equation is estimated:

$$Y^{\mathrm{a}}_{ik} = \mathbf{A}_i\boldsymbol{\beta} + \mu_k + e_i, \tag{3}$$

where Y^{a}_{ik} represents each individual i's log wage in industry k, where $\boldsymbol{\beta}$ represents fixed effects of observable attributes $\mathbf{A}_i$ of individuals, and where μ_k are fixed industry effects for each industry k.

In this tradition of analysis, the net industry coefficients μ_k are considered measures of relative industry rents captured by workers (after the wage differences produced by alternative attributes have been partialled out across the sample under consideration). Accordingly, if there are no wage advantages associated with employment in one industry rather than another, net industry effects would be 0. In this case, the variance of all industry effects would be zero. But, when relative wage advantages exist, the variance of these net industry effects is positive, for some workers will be receiving relative advantages in comparison to others.[7]

In our analysis, we will estimate relative rents by conditioning on attributes in two different ways. For our first set of models, we will condition with statistical methods, as in the economics literature, using covariance adjustments for characteristics such as education, race, and gender. Thereafter, we will further condition deterministically on individuals' social class, estimating all models by class with covariance adjustments separately within class. It is this within-class analysis of relative industry effects that is crucial for our test of the rent destruction conjecture, as we specify next.

3.2. Testable implications

If increases in inequality over the past two decades reflect selective rent destruction for the working class, and if industry wage premia reflect rent payments, then two testable implications follows:

Implication 1. The relative earnings of those at the bottom of the class distribution have declined.

Implication 2. The variance of net industry effects has declined relatively more for those at the bottom of the class distribution.

Both of these implications must be supported in empirical analysis if either set of results is to be seen as supportive of the rent destruction conjecture.

Were we to observe, for example, relative declines in the variance of net industry effects for the working class without a contemporaneous relative decline in average wages for the working class, this would not be evidence that rent destruction has contributed to the increase in earnings inequality. We would merely have evidence of a convergence of wages toward average levels of earnings that have kept pace with those of other classes. The goal of analysis is therefore to evaluate whether the data are jointly consistent with these two implications in an attempt to determine whether the wages of the working class have converged around lower average levels unrelated to traditional sources of structural advantage identified by the industrial structure in the extant literature.

In order to evaluate these implications, we must adopt a definition of social class. Sørensen did not offer a class schema as a standard against which one could examine change, even though he was attempting to use a form of class analysis to account for important structural change. In this article, we will adopt two different schemas: (1) the class schema introduced by Erikson, Goldthorpe, & Portocarero (1979), which we label the EGP class schema hereafter, and (2) the major occupation categorization of the U.S. census. We will not offer any defense of the major census occupation categories, as we will invoke them only briefly in the course of analysis. But, we will offer a justification for our reliance on the EGP schema in the course of presenting our data and results. By looking at educational groups as well, we will also show that our findings are not likely a function of our reliance on these two class schemas.

[7] Sørensen (1996:1354–5) identifies the wage premia associated with industry location as a type of composite rent (that is, one which emerges because the industry and the worker are uniquely matched to each other). We believe that this is incorrect, or at least too narrowly conceptualized, both in terms of Sørensen's own framework and with reference to the inter-industry wage differentials literature. One need not assume a unique productivity boost for workers employed by firms that generate rents. All that is required is that firms command rents in the consumer market for their goods, that workers recognize the existence of these rents, and that workers either demand a share of the rents or that firms decide to share them because of norms of fairness.

4. Data

We analyze data drawn from the 1983 through 2001 merged Outgoing Rotation Groups (hereafter, ORG) of the Current Populations Surveys (hereafter, CPS). Each household entering the CPS is interviewed for 4 months, then ignored for 8 months, and then interviewed again for 4 more months. At the end of each of the 4-month interview periods, households rotate out of the sample, returning after the first rotation but then leaving the sample permanently at the end of 16 months. Individuals in the fourth month of each rotation group are designated as the outgoing rotation group and asked additional questions during the interview, such as their usual weekly earnings and whether or not they are union members.

Table 1 presents the means (and, where relevant, the standard deviations) of the variables that we will use in analysis. Our main analysis sample includes 2,508,500 individuals between the ages of 18 and 64, who reported working 35 hours or more per week, earning $50 or more per week, and who we identified as members of one of the 7 EGP social classes listed and described in Table 1.[8] We adjusted the earnings variable for inflation, using the Bureau of Labor Statistics' Personal Consumption Expenditures Deflator to convert nominal dollars to 2000 dollars. Details of our coding procedures, especially the handling of the topcoding of the earnings variables and our construction of the EGP class schema from 1980 and 1990 three-digit census occupation codes is available in an extensive Supplementary Appendix S. This appendix is available from the authors by request and is posted on the website of the first author.

Why utilize the EGP class schema? Although many class mappings exist in the sociological literature, in the past two decades the EGP schema has become the most prominent, primarily because it has been effectively deployed in a wide variety of substantive contexts, such as in cross-national studies of social mobility (e.g., Erikson & Goldthorpe, 1992; Goldthorpe, 1987; Hout, 1989) and voting (e.g., Heath, Jowell, & Curtice, 1985; Manza & Brooks, 1999). The EGP schema has also received a recent theoretical justification

[8] In other words, our analysis sample excludes non-full-time workers and all members of EGP classes IVa, IVb, IVc, and VIIb (see notes to Table 1 for specific descriptions). We focus on full-time workers because they are the core of the labor market. And, we exclude classes IV and VIIb from our analysis because it is customary in the literature in both sociology and economics on the distribution of self-reported earnings to treat the earnings of farmers, self-employed artisans, and small-scale proprietors as too severely complicated by taxation and accounting issues.

Table 1
Means and S.D. of variables

Variable	Mean	S.D.
Natural logarithm of weekly earnings	6.306	.583
EGP social class		
I: Higher-grade professionals, administrators, and officials; managers in industrial establishments	.174	
II: Lower-grade professionals, administrators, and officials; higher-grade technicians; supervisors of non-manual employees	.202	
IIIa: Routine non-manual employees, higher-grade (administration and commerce)	.165	
IIIb: Routine non-manual employees, lower-grade (sales and service)	.048	
V: Supervisors of manual workers; lower grade technicians	.052	
VI: Skilled manual workers	.114	
VIIa: Semi- and unskilled manual workers (not in agriculture)	.246	
Major census occupational group		
Managerial and professional	.287	
Technical, sales, and administrative support	.300	
Service	.106	
Precision production, craft, and repair	.131	
Operators, fabricators, and laborers	.170	
Agricultural laborers	.006	
Level of educational attainment		
More than a bachelor's degree	.094	
Bachelor's degree	.171	
Some college	.250	
High school degree	.376	
Less than a high school degree	.110	
Industry		
Mining	.008	
Construction	.061	
Lumber and wood products, except furniture	.007	
Furniture and fixtures	.007	
Stone, clay, glass, and concrete product	.006	
Primary metals	.009	
Fabricated metal	.014	
Forestry and fisheries	.001	
Machinery, except electrical	.027	
Electrical machinery, equipment, and supplies	.023	
Motor vehicles and equipment	.014	
Aircrafts and parts	.006	
Other transportation equipment	.008	
Professional and photographic equipment	.008	
Toys, amusements, and sporting goods	.001	
Miscellaneous and not specified manufacturing industries	.004	

Table 1 (*Continued*)

Variable	Mean	S.D.
Food and kindred products	.018	
Tobacco manufactures	.001	
Textile mill products	.007	
Apparel and other finished textile products	.011	
Paper and allied products	.008	
Printing, publishing, and allied industries	.017	
Chemicals and allied products	.014	
Petroleum and coal products	.002	
Rubber and miscellaneous plastics products	.009	
Leather and leather products	.002	
Transportation	.049	
Communication	.018	
Utilities and sanitary services	.017	
Wholesale trade	.044	
Retail trade	.129	
Banking and other finance	.035	
Insurance and real estate	.037	
Private household services	.004	
Business services	.044	
Repair services	.011	
Personal services, except private household	.020	
Entertainment and recreation services	.011	
Hospitals	.046	
Health services, except hospitals	.039	
Educational services	.081	
Social services	.018	
Other professional services	.038	
Agriculture service	.005	
Other agriculture	.001	
Public administration	.061	
Covariates for the earnings equations		
Female	.429	
Race		
Black (includes black Hispanics)	.117	
Hispanic (includes white Hispanics)	.085	
Other (primarily Asians)	.035	
Years of education	13.242	2.631
Years of experience	18.577	11.541
Union member	.181	
Ever married	.761	
Region		
Central	.241	
South	.351	
West	.206	

Notes: The number of respondents for the table is 2,508,500 and includes all individuals between the ages of 18 and 64 in the 1983 through 2001 ORGs who reported usual work hours of 35 or more per week, average weekly earnings of $50 or more, and were employed in one of the 7 EGP classes listed in the table. Data are weighted by the earnings weight (earnwt) provided by BLS. The omitted reference categories for the earnings equations are public administration for industry, white non-Hispanic for race, and east for region.

(see Goldthorpe, 2000) grounded on a broad set of literature from both economics and sociology, and it has been introduced into the economics literature (see Erikson & Goldthorpe, 2002).

Ultimately, the utility of the EGP schema is a function of its explanatory power. As we will show in the empirical analysis of the next section, the EGP schema fairs well in this regard. However, we are unable to offer any systematic evidence that the EGP schema outperforms its competitors in class analysis. With the ORG data, we cannot implement any of the intriguing Marxist schemas that have been proposed, as they require information beyond the census occupation codes. Moreover, because we need to estimate industry effects within each year and within each class, we do not have a large enough sample to enable an analysis with the micro-classes proposed by Grusky and Weeden (2001, 2002).

5. Results

5.1. Three complementary depictions of the growth in inequality

Fig. 1 presents three alternative depictions of the growth in earnings inequality across subgroups of the core full-time labor force from 1983 to 2001. For panel (a), the mean earnings of the seven main EGP classes are presented. The largest increase in earnings is present for class I, with an increase of 23% (from 973 to 1201 in inflation-adjusted year $2000 per week). In contrast, the mean earnings of classes V and VIIa increased by only 7% (from $722 to 769 and $438 to 469 per week, respectively). There is a clear and interpretable pattern in the trends taken together. On average, the mean earnings of classes I, II, IIIa, and IIIb increased relative to the mean earnings of classes V, VI, and VIIa (i.e., 23, 21, 18, and 20% relative to 7, 10, and 7%).

Panel (b) presents the archetypical representation of the recent growth in labor market inequality. For this depiction, the same individuals from panel (a) were remapped into five educational attainment groups. The mean earnings of those who obtained advanced post-secondary degrees increased by 37% (from $940 to 1287 per week). In contrast, the mean earnings of those without a high school diploma declined by 10% (from $454 to 413 per week). All other educational attainment groups experienced relative earnings increases in between these two extremes. And, in general, the amount of increase was greater for more educated groups.

Finally, panel (c) presents one further representation of the growth of inequality, where the same individuals from panels (a) and (b) were remapped into the six coarse

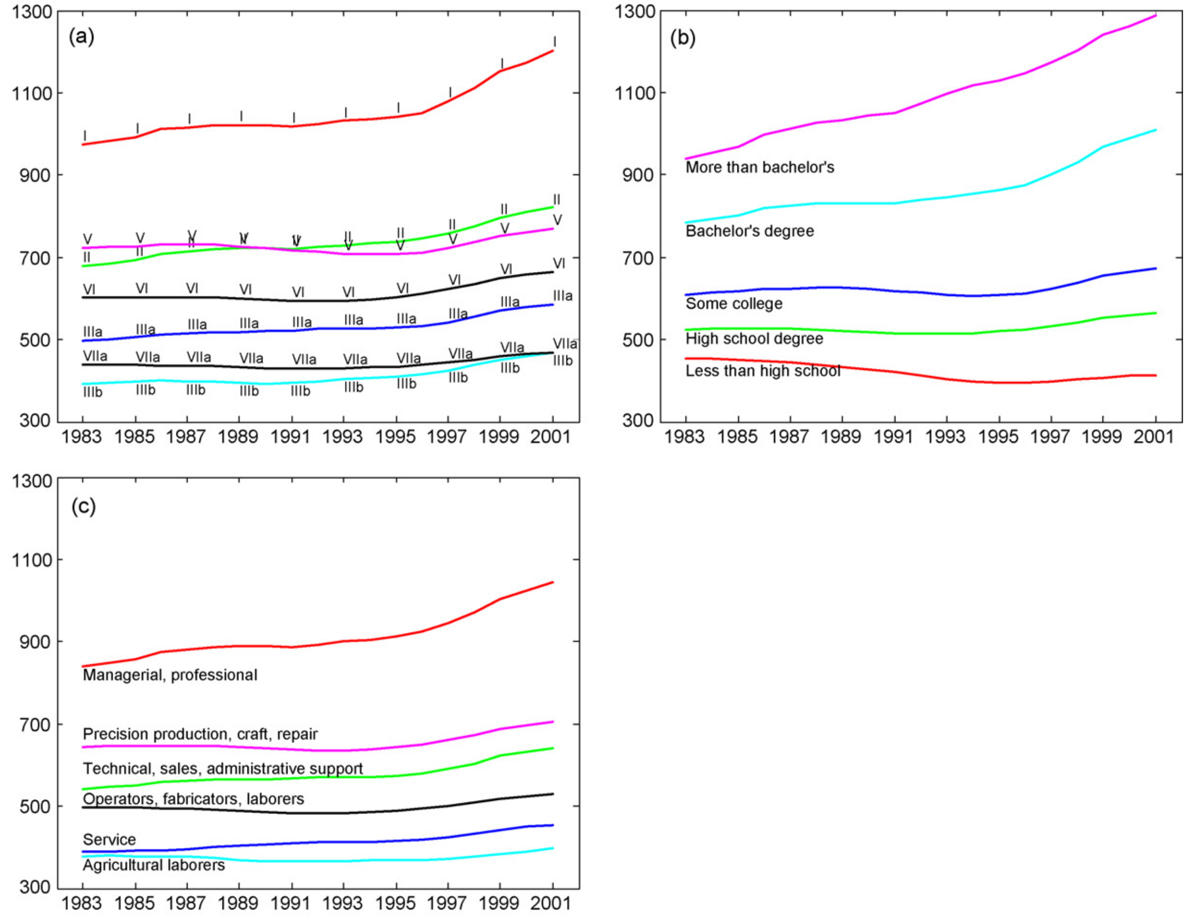

Fig. 1. Mean weekly earnings by EGP class (panel a), by educational category (panel b), and by major census occupational category (panel c).

major occupational categories adopted by the U.S. Census Bureau. The earnings of managers and professionals increased by 25% (from $838 to 1044 per week) in contrast to the modest 6% gain in the earnings of operators, fabricators, and laborers ($497–528 per week).

Table 2 demonstrates why the three panels of Fig. 1 are similar and yet not identical. Within columns, each of the seven main EGP classes is subdivided into the major occupational categories of the census and then the educational attainment categories from Fig. 1. The cross-classifications are offered for two separate time periods that correspond to the last 5 years of the two most recent economic expansions in the United States, 1985–1989 and 1996–2000.

Because the EGP class schema was constructed by appropriately categorizing three-digit census occupational codes, there is a strong relationship between the major census occupational categorization and the EGP class schema, as shown in the first panel of Table 2. And yet, there are differences. There is no simple correspondence between EGP classes V, VI, and VIIa and the occupations that would be categorized as service occupations, precision, production, craft, and repair occupations, or operating, fabricating, and laboring occupations. Even class I is drawn from more than just occupations that the census bureau categorizes as managerial, professional, and specialty occupations.[9]

The bottom panel of Table 2 breaks down the same EGP social classes by level of educational attainment.

[9] For example, class I includes a few occupations categorized as technicians (airline pilots and navigators) and others categorized as operators (ship captains and marine engineers) by the census bureau. Also, a few "farming, forestry, and fishing" occupations are included in class VIIa instead of class VIIb, and as a result are included in our sample. These occupations include "groundskeepers and gardeners, not farm," "graders and sorters, agricultural products," and "fishers." As detailed in the supplementary appendix available from the author's by request, we made these coding decisions by following closely the extant literature on the EGP class schema.

Table 2
Percentage distributions of EGP classes across major census occupational groups and levels of educational attainment in 1985 through 1989 and 1996 through 2000

	EGP classes from 1985 to 1989							EGP classes from 1996 to 2000						
	I	II	IIIa	IIIb	V	VI	VIIa	I	II	IIIa	IIIb	V	VI	VIIa
Major census occupational group														
Managerial and professional	99.3	52.5	.3	1.8	2.1	0	0	99.5	57.5	.5	1.9	1.9	0	0
Technical, sales, and administrative support	.5	42.9	99.7	85.9	16.5	0	2.3	.4	38.0	99.5	82.9	16.5	0	1.9
Service	0	4.7	0	12.3	17.5	3.0	29.5	0	4.5	0	15.2	21.0	3.0	32.5
Precision, production, craft, and repair	0	0	0	0	61.1	83.1	2.0	0	0	0	0	57.5	84.8	2.0
Operators, fabricators, and laborers	.2	0	0	0	2.8	14.0	64.0	.1	0	0	0	3.1	12.2	60.8
Agricultural labor	0	0	0	0	.1	0	2.2	0	0	0	0	.1	0	2.8
Total	100	100	100	100	100	100	100	100	100	100	100	100	100	100
Educational attainment level														
More than a bachelor's degree	29.3	20.5	3.4	2.4	3.8	.7	.7	27.2	15.0	2.4	1.6	2.0	.5	.5
Bachelor's degree	30.7	28.9	12.2	9.1	10.1	2.9	3.0	38.0	34.1	16.5	11.2	12.1	3.6	4.1
Some college	19.5	24.2	29.1	24.4	26.0	16.9	13.7	21.9	30.3	40.0	34.0	39.0	28.2	21.7
High school degree	18.4	23.6	51.0	52.4	48.3	56.4	51.7	12.1	18.7	38.5	44.5	40.1	51.4	50.9
Less than a high school degree	2.2	2.8	4.2	11.7	11.8	23.2	30.8	0.8	1.8	2.6	8.9	6.8	16.3	22.8
Total	100	100	100	100	100	100	100	100	100	100	100	100	100	100

Not surprisingly, individuals in class I are the most likely to obtain college and postgraduate educational degrees while those in classes VI and VIIa are the least likely. And yet, there remains a good deal of heterogeneity in educational attainment within social classes. Approximately 21% of those in class I have no more than a high school degree in both time periods. The proportion of those in class VIIa who have obtained some education beyond high school is approximately 17% in the first time period, and increases to nearly 26% for the second time period.

Return now to Fig. 1. All three panels show similar trends, but they do so in slightly distinct ways because the underlying categorization of individuals differs. We see no need to assert any arguments for the *a priori* superiority of any of the three depictions in Fig. 1. The only justification for relying solely on one of them would be to assert that increases in inequality are attributable to a single causal narrative that depends only on returns to the skills measured by educational degrees or only on the employment relations captured by the EGP class schema or the census occupational categorization. One might argue for the educational breakdown of the trends if one were willing to accept the skill-biased technological change explanation as a master causal narrative (and further that coarsely measured educational credentials reveal the overwhelming majority of skills relevant to these changes), but even the recent economics literature does not adopt this position (see Card & DiNardo, 2002).

5.2. A general decline in the dispersion of industry wage premia

Before analyzing class differences in industry rents, we first analyze trends in industry rents across the entire labor market. We use least squares to estimate the fixed industry effects μ_k in Eq. (3) every year from 1983 to 2001, and we then assess whether or not there has been a decline in the standard deviation of industry effects (see Table 1 for descriptions of the 46 industries).

For this set of models, we adopted the customary specification of the earnings equations in past studies of industry rents in labor economics (see Katz & Summers, 1989; Krueger & Summers, 1988). For the first specification of Eq. (3), we included among the attribute covariates $\mathbf{A}_i$ of individuals the variables listed in the last panel of Table 1: years of education (and its square), years of potential labor force experience (and its square and cube), a dummy variable for gender, three dummy variables for race/ethnicity, three dummy variables for region, and a dummy variable for marital status (along with interactions between the gender dummy and the

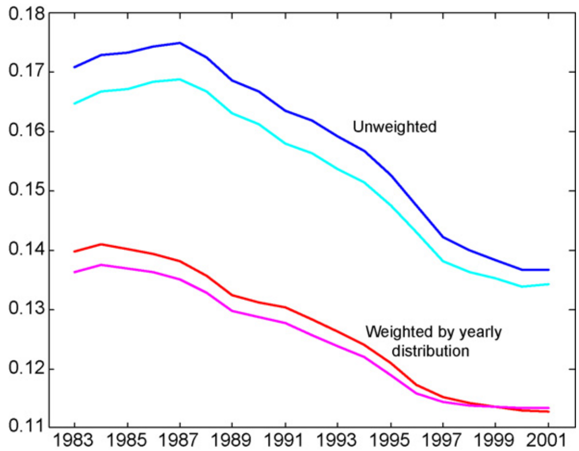

Fig. 2. Five-year moving averages of the standard deviation of industry effects. *Note*: The lower lines of each pair are further adjusted by union status.

variables for education, labor force experience, and marriage). For the second specification, we added a dummy variable for union membership. The first specification yields estimated industry effects $\hat{\mu}_k$ for each year that are covariance-adjusted for location and demographic characteristics, along with years of educational attainment. The second specification yields estimated industry effects that are further adjusted for union status.

Fig. 2 presents trends in the dispersion of estimated industry effects between 1983 and 2001. The dispersion of industry effects is captured by two different summary statistics, an unweighted standard deviation and a corresponding employment-share-weighted standard deviation. For the top two lines, we calculated the standard deviation of the industry effects, using

$$\sqrt{\sum\nolimits_k \frac{1}{K}(\hat{\mu}_{kt} - \hat{\bar{\mu}}_t)^2} \quad (4)$$

where K is the total number of industries under consideration (i.e., 46 for the models in Fig. 2). For the standard deviation computed in each year t by Eq. (4), the squared deviation of each industry coefficient μ_k is given equal weight. For the bottom two lines in Fig. 2, we summarized trends in the dispersion of the same estimated industry effects, using

$$\sqrt{\sum\nolimits_k \frac{n_{kt}}{N_t}(\hat{\mu}_{kt} - \hat{\bar{\mu}}_t)^2} \quad (5)$$

where n_{kt} is the number of respondents in industry k in year t and N_t is the number of respondents in the sample in year t. This weighted standard deviation is the common summary statistic for the size of industry differentials in the industry rents literature (e.g., Katz & Summers, 1989, Table 2). It also has two inherent advantages over the unweighted standard deviation. First, it enables a desirable individual-level interpretation—the expected magnitude of the industry wage differential for a randomly selected individual from the population (i.e., not the expected industry differential across industries of varying size, as for the standard deviation from Eq. (4)). Second, the employment-share-weighted standard deviation minimizes the noise contributed by random sampling error across years, since such errors for each fixed industry effect are inversely proportional to the number of workers in each industry.

Within each pair of trend lines, the dispersion of industry effects from the two different specifications of Eq. (3) is summarized. The upper line of each pair is the 5-year moving average of the standard deviation of the industry effects without an adjustment for union status whereas the lower line of each pair is the 5-year moving average of the standard deviation of industry effects with an adjustment for union status. All four lines show the same basic pattern, which is a general decline in the dispersion of industry effects. The estimated decline is less severe when the underlying industry effects are adjusted for the union status of individuals. We prefer the weighted standard deviation for the reasons just mentioned and therefore will use it as our summary measure in the rest of this article.

Across the entire labor market, and after adjusting for the baseline demographic and human capital covariates only, the estimated industry differential in earnings that an individual would obtain by being in one industry rather than another declined by 19% between 1983 and 2001 (i.e., from .140 to .113 in the natural logarithm of weekly earnings). Further adjusting for union status in each year, the estimated standard deviations are smaller and decline less (only about 17% from .136 to .113 in the natural logarithm of weekly earnings). Industries that pay their workers more tend to be industries that are more heavily unionized, and this was more so in the early 1980s than at the turn of the century. Thus, as the earnings advantages associated with being a union member have declined so have the advantages of being employed in one industry rather than another. But, as best one can determine with CPS data, neither the existence of industry wage premia nor trends in the size of industry wage premia can be easily explained away by a simple unionization explanation.

Although we do not report the full results from the underlying regression models (because Fig. 2 summarizes 1748 industry coefficients – 46 parameters for each industry, for 2 model specifications, in 19 separate yearly regressions – under two different weighting schemes), two sets of unreported results deserve mention. First, the 19 yearly sets of regression models fit the data well, and the associated R^2 values reveal an important trend. For models without an adjustment for union status, R^2 declined steadily from .457 to .401 from 1983 to 2001. Likewise, for the models with an adjustment for union status, R^2 declined steadily from .461 to .401. This trend is consistent with much of the literature on changes in inequality over this time period, wherein the unexplained variance of earnings increases throughout the labor market (see Morris & Western, 1999). In our models, this increase in "within-group" inequality is part and parcel of our finding that industry location is less predictive of individuals' earnings.

Second, after estimating the models that generated Fig. 2, we sought confirmation that the trends in the industry coefficients reflect our expectations of a relative decrease/increase in the industry effects of historically higher/lower paying industries (see Slichter, 1950). We therefore estimated a pooled model that constrained the industry effects to linear trends across years.[10] In this model, we confirmed that historically higher paying industries (such as mining, motor vehicles and equipment, aircrafts and parts, chemicals, petroleum and coal, and communications) had declining relative industry effects while historically lower paying industries (such as miscellaneous manufacturing, leather and leather products, private household services, personal services, health services, educational services, and public administration) had increasing relative industry effects.

In sum, the evidence summarized in Fig. 2 is consistent with a general decline in the advantages (and disadvantages) that workers of all types capture (or suffer) merely by finding themselves employed by firms in different industries. This apparent decline has implications for how the post-industrial economy is evolving, but a uniform decline could not have contributed to the growth in inequality between social classes.[11] In the next section, we investigate this decline across social classes, thereby directly examining the implications of the rent-destruction conjecture outlined in the introduction.

[10] The pooled model had 130 parameters: an intercept and 45 industry dummies, education (and its square), experience (and its square and cube), three race dummies, three region dummies, a marriage dummy, and gender (along with interactions between gender, marriage, education, and experience). All of these variables were then interacted with a linear term for time, which was also included as a main effect. The pooled model had an R^2 of .433.

[11] The decline might plausibly be interpreted as a beneficial transfer to consumers, as more perfect competition within the markets for final goods would lower prices for all goods. Indeed, this would be the position of Hayek (1944:46) who argued that industry rents of the past have been little more than a tax on the consumer.

5.3. A greater relative decline in the dispersion of industry wage premia for classes with declining earnings

First recall that there has been a relative decline in the typical wages of workers in EGP classes V, VI, and VIIa relative to those in classes I, II, IIIa, and IIIb (see Fig. 1). By estimating Eq. (3) separately for each EGP class and then calculating yearly within-class estimates of the employment-share-weighted standard deviation of industry effects, we can assess whether or not there has been a sympathetic relative decline in the dispersion of industry effects for those employed in EGP classes V, VI, and VIIa. If so, then the wages of the working class have not only fallen, but they have converged around lower levels unrelated to industry position.

Fig. 3 presents 5-year moving averages of the weighted standard deviation of industry effects within EGP classes. As shown in panel (a), without adjusting

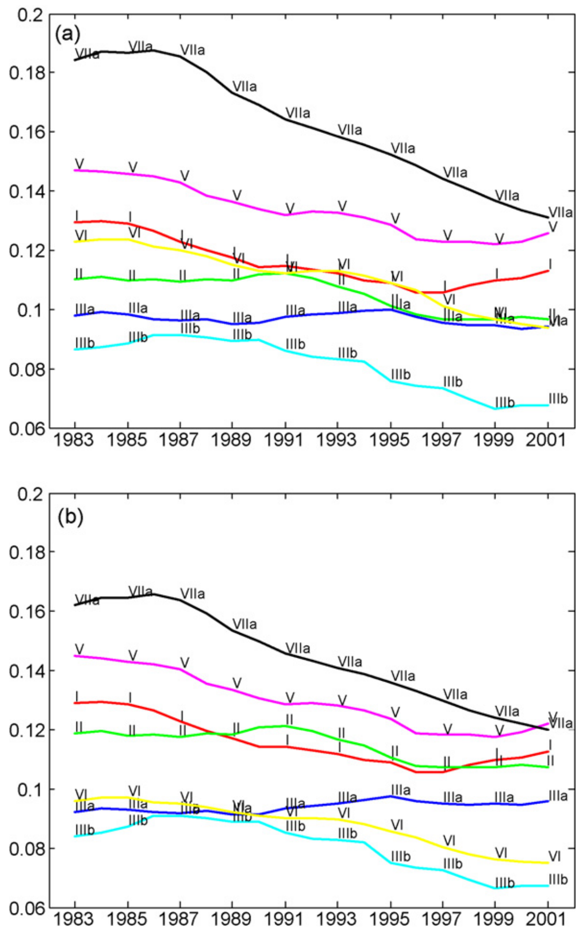

Fig. 3. Five-year moving averages by EGP class of the weighted standard deviation of industry effects (and where panel b includes a further adjustment for union status).

for union status the weighted standard deviation of industry effects declined by 14% for class V (.147 to .126), by 24% for class VI (from .123 to .094), and by 29% for class VIIa (.184 to .131). Trends for the other classes were less clear but were in general more modest. For classes I and II, the weighted standard deviation declined by 13% (from .130 to .113 and .110 to .096, respectively). As noted by Sørensen (2000:1552), some decline in these sorts of industry effects should be present as well, just not as much as for workers at the bottom on the class distribution.

For class IIIa, the weighted standard deviation declined by 4% (from .098 to .094), and for class IIIb it declined by 23% (from .087 to .067). Class IIIb is hard to interpret for several reasons. It is small (less than 5% of the sample), and it is not spread across the full distribution of industries. As described in the last two paragraphs of Appendix A, we had to drop five industries from consideration for class IIIb, reducing the number of industry effects from 46 to 41. Moreover, class IIIb includes a slightly more heterogeneous mix of employment contracts among its members. It is dominated by sales occupations, for which payment by commission is common, and yet the two largest occupations are cashiers and receptionists. Thus, although it is not unreasonable to regard class IIIb as reflecting one constituency within the working class (that is, along with classes V, VI, and VIIa), we will nonetheless de-emphasize findings based on class IIIb because of these concerns about comparability.[12]

As shown in panel (b) of Fig. 3, adjusting for union status does not change the basic trends. The declines are still evident and in the same basic pattern, although, as for the trends in Fig. 2, they are less dramatic. Moreover, in every year, the apparent industry effects characterizing classes VI and VIIa are reduced, as these are the most heavily unionized classes.

We therefore conclude that the evidence, at least in direction, is consistent with the selective rent destruction narrative for increases in inequality in the 1980s and 1990s. The average relative gain from being in a relatively high paying industry instead of a relatively low paying industry declined more for classes V, VI, and VIIa than for classes I and II. At the same time, the average earnings of workers in classes V, VI, and VIIa declined relative to those of workers in classes I, II and IIIa. Thus, there is some evidence that the wages of those in classes V, VI, and VIIa converged on lower levels, suggesting that either the number of jobs in profitable firms declined for workers in classes V, VI, and VIIa or that these workers received less of the profit accruing to firms. Class IIIb is somewhere in the middle, although it is hard to interpret for the reasons just mentioned.

Two questions remain. Does the EGP class schema have a claim to uniquely reveal these trends? Are any such trends, such as those in Fig. 3, an artefact of sampling error? In our remaining two sections of results, we justify the same answer for both questions: “Probably not.”

5.4. *An assessment of the unique explanatory power of the EGP class schema*

As shown earlier in Fig. 1, the growth in inequality can be depicted as a growth in the dispersion of mean earnings across educational groups, social classes, or census occupational groups (and indeed in many other ways; see Katz & Autor, 1999). The evidence for the relative convergence of net industry effects has been presented so far only with reference to the EGP class schema.

Fig. 4 presents trend lines analogous to those in Fig. 3 but with major occupational groups of the U.S. census rather than with EGP classes serving as the cross-sectional partition. In Fig. 4, the weighted standard deviation of industry effects declines more for the occupational groups one might consider to be lower classes (service occupations along with operators, fabricators, and laborers) in comparison with upper classes (managerial and professional occupations). Again, since there is some correspondence between the EGP classes and the major census occupational groups, this is not a surprising finding. But, at a very coarse level, major census occupational groups reveal the same basic story as EGP social classes. The wages of workers at the bottom of the distribution of earnings declined relative to those at the top, and this relative decline in the level of wages was accompanied by a decline in the advantages and disadvantages formerly associated with employment in alternative industrial sectors of the economy.

Fig. 4 suggests that the EGP class schema does not have a claim to uniquely reveal a decline in industry effects. It is unclear whether this result should be seen as support for the claim of Sørensen (2000) that social class analyses of inequality could benefit from the abandonment of coarse, occupation-based measures of social

[12] Again, we do not report full results for these models, since Fig. 3 summarizes 12,236 industry effects (46 industries by 7 classes by 19 years by 2 specifications). However, it is again noteworthy that the R^2 values decline steadily through time, and especially so for class VIIa. In particular, R^2 declines between 1983 and 2001 from .36 to .29 for class I, from .30 to .28 for class II, from .33 to .27 for class IIIa, from .31 to .27 for class IIIb, from .36 to .26 for class V, from .34 to .25 for class VI, and from .43 to .28 for class VIIa.

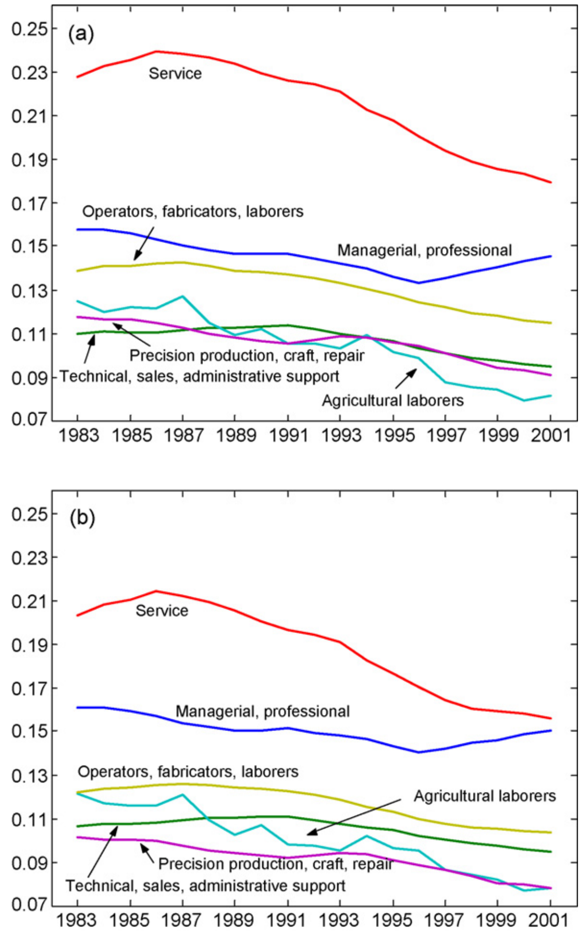

Fig. 4. Five-year moving averages by census occupational category of the weighted standard deviation of industry effects (and where panel b includes a further adjustment for union status).

class (see also Grusky & Weeden, 2001, 2002). Our position is that the EGP schema does generate reasonable results in this application.

What about the third depiction of the growth in inequality from Fig. 1? Has there been a relative decline in the dispersion of industry coefficients for less educated workers? As we show in four figures in Appendix B (available by request and posted on the website of the first author), the standard deviation of industry coefficients declines more for less educated workers when analogous models are fit for educational groups rather than EGP classes or census occupational groups. These patterns hold (and, in fact, become more pronounced) when adjustments are applied for union status, EGP classes, and major census occupational categories.

This is not a surprising finding because the EGP classes for which the dispersion of industry coefficients declined are also on average the least educated classes. Even so, the direct relevance of these patterns across educational groups for the rent destruction conjecture is unclear. The findings for educational groups are hard to interpret while remaining true to the literature that establishes industry differentials as indicators of industry rents. For consistency with that literature, when estimating Eq. (3), education must be included among the observable attributes $\mathbf{A}_i$ that are used to partial out individual-level determinants of earnings. And, as a result, education cannot simultaneously serve to partition the sample into alternative groups for which Eq. (3) is estimated.

5.5. An assessment of the consequences of sampling error and model uncertainty

Could the greater relative decline in the dispersion of industry effects for lower classes be an artefact of sampling error? In order to address this possibility, one must determine how often, by chance, one would expect to generate a set of data in which there is a greater relative decline in the weighted standard deviation of industry effects in classes such as VIIa instead of class I or II, and so on.

In order to make such a determination, the sampling variance of the trend lines in Figs. 3 and 4 must be estimated across hypothetical alternative samples. We generated this distribution using a bootstrap-type procedure, which amounted to developing 10,000 versions of Figs. 3 and 4 consistent with the amount of expected noise in the observed data. As described in Appendix A, we first simulated 10,000 sets of industry effects by sampling from the posterior distribution of industry effects (assuming an uninformative prior distribution and a linear trend in the within-class effect of each industry on wages). We then calculated the employment-share-weighted standard deviations for these 10,000 sets of industry effects and a corresponding set of linear versions of Figs. 3 and 4 (i.e., using best fitting linear trend lines rather than the non-parametrically smoothed lines in the original figures).

Table 3 presents the means, standard deviations, and 2.5th and 97.5th percentiles of these class-specific linear trends in the weighted standard deviations of industry effects (separately in four panels analogous to those in Figs. 3 and 4). All entries of the table are multiplied by 100 to save space. Accordingly, the −.126 in the first cell of Table 3 is 100 times the mean linear decline of the weighted standard deviation of industry effects for EGP class I without adjusting for union status (i.e., the mean of the slopes for the linear decline in the weighted standard deviation for class I across 10,000 hypothetical

Table 3
Descriptive statistics of the simulated distribution of the linear decline in the weighted standard deviation of industry effects by social class

	No adjustment for union membership				With adjustment for union membership			
	Mean decline	S.D.	2.5th percentile	97.5th percentile	Mean decline	S.D.	2.5th percentile	97.5th percentile
EGP class								
I	−.126	.051	−.226	−.026	−.123	.051	−.224	−.021
II	−.099	.047	−.191	−.006	−.084	.050	−.184	.014
IIIa	−.030	.045	−.118	.061	−.010	.047	−.084	.102
IIIb	−.150	.071	−.288	−.010	−.141	.068	−.278	−.008
V	−.155	.048	−.248	−.064	−.171	.047	−.264	−.079
VI	−.171	.048	−.264	−.076	−.130	.046	−.224	−.042
VIIa	−.354	.041	−.434	−.275	−.287	.039	−.365	−.211
Major census occupational group								
Managerial and professional	−.110	.050	−.210	−.010	−.090	.054	−.197	.014
Technical, sales, administrative support	−.097	.068	−.231	.038	−.077	.066	−.207	.055
Service	−.349	.061	−.468	−.229	−.358	.060	−.476	−.238
Precision, production, craft, and repair	−.136	.040	−.214	−.057	−.125	.039	−.201	−.049
Operators, fabricators, and laborers	−.167	.048	−.263	−.073	−.138	.048	−.233	−.044
Agricultural labor	−.278	.129	−.537	−.025	−.271	.131	−.534	−.017

Notes: All coefficients are multiplied by 100.

figures). As such, it is 100 times the best fitting linear slope of the trend line for class I in panel (a) of Fig. 3. Consistent with the general pattern of Fig. 3, this mean linear decline for class I reported in Table 3 is closer to zero than the relevant mean linear declines presented in the same column of Table 3 for classes V, VI, and VIIa (i.e., −.126 versus −.155, −.171, −.354). Classes II and IIIa are characterized by even less of a decline, and class IIIb is hard to interpret because of its small size and limited spread across industries.

Table 3, however, also gives the posterior predictive uncertainty associated with these linear declines. For example, for 95% of the sets of industry effects consistent with the data, the mean decline for class I falls within an interval from −.226 to −.026. For class VIIa, in contrast, the same interval is shifted to the left, stretching from −.434 to −.275. In this case, the upper end of the interval for class VIIa is more negative than the lower end of the same interval for class I.[13] For other pair-wise comparisons between classes, there is more overlap in the distributions but still considerable precision in the estimates of the weighted standard deviations.

There is no ironclad proof against the claim that some of the patterned relative decline present in Figs. 3 and 4 has been produced by sampling error. On balance, the table suggests that the findings most relevant to the rent destruction conjecture (that is, the decline in the variance of industry differentials for class VIIa relative to other classes in Fig. 3) are unlikely to have been generated solely by sampling error. That classes IIIb, V, and VI fall somewhere in between classes I and VIIa is also encouraging for the conjecture, but there is no strong evidence in favor of a relative ranking of these classes because sampling error and model uncertainty are non-trivial.

6. Conclusions

To evaluate the claim that selective rent destruction is one component of the recent increase in earnings

[13] We invite readers more inclined to take a frequentist view of statistical inference to form quasi *t*-tests by (1) considering the simulated standard deviations in columns 2 and 6 as standard errors, (2) subtracting any two coefficients from another, and (3) dividing each resulting difference by the square root of the summed squares of the relevant standard deviations. For example, divide .00228 (i.e., −.00126 + .00354) by .00065 (i.e., $\sqrt{.00051^2 + .00041^2}$), resulting in a quasi *t*-statistic of 3.51. This *t*-statistic is duly labeled quasi as the simulated standard deviation incorporates model uncertainty along with sampling error. As such, the resulting ratio is not distributed as Student's *t*, and is best interpreted as more conservative than a *t*-ratio.

inequality and more generally to evaluate whether structural analysis of trends in labor market inequality continue to have explanatory power, we analyzed yearly data from the Outgoing Rotation Groups of the Current Population Surveys from 1983 to 2001. We presented two sets of findings. The relative earnings of those at the bottom of the class distribution declined over time, as did the variance of wage premia associated with employment in alternative industries. We conclude that these results jointly support the two implications of the rent destruction conjecture outlined in the introduction. Those implications are grounded in the sociological and labor economics literature that argues that net industry wage premia can be interpreted as rent payments. Accordingly, we infer from our results that the bottom of the labor market now contains fewer privileged positions in which an individual can obtain what might be regarded as above-market wages. As a result, a case can be made that Sørensen is correct: a disproportionate share of at least one type of structural advantage enjoyed by some lower class workers has been eliminated in the last two decades.

7. Discussion

In this section, we discuss six explanations for the findings that we present, some of which are complements to the rent destruction narrative. We conclude with a statement on the implications of our results for the structural tradition of labor market analysis in sociology.

7.1. Plausible complementary and alternative explanations

Among the following six plausible explanations for the pattern of findings we present, three can be represented as alternatives to the rent destruction narrative:

1. The industry classification of the census bureau no longer captures the dimensions of the industrial location along which industry rents are evolving. Although this claim is plausible and indeed may account for the general decline in industry effects shown in Fig. 2 (and perhaps for the decline in industry effects for class I in Fig. 3), we see no reason why the industrial structure should have become relatively worse at picking up good and bad jobs for the working class. Our intuition suggests the opposite, as dynamic job growth has been more pronounced at the top of the class distribution.
2. Net industry wage differentials may reflect nothing more than selection bias on unobserved ability. If individuals with relatively high unobserved abilities are allocated by market forces to positions that have unobservable skill demands, movement in industry wage premia may reveal nothing other than changes in the market for unobserved skills. We do not see this explanation as a strong threat to our conclusions. It does not provide a unique explanation for the trends we have observed, as there is nothing in the rent destruction conjecture that excludes mechanisms relating to the market for unobserved ability.[14] Indeed, if this explanation is true, it could be that firms with rent to share may have remained willing to use some portion of that rent to attract relatively high ability individuals to management positions. At the same time, these same firms may have become less willing to use that rent to attract relatively high ability individuals to positions at the bottom of the organizational chart. Exactly why such a cleavage in hiring strategies may have emerged then demands its own explanation, one which need not be distinct from the basic components of the rent destruction narrative offered later.
3. The findings are an artefact of changes in educational institutions and patterns of selection into them. If educational institutions have become better at sorting students by levels of ability and potential productivity, the rate of return on education may have increased while market imperfections such as industry wage premia have declined. Individuals who would have been relatively highly paid members of the working class because of inherent ability may now more likely have been sponsored to higher levels of educational and occupational attainment. As such, there is more homogeneity of ability within the working class, and there are fewer potential individuals of relatively high ability who are willing to take employment in low-skill jobs. As for the prior alternative explanation, we know of no evidence that supports this position (such as an increase in the relationship between college graduation and cognitive skills measured in early adolescence). But, it is clearly an explanation deserving of investigation.

   The next three explanations can be interpreted as rival explanations or, as we prefer, central components of the rent-destruction narrative that identify plausible ultimate causes:

[14] Furthermore, we do not know of any evidence that supports it. Rather, it is somewhat inconsistent with the expressions of employers that relatively high ability individuals willing to take entry level jobs are hard to find and hence worth enticing with special accommodations (see Moss and Tilly, 2001, Chapter 3; Rosenbaum, 2001, Chapter 5).

4. International trade (perhaps coupled with deregulation) may be an ultimate cause of the patterns we observe, either by (a) eroding the bargaining power of the working class through a relative decline in the number of jobs for low-skill workers or (b) by eliminating some of the rent formerly earned by quasi-monopolist producers (see Katz & Autor, 1999:1536–1538). If trade is an ultimate cause of the decline in industry rents for the working class and (b) is true, then either (a) is also true or some other process must be invoked to explain why workers' negotiated rent would have declined relatively more at the bottom of the class structure. For Bourdieu, no separation is possible, as the "individualization of the wage contract" and attendant "institutionalized precariousness" are made possible because national labor markets have been drawn into a competitive global field (see Bourdieu, 2003[2001]:28–29 and 91–96).
5. There may have been an exogenous decline in union bargaining power, and this decline in union power may be unrelated to the size of the rent generated by firms. If this explanation is anything other than a mechanistic elaboration of explanation (4), the decline in union power would have to be one based on something other than a weakened bargaining position, and it must be attributable to something such as poor union strategy or an erosion in government protection of the right to organize (see Bronfenbrenner, 2000).
6. The working class has been disproportionately harmed by basic changes in employment relations, both in the norms by which wages are set and in organizational employment schemes (see Appelbaum, Bernhardt, & Murnane, 2003; Powell, 2001). The ultimate cause of these changes may be the greater responsiveness from the 1980s onward of managerial decisions to the value of publicly held shares, as argued effectively by Dore (2000), Fligstein (2001), and Sørensen (1996, 2000). And yet, for this explanation to hold, one must explain why the working class has become relatively less able to defend against the evolution of these new employment relations, perhaps relying upon explanations 4 and 5.

In sum, in our judgment, explanations 1 through 3 are not strong alternatives to the rent destruction narrative, and explanations 4 through 6 represent ultimate causes that may have produced rent destruction. Insofar as explanations 4 through 6 invoke deeply intertwined causes, they may be so inextricably linked that attempts to definitively separate them with observational data will founder.

Our view on this matter represents a realistic stance toward the challenges of causal inference from limited observational data. Structural advantages are elusive objects of investigation because they are counterfactually defined with reference to social structural positions rather than the individuals who occupy them. And, any such structural effects are probably generated by heterogeneous processes only weakly predictable by the sort of aggregate variables that can be constructed with currently available data.[15] However, it could be the case that more fine-grained data on the dimensions of social structure would reveal structural effects of considerable size which can be attributed to specific ultimate causes, as we discuss next.

7.2. *Implications for structural analysis of labor markets in sociology*

For work that continues to trace its origins to the new structuralism (see Kalleberg et al., 1996, 2003), a subtle accounting of the distribution of earnings inequality is now offered. Firms operate in competition with each other, but also under institutional and regulatory constraints that apply to all firms and which shape their human resource practices. Workers compete within organizations for promotion opportunities, once they have achieved some degree of protection from competition with those outside of the firm. Increasingly since the 1980s, workers have had to compete with a pool of casual labor, which represents the post-industrial equivalent of the classical Marxist reserve army of the unemployed. A casualization of the employment relationship results, in which the weakened bargaining position of workers hurts their wages and fringe benefits.[16]

[15] One might argue that, if this position is correct, then unique structural effects on wages are probably somewhat small in comparison to those which can be accounted for fairly well with individualist human capital investment and efficient labor market matching models (under the assumption that individuals can forecast, with some degree of accuracy, the relative attractiveness of alternative structural positions and then strive to enter them). We are comfortable with this position, as it furnishes a convenient explanation for why structural models have never been able to achieve clear empirical supremacy over human capital and other alternatives. That being said, even if uniquely identified structural effects are relatively small, it does not follow that they are necessarily unimportant, especially as explanations of exogenous overtime change that occur in intervals too small for individuals to forecast and then respond to.

[16] This conception of the labor market remains very much in line with the original conception of Sørensen and Kalleberg (1981), where

In this article, we implicitly adopt this framework. Unfortunately, we cannot offer direct support for it with the CPS data, since precious little is known about the particular employment contracts under which CPS respondents work. Nonetheless, our results should be regarded as encouraging for the prospects of structural models in this tradition, as well as further justification for the new forms of data collection that the new structuralists have launched in collaboration with scholars of organizations. The multivariate new structuralist approach advocated by Kalleberg and Berg (1994), if applied to our research, suggests that we should partial out structural effects across overlapping work structures by specifying cross-sectional and over-time differences in the features of firms within our 46 industries. Kalleberg et al. (1996), in reporting on the National Organizations Survey, offer a comprehensive accounting of the many features of organizational structures that may have consequences for individuals' wages. In principle, if such a survey of firms within the 46 industries we analyze were available for every year and were of sufficient size, then we could attempt to decompose the trends in industry effects that we document, using models that examine the changing relative power of firms and workers as they interact with shifting compensation practices. For now, such data are not available. But, we would hope that our results can serve as further justification for investment in their future collection, since our results suggest that there are potential structural effects worthy of more fine-grained examination.

One complementary strand of research also deserves mention in this regard, in part because it supports our interpretations but also because it gives support to the new structuralist literature as well. Addressing the weaknesses of the inter-industry wage premia literature, John Abowd and his colleagues in economics have developed a set of models for longitudinal matched employer-employee data (see Abowd & Kramarz, 1999; Abowd, Kramarz, & Margolis, 1999). With confidential data from the State of Washington, Abowd and his colleagues have estimated that, on average, 50% of the raw inter-industry effects estimated from models such as the one represented in our Eq. (3) is a function of unobserved individual-level heterogeneity. We regard this as somewhat encouraging. A full 50% of raw inter-industry effects is a function of firm-level characteristics, and hence adjusted industry effects remain amenable to a rent-sharing interpretation.[17] Even though we therefore have some confidence that this new research from labor economics suggests that our 46-fold industry coding is reliable, it also demonstrates that firm-level analysis is important and that further work following upon Kalleberg et al. (1996) is crucial for advancing knowledge about the contours of inequality in post-industrial labor markets.

employment relationships are given an ideal type characterization as either "open" or "closed." The full range of open and closed employment relationships is now known and fully integrated with the literature in the sociology of organizations. Moreover, as reviewed by Kalleberg (2001), the share of open employment relationships has expanded so much that a new vocabulary for describing nonstandard employment relations has been invented.

Appendix A

Calculating replicate distributions of trends in class-specific industry differentials

In order to generate a replicate distribution within each class of the weighted standard deviation of industry effects, we chose to create 10,000 replicate distributions of the underlying industry effects and then to form 10,000 corresponding weighted standard deviations as deterministic summaries of these underlying replicate distributions. We follow the tradition of posterior predictive simulation in Bayesian statistics, and in particular the simple Bayesian treatment of classical regression (see Chapter 14 of Gelman, Carlin, Stern, & Rubin, 2004).

To construct a posterior from which to simulate industry coefficients, we introduce the assumption that, for each industry k, the industry effects follow an underlying linear trend around which they are normally distributed with a constant variance. That is, for a design matrix $\mathbf{T}$ filled with ones in its first column and an index for time in its second column, we assume that for each industry k, the industry effects are $\mu_k \sim N(\mathbf{T}\boldsymbol{\tau}_k, \sigma_k^2)$. Maintaining an uninformative joint prior for τ_k and σ_k^2, we then generated 10,000 replicate industry effects with the following 5-step simulation algorithm within each class:

1. *Generate observed industry effects from individual-level data*. For each year t, estimate Eq. (3) and collect a set of industry-and-year specific coefficient estimates in $\hat{\mu}_{\mathrm{kt}}$.
2. *For each industry, find the best linear predictor of the industry effects*. For each k, specify $\hat{\mu}_{\mathrm{kt}}$ as the

[17] And, since we argued earlier that the changes in the market for unmeasured skills are not necessarily inconsistent with the rent destruction conjecture, we regard the other 50% of the apparent industry effects as still amenable a rent destruction explanation as well (see explanation 2 earlier).

outcome variable and $\mathbf{T}$ as the predictor matrix. Then, use ordinary least squares to estimate $\hat{\boldsymbol{\tau}}_k$, which, as a result, contains the intercept and slope parameters of the best linear predictor.

3. For each industry, sample from the joint posterior distribution of σ_k^2 and $\boldsymbol{\tau}_k$. For each k, draw 10,000 values from the marginal posterior distribution for σ_k^2 (i.e., draw 10,000 values from a scaled inverse chi-squared distribution with parameters $t-2$ and the degree-of-freedom-corrected error sum of squares from the k-specific regression of the last step). For each draw from the marginal posterior distribution for σ_k^2, draw conditional values for the elements of $\boldsymbol{\tau}_k$ (i.e., draw values from a bivariate normal distribution with expectation $\hat{\boldsymbol{\tau}}_k$ and variance $(\mathbf{T}'\mathbf{T})^{-1}\sigma_k^2$, where the σ_k^2 is the draw from the marginal posterior distribution in the last sub-step).
4. *Generate a posterior predictive distribution for the industry effects*. For each of the 10,000 sets of draws from the joint posterior distribution of $\boldsymbol{\tau}_k$ and σ_k^2, generate μ_{kt}^{rep} from a $N(\mathbf{T}\boldsymbol{\tau}_k, \sigma_k^2)$. The resulting 10,000 replicates of the predicted industry coefficients reflect the inherent uncertainty of the joint posterior distribution (and the uninformative prior distribution it implicitly absorbs) along with the predictive uncertainty of the assumed linear model for the industry-specific time path.
5. *Calculate a replicate distribution of the weighted standard deviations of industry effects*. For each year t, calculate 10,000 weighted standard deviations of industry effects using the replicated values for μ_{kt}^{rep}. To do this, first generate 10,000 replicate industry distributions to mimic an underlying sampling process. For each industry k and each year t, draw 10,000 random values from a Poisson distribution with mean n_{kt} and then randomly pair each random industry distribution with a set of replicate industry coefficients.

Because the data are sparse in some dimensions, the process differs slightly by class. Within each class, we dropped industries from further analysis if no coefficient could be estimated in step 1 for 7 or more of the 19 total years. This is equivalent to dropping industries if the number of respondents n_{kt} in industry k in year t is equal to 0 for less than 12 of the 19 years analyzed. The consequences of this restriction were not severe. We were able to estimate industry effects for all 46 industries in all years for classes I, II, IIIa, and VIIa (i.e., 874 coefficients for each class). For class IIIb, we dropped five industries, and for class V we dropped 1 industry. For class VI, we had to impute a few coefficients when developing the non-linear trends for Fig. 3 (as described in the next paragraph), but we were able to keep all 46 industries in the analysis.

In order to plot the non-linear trends in the weighted standard deviations of industry coefficients for Fig. 3, we had to impute 27 of the 6004 coefficients that generated the summary standard deviations presented in the figure. For industries for which it was not possible to estimate a coefficient for all 19 years (but which survived the prior step by passing the 12-or-more-years-limitation), we adopted the following simple imputation scheme. After enacting the same first two steps of the simulation algorithm, we (1) identified the predicted value for the missing year t^* from the best linear predictor, and then (2) added noise to this predicted value with a random draw from a normal distribution with the variance equal to the degree-of-freedom-corrected error sum of squares from the relevant least squares regression. With this procedure, we imputed 21 industry coefficients for class IIIb (i.e., 21 out of 779 coefficients), 3 coefficients for class V (i.e., 3 out of 855 coefficients), and 3 coefficients for class VI (i.e., 3 out of 874 coefficients). Since this was a rather small proportion of coefficients, we did not use the more sophisticated simulation from the posterior distribution in our imputation routine. We therefore maintain a rigid assumption in the imputation scheme (that the slope and variance of the underlying regression are known after they are estimated in step 2), but we do so with little consequence since the small number of imputations had virtually no influence on the weighted standard deviations that are the main focus of the empirical analysis.

Appendix B. Supplementary data

Supplementary data associated with this article can be found, in the online version, at doi:10.1016/j.rssm.2007.08.003.

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