Education and its Discontents: Overqualification in America, 1972-2002

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The study of education-occupation mismatch, once central to the sociological investigation of the labor market, has been largely abandoned. While labor economists and scholars in other nations continue to investigate overqualification, it has been more than two decades since its last sociological assessment in the United States. Drawing on previous work and guided by Bourdieu's concept of habitus, I hypothesize that workers who have more educational attainments than needed for their jobs will be less satisfied with their jobs, be more politically liberal, and be less likely to endorse an effort-based achievement ideology. Using the 1972-2002 General Social Survey, I find that overqualification has increased substantially, that the expected effects are generally found, and that these effects remain relatively stable over time. I discuss the implications of these findings for understanding the persistence of existing stratification hierarchies.

The study of mismatch between educational and occupational attainment, once central to the sociological investigation of the labor market, has now been largely abandoned. Burris (2005) recently noted that while interest in this phenomenon has not waned among all social scientists, sociologists (and American sociologists in particular) have contributed little to our current understanding. Of 42 articles published on overqualification[1] in the previous six years (as listed in the Social Science Citation Index), only four have been by sociologists and only three have treated the United States. For interesting international work on the subject see, Batenburg and de Witte 2001; Brynin 2002; Chauvel 2002; de Grip, Heijke and Willems 1998; Livingstone 1998; Lowe 2001).

Overqualification – which occurs when a worker has more education than is required for the performance of his or her job – is an important social problem for several reasons. At a macro level, it deals with the intersection of two of the major institutions of modernity – education and the labor market. Though in recent years this terrain has been largely ceded to economists and human capital theorists (Burris 2005), sociologists have intellectual resources that make them well suited to understanding these processes as well (Gibbons 2005). Rather than assuming an unproblematic or automatic link between educational institutions and the labor market, overqualification research makes the relationship between them an empirical problem. As Berg (1971) argued in his seminal (and newly reissued [2003]) book, *Education and Jobs: The Great Training Robbery*,

I would like to thank Arne Kalleberg, Ken Bollen, Barbara Entwisle, David Autor, Rebekah Vaisey and the Social Forces *reviewers for their help and insights. A National Science Foundation Graduate Research Fellowship supported me during this project.* *Direct correspondence to Stephen Vaisey, Department of Sociology, CB #3210, University of North Carolina, Chapel Hill, NC 27599. E-mail: vaisey@unc.edu.*

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educational expansion may have unintended and negative consequences for society. To take one example, Berg (1971) argued that employers' increasing reliance on educational credentials for hiring can place severe limitations on the achievement of disadvantaged groups (see also, Burris 2005). At a micro level, overqualification is important because it matters for the way individual actors experience their lives. Work is not only about obtaining money or other material goods, but also about creating and sustaining worldviews and identities (Akerlof and Kranton 2005; Sayer 2005).

Though labor economists have done excellent work developing more precise models to estimate the effects of overqualification on economic outcomes (Sicherman 1991; Cohn and Kahn 1995), there is much we do not know about its effect on social and political outcomes. By contrast, Berg (1971) looked at a host of worker outcomes, such as job satisfaction and political orientations. Burris (1983) continued this multidimensional approach with an assessment of the social consequences of overqualification, but his work is now more than two decades old. The time has come for an update. This paper therefore has two objectives: first, to assess the prevalence and social distribution of overqualification in the U.S. labor market between 1972 and 2002; and second, to explore the effects of overqualification on several economic, social and political outcomes. Based on these analyses, I conclude that the incidence of overqualification has substantially increased among all social groups since 1972, that it has significant effects on diverse outcomes, and that these effects have not diminished with time.

Theory and Previous Research

Many theorists have considered the question: "Where does overqualification come from?" This question has been intensely debated among and between advocates of human capital theory (Becker 1993; Freeman 1976) and those favoring a credentialist view of the labor market (Berg 1971; Brown 2001; Collins 1971, 1979; Livingstone 1998, 1999, 2001). The intractability of this debate stems from the fact that both major schools of thought – human capital theory and credentialism – make similar macro-level predictions (i.e., a strong link between educational and occupational attainment) but rely on vastly different explanatory mechanisms. I focus here on the distribution and consequences of overqualification in the labor force.

Though overqualification was once the object of intense interest among sociologists, one probable reason for declining interest is that it failed to live up to the theoretical predictions originally associated with it (Burris 2005). From the late 1960s, social observers from radical Marxists to government planners linked the coming glut of college graduates to predictions of wholesale dissatisfaction and revolt against the system that blocked their advancement (Berg 1971; Bowles and Gintis 1976; Carnegie Commission on Higher Education 1973; Gorz 1967; see also Burris 1983). Blumberg and Murtha (1977:46) went so far as to call it "the central dynamic for social discontent emerging in America." While early research seemed to support the alarmist view, subsequent findings (not to mention relative political calm) slammed the door on it. Burris' (1983) analysis yielded little evidence for the radicalizing effects of overqualification (see also

Glenn and Weaver 1982). A few years later, Smith (1986) showed that the fear of declining economic returns to education (advanced primarily by Freeman [1976]) was also unfounded. After a brief downward blip, the economic returns to a college degree had not only remained stable, but had in fact *risen* over time. Thus the two main predictions of overeducation's theoretical edifice had crumbled under the weight of the evidence. Moreover, these predictions had come with no explicit micro-level mechanism that would explain why a discrepancy between educational and occupational attainments – net of wage returns – would matter to workers in the first place.

There is no single, unitary framework that will allow us to theorize the micro-level effects of overqualification. Yet, where this subject is concerned, seemingly disparate theories point in complementary directions. The work of Pierre Bourdieu, especially his notion of habitus, can provide a useful framework for understanding the how a mismatch between educational and occupational attainments might cause subjective problems for actors (Bourdieu 1990, chapter 3; Sayer 2005, chapter 2). Though habitus can be a slippery concept, recent work in the sociology of education has profitably operationalized the habitus as occupational aspirations – or more generally, as an active orientation to probable future structural position that has been formed out of one's past structural position (Dumais 2002; McLelland 1990). What does this mean in practice? It means that individuals moving through the educational system come to expect an "appropriate" type of work upon completing their education. By this I do not mean simply a high paying job, but rather one that will allow the "right" kind of income, working conditions, associates and identity, to name a few (see Sayer 2005, chapter 2; also Akerlof and Kranton 2005). Thus a lack of fit in the labor market can lead to cognitive dissonance and dissatisfaction when the worker finds himself in a position "beneath" that for which he was prepared (Festinger 1957). In a similar vein, Merton's (1968) "anticipatory socialization" speaks to the general social process of generating occupational expectations.

Given these considerations, the mechanism of the habitus can serve to explain the importance of status consistency to which several theorists have pointed (see Lenski 1954 for a related argument). Bourdieu, Merton, Festinger, Lenski and many others have approached this phenomenon with different theoretical vocabularies, but each points to a similar, basic conclusion: that actors desire to achieve consonance between their various social statuses and that social and cognitive discomfort that can result when they are not successful. This "drive" can perhaps be most profitably seen as grounded in the habitus; that is in one's internalized expectations about future trajectories. This theoretical orientation would lead us to believe that when actors' expectations are not met, dissonance and dissatisfaction will likely result (Festinger 1957). The question remains, however, whether workers will adapt to undesirable circumstances perhaps by blaming themselves (see Burris 1983; Smith 1986) or whether they will channel their dissatisfaction into more collective outlets. With this in mind, the primary emphasis here will be on testing whether overqualification is related to outcomes such as job dissatisfaction, political liberalism or rejecting the narrative of the "American Dream." All of these represent some degree of the externalization

of blame – respectively, toward the employer, the government or widely shared views about the nature of success.

Measurement and Data

Measuring Mismatch

Though overqualification is a seemingly simple concept, it has proven difficult to measure. The problem is determining how much education a job really requires. There are three general approaches to this. The first is to rely on the expert assessment of a job analyst, who rates the requirements of a given job (Burris 1983; Livingstone 1998). The second is to ask workers how much education their job requires (Halaby 1994). The third is to estimate the average level of education in each occupation, and then to classify those who are (say) more than one standard deviation above the mean as overqualified (Clogg and Shockey 1984; Verdugo and Verdugo 1989). In this paper, I rely on the first method, basing my estimates of required education on the General Educational Development (GED) scale developed by the U.S. Department of Labor. The GED scale is really the maximum value of three scales – reasoning, math and language (Livingstone 1998). The GED scale does not capture "soft" (e.g., interpersonal) skills, but the presence of the language component implies that it does not measure technical

Table 1: General Educational Development Scale

1	Applies common sense understanding to carry out *simple* one or two step *instructions*. Performs simple counting. Learns job duties from oral instruction or demonstration
2	Applies common sense understanding to carry out *detailed* but uninvolved *instructions*. Uses arithmetic with whole numbers. Files and copies data.
3	Applies common sense understanding to carry out *instructions involving several concrete variables*. Makes arithmetical calculations involving fractions. Files and copies data.
4	Applies principles of rational systems (e.g. bookkeeping, electric wiring systems) to *solve practical problems*. Performs algebraic/geometric procedures. Interprets technical manuals.
5	Applies principles of logical or scientific thinking to *define problems* and interpret extensive technical instructions. Applies advanced mathematical techniques. Evaluates technical data and writes reports.
6	Applies principles of logical or scientific thinking to a *wide range of problems* and applies them to a variety of abstract and concrete variables. Applies advanced mathematical techniques. Evaluates technical data and writes reports.

Source: Livingstone (1998, Figure 1.1)

proficiency alone. Many jobs with a high language requirement necessitate, for example, the ability to communicate complex concepts to others. The values and definitions of the GED scale are given in Table 1.

My choice of the job analyst method requires some explication. The general consensus in the literature is that the job analyst method and the self-report method are superior to the "plus one standard deviation" method in that they are less arbitrary and non-dichotomous (Halaby 1994; Cohn and Kahn 1995; van der Velden and van Smoorenburg 1997). Analysts have also found that while the job analyst and self-report approaches provide different estimates of the overall level of overqualification in a sample, they yield very similar coefficients as predictors, at least when predicting wages (van der Velden and van Smoorenburg 1997). There is therefore no clear advantage to choosing one method over the other. The current analysis is constrained by data availability. Very few U.S. surveys ask respondents about the educational requirements of their job *and* about non-monetary outcomes of interest. Those that do (notably the 1973 and 1977 Quality of Employment Surveys) are three decades old. Other data sources must therefore be used if we wish to make inferences about the incidence and consequences of overqualification over a longer period.

Data

The data come from the 1972-2002 General Social Survey (GSS). The GSS is a repeated cross-sectional study conducted by the National Opinion Research Center (NORC) that interviews approximately 1,500 non-institutionalized adults in the United States about every year. The data collection procedures have generally remained the same since 1972, although the 1972-74 surveys used a multi-stage area probability sample rather than a full probability sample. This may lead to slightly underestimated standard errors in these years, but correcting for such differences would not change the results. I exclude the 1982 and 1987 oversamples from the analyses, and use the unweighted data. Aggregating surveys from the entire period presents an unparalleled opportunity to examine this issue over a 30-year span. Following previous work in this area (e.g., Burris 1983; Clogg and Shockey 1984; Halaby 1994) I restrict the sample to full-time workers between the ages of 25 and 65. This is not because the effect of mismatch would be less interesting for part-time workers; in fact, a comparison between full- and part-time workers would probably be interesting and important. Yet to maintain comparability with previous research, it seems that focusing on full-time workers is a better choice. With these restrictions, the final sample over the 30-year period is 18,172.

To estimate the educational requirements for each occupation, I make use of data compiled by Autor, Levy and Murnane (2003) to measure changes in skill requirements over time. Autor and colleagues used information from the *Dictionary of Occupational Titles* (DOT 1977, 1991), in combination with special CPS and census samples containing both 1970 and 1980 three-digit census occupational codes, as well as more detailed DOT job codes, to estimate the GED values for 1970 and 1980 census occupations at two time points – 1977 and 1991.[2] These values were separately estimated for men and women, which (because men and

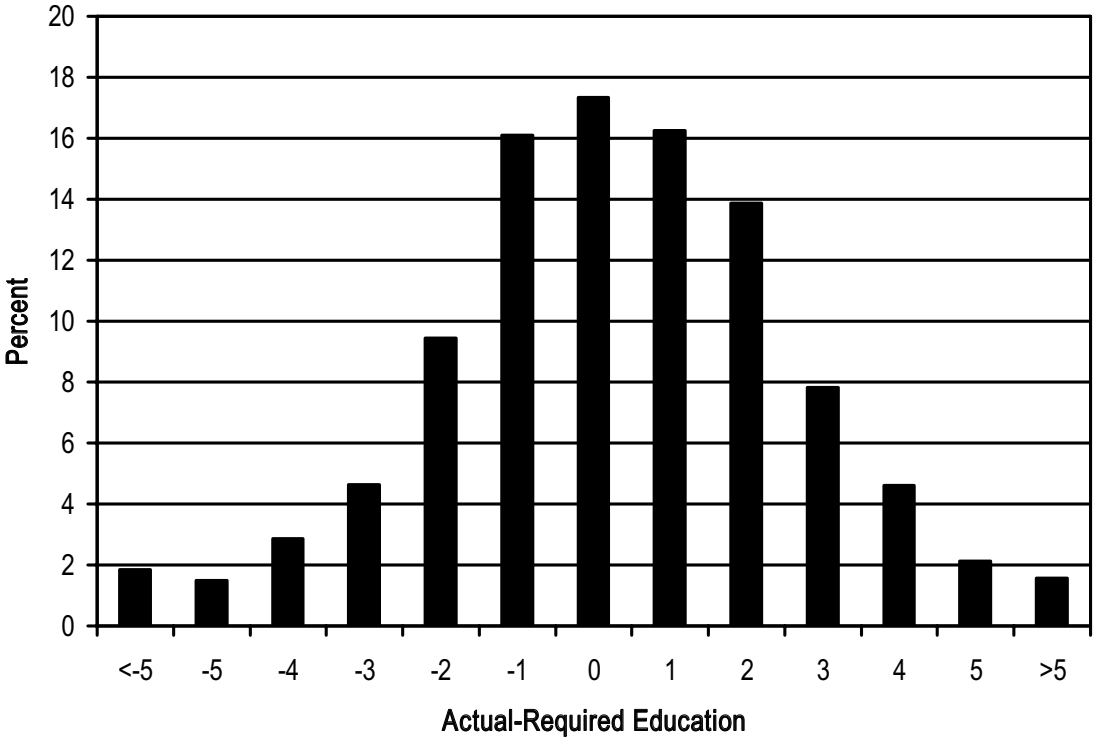
women tend to have different jobs within a given occupational title) reduces the jobs-within-occupation distribution bias that comes with operating at the three-digit census level (Halaby 1994). I attach these values to the 1970 and 1980 three-digit occupational codes available in the GSS. (Where both 1970 and 1980 occupation codes are available [1998-1990], I use the 1980 codes.)

With this information, I have a reasonable indicator of GED educational requirements at two widely-spaced time points. To apply these estimates to other years, I use a linear interpolation, extending the observed average yearly change between 1977 and 1991 back to 1972 and forward to 2002 for each occupation. The rationale behind this is that the change between 1977 and 1991 for a given occupation represents an educational requirements trajectory generalizable to the entire 1972-2002 period.[3] Though this rationale is debatable, it ultimately will have few consequences for two reasons. First, the difference between the 1977 and 1991 overall means is slight – only .11 GED units. Though perhaps surprising, what this means is that relatively little skill upgrading (within or between occupations) took place over that 14-year period (at least according to Department of Labor experts). Second, the 1977 and 1991 GED values are correlated at .99, which means that what skill upgrading occurred almost entirely uniformly across occupations.

To convert the final GED score into an estimate of required schooling, I opt not to use the conversions suggested by Burris, Berg, Eckhaus and others (see Livingstone 1998 for an overview). Though they have several commonalities, these conversions are based on widely divergent assumptions about what kinds of educational distinctions matter and about what level of education is required to do certain tasks. Each therefore produces a very different estimate of the incidence of overqualification. For example, Livingstone (1998) shows that these measures produce estimates of overqualification ranging from 24 to 76 percent of the full-time workforce in 1990. Rather than rely on unsupported a priori assumptions, I derive my GED-to-required education conversion from one of the few data sources that contains both three-digit occupation codes and self-reported job education requirements – the 1977 Quality of Employment Survey (QES). The QES is a national probability sample containing data gathered from 1,515 workers aged 16 and older who were working for pay for 20 or more hours per week in the United States in 1977. To convert GED values to an estimate of required schooling, I regressed the respondent's self-estimated required years of schooling on the GED values from 1977. I did not assume that this relationship was linear and used Bayesian Information Criterion (BIC) comparisons[4] and likelihood ratio tests to evaluate different polynomial specifications. Both criteria confirmed that a third order polynomial offered the best fit (required education = -.52 + 6.96**GED* -1.46**GED*^{2} +.14**GED*^{3}). Though the correspondence between the two is not perfect, the polynomial-transformed GED and self-reported required education are relatively highly correlated (.59, R^2 = .35).

There are several indications that this conversion establishes a good correspondence between GED and required education. First, as Figure 1 demonstrates, the modal difference between actual and required education is 0, centering the conversion on matched workers. Furthermore, half of all workers are

Figure 1. Matching between Actual and Required Education



classified within one year of a perfect match. Second, the full cross-classification of actual and required education (Appendix 1) also shows a significant grouping around the diagonal. Of course, the fact that some categories are more populated than others prevents a perfect correspondence, but the overall fit between the two appears to be a good one. Thus, there is considerable prima facie evidence that my estimate of required education is at least reasonably accurate at the three-digit census occupational level.

At this point, readers may raise two objections. First, there is an assumption of a uniform level of school quality and student performance, and both may have decreased dramatically between 1972 and 2002. Second, because the latest estimate of job requirements is from 1991, the data do not consider the influence of widespread personal computing or other recent technological advancements in the workplace. Both are serious objections discussed in the conclusion where their plausibility is assessed in light of empirical results.

Methods

The first part of the analysis will describe the extent of overqualification in the working population over the 1972-2002 period. In this section, I follow Tsang and colleagues (1991) and use two benchmarks – the simply overqualified (at least one more year of education than required) and the highly overqualified (three or more years of education than required). Because one cannot be sure that those with only one or two years of "surplus schooling" are truly overqualified, I will focus mainly on the highly overqualified.[5] I further decompose the incidence of overqualification by gender, age, race and type of degree.

The second part of the analysis will estimate the effects of education-occupation mismatch on three outcomes predicted to be influenced by overqualification: job satisfaction, left political views, and achievement ideology. In addition, I look at the effect of overqualification on earnings to see if this method of estimation yields similar results to those found by Sicherman (1991) and Cohn and Kahn (1995) using a self-estimate of educational requirements. To measure earnings, I use the natural log of yearly income, adjusted for inflation to year 2000 dollars. To measure job satisfaction, I use the variable SATJOB, and contrast those with full job satisfaction to those without full job satisfaction. For left political views, I use the variable POLVIEWS, and contrast those who call their views "liberal" or "extremely liberal" to those who offer other responses. Finally, to measure belief in achievement ideology (i.e., endorsement of the "American Dream"), I use the response to the survey question GETAHEAD: "Some people say that people get ahead by their own hard work; others say that lucky breaks or help from other people are more important. Which do you think is most important?" The response is coded 1 if the answer is "hard work" and 0 if another response was given. The equation used to estimate mismatch effects has been used for both monetary (Sicherman 1991; Cohn and Kahn 1995) and non-monetary (Tsang et al. 1991) outcomes and is as follows:

$$Y_i = \beta X_i + \alpha E_i^{\mathrm{r}} + \tau E_i^{\mathrm{o}} + \delta E_i^{u} + \varepsilon_i$$

where $\mathbf{X}$ is a vector of independent variables, E^r represents an estimate of the required years of education for the respondent's job, E^o represents the number of years of education beyond the job requirement (or 0, if the respondent is either matched or underqualified), and E^u stands for the number of years of education below what is required for one's job (or 0, if the respondent is matched or overqualified). Thus α is a parameter estimating the effect of the education that is actually used (i.e., the effect of the job itself), τ is a parameter estimating the effect of additional years of education beyond what is required, and δ is a parameter which estimates the effect of the years of underqualification.

This model assumes that a match has occurred when $E^o = E^u = 0$. While this assumption is relatively unproblematic when workers are self-reporting job requirements, some may object that one cannot be sure that a match has occurred when using GED-derived measures. This objection, while of course true, should not significantly affect the analysis because in this case, a match simply refers to the most *likely* point of matching given the GED-based estimate. There is no reason to believe that the estimates should be biased in either direction. Even so, as a sensitivity test, I conducted all eight analyses again using over- and underqualification variables that treat one or two years of mismatch as a perfect match; in essence, this specification allows a grace period before classifying a worker as mismatched. In every case except one, the substantive results for both measurement strategies are identical. For all analyses, I use the model in equation (1) (or a binomial logistic version for the non-monetary outcomes). Controls for age, age squared, the year of the survey, nonwhite ethnicity, never-married status and (for non-monetary outcomes) respondent income are also included.[6] In addition, because it is possible that men and women experience education-occupation matching differently, I conduct separate analyses by gender.

Whether a genuine mismatch effect can be inferred from these models depends on the sign and magnitude of these three parameters, and varies depending on the outcome variable under consideration. In the case of earnings, for example, an overqualification effect would be inferred if the returns to overqualification ("surplus education") were less than the returns to the required education (i.e., $\alpha > \tau$; see Sicherman 1991; Tsang et al. 1991; Cohn and Kahn 1995). This is because the model would predict that a worker with 16 years of education (for example) would earn less as an overqualified worker than in a matched job. On the other hand, if the overqualification and required education effects were identical ($\alpha = \tau$), then this would be the straightforward case of job-independent "returns to education."

The criteria for evaluating mismatch effects are different for different outcomes, depending on theoretical predictions. In the case of job satisfaction and achievement ideology, a mismatch effect can be inferred if the overqualification effect (τ) is negative. This would reflect the fact that educational attainments that are matched by corresponding occupational attainments lead to social consonance, while educational attainments in excess of occupation achievements lead to dissonance and discontent. In the case of liberal political views, a genuine overqualification effect could only be inferred if the positive effect of overqualification (τ) on left political views is significantly greater than the effect of required education (α). This is so because the effect of education on liberalism is already positive – for mismatch to have an effect beyond the primary education effect, there must be an acceleration of liberalism beyond the point of required education.

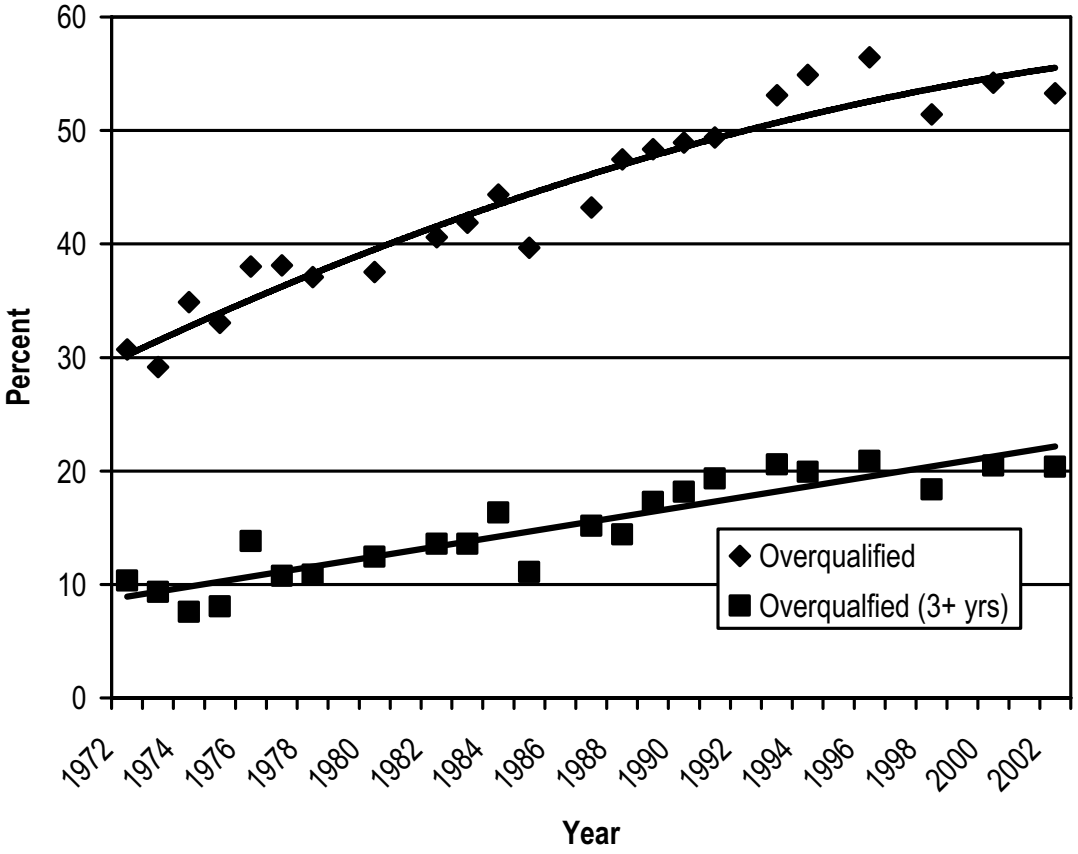
Results

The Incidence of Overqualification, 1972-2002

Figure 2 shows trends in overqualification using both simple (1+ years) and highly overqualified (3+ years) benchmarks. Using the simple benchmark, overqualification has increased from around 30 percent to near 55 percent of the full-time working population. This estimate falls roughly in the middle of those that would be generated by other conversion schemes (see Livingstone 1998). The more conservative benchmark shows a trend from roughly 10 percent in 1972 to around 20 percent in 2002. Both trends appear to be approximately linear. This visual impression was confirmed with BIC comparisons between binomial logistic models using linear and non-linear specifications of the survey year as predictors of overqualification. Contrasting linear, quadratic, cubic and all possible functional forms specifications of the model, the linear model comes out as the best for estimating the yearly trend for the percentage of the workforce with three or more years of overqualification. For the measure using the simple benchmark, a quadratic model offered the best fit, and showed a leveling off in recent years.

If the measure used here is capturing overqualification as opposed to random noise or a measurement artifact, it should behave according to plausible predictions. While data limits the ability to test such predictions, it is

Figure 2. Overqualification 1972-2002, Full-time Workers Aged 25-65



reasonable to infer that the state of the economy should have an impact on the incidence of overqualification. During hard economic times (i.e., when loose labor markets prevail), workers may be more willing to accept a job whether or not it fits perfectly with their level of education. One would therefore expect that the unemployment rate should have a positive effect on the incidence of overqualification. When lagged one year, the unemployment rate is indeed a significant positive predictor of overqualification (highly overqualified only; $p = .003$). The linear year measure is still highly significant net of unemployment, however ($p < .000$), and comparing the range of predicted probabilities for the year and unemployment rate demonstrates that the effect of the linear trend is approximately four times greater than the effect of the unemployment cycle. Thus, while economic cycles appear to play a role in promoting mismatch, they cannot account for the large, monotonic rise in its incidence. This result also suggests that the "highly overqualified" benchmark is probably a more realistic measure of overqualification than the simple benchmark, since only the former behaved according to predictions.

By examining the trends in both educational attainment and educational requirements over the same period, the reason for the increase becomes clear – the level of education of the average full time U.S. worker has increased by 1.75 years from 1972-2002, while the educational requirements of the average full-time job have increased by only .33 years.[7] This latter figure represents not only skill upgrading and downgrading within occupations, but also reflects changes in the

Table 2: Overqualification Among Full-time U.S. Workers by Selected Characteristics, 1972-2002

	Percent Overqualified (3 or more years)		
	1972-82	1983-92	1993-2002
Gender			
Men	11.6	16.1	20.3
Women	9.4	14.5	19.8
Age			
25-34	13.9	16.1	20.8
35-44	11.0	17.8	19.0
45-54	9.2	13.3	20.1
55-65	6.6	10.1	18.9
Race			
White	10.4	15.0	19.4
Non-white	14.0	17.8	22.6
Education			
0-11	.4	.6	1.3
12	5.5	4.7	6.0
13-15	12.5	15.0	17.1
16	17.7	25.7	29.8
17-20	39.7	45.3	51.0

kinds of jobs workers hold. For example, the proportion of workers in technical and managerial work has increased, while the proportion in manufacturing has decreased. Changes in the mix of jobs held by workers account for a little more than half of the increase, while skill upgrading within occupations accounts for the rest. Thus, while there has been some upgrading of educational requirements, these increases have been exceeded by the expansion of educational attainments. Of course we cannot be sure about exactly what has actually happened on the demand side (especially since 1991), so these are rough estimates at best. It is noteworthy however, that Department of Labor experts saw little change in requirements between 1977 and 1991 – a period of significant computerization in the workplace (Autor, Levy and Murnane 2003). It may therefore be the case that the massive "upskilling" of jobs so often talked about in the media is less common than imagined (Rothstein 1999).

In order to see more clearly the social patterning of overqualification, a number of descriptive statistics are included in Table 2: overqualification by

gender, race, age and educational level. The statistics in this table are based on the highly overqualified benchmark. Table 2 demonstrates that overqualification has increased over time in all categories. Men and women are about equally affected, with men slightly more likely to be overqualified. Non-whites are still significantly more likely to be overqualified than whites, despite their lower average level of education. These findings are not surprising given the numerous mechanisms that keep minorities from obtaining high quality work (Tilly 1998). Also not surprising is the finding that higher levels of education increase the risk of overqualification.

A significant trend that appears in the table is the decreasing salience of age for overqualification; while the youngest workers are more than twice as likely as the oldest to be overqualified in the earliest period, this gap shrinks to almost nothing in subsequent years. A comparison using simple binary logistic regressions of overqualification on age in all three periods confirms this finding. The age coefficient in the third period is not statistically significant, while the age coefficient in the first period is highly significant ($p = .000$). This pattern undermines the objection that overqualification is simply a life course phenomenon; that is, it takes younger workers time to sort themselves into occupations that fit their credentials and training (Sicherman 1991). While this "bad-job-as-stepping-stone" argument is undoubtedly true as far as it goes (and may once have been truer than it is today), the age trends in Table 2 show that a life course view of occupational matching cannot explain *overall* increases in overqualification. This may reflect the fact that younger cohorts are increasingly facing job markets where the number of qualified workers is greater than the number of commensurate jobs. It may also reflect breakdowns in traditional advancement structures such as the firm-internal labor market. Taken as a whole, what Table 2 demonstrates is that, while there are systematic differences in the risk of overqualification, its absolute incidence has increased among all sociodemographic groups.

The Effects of Overqualification, 1972-2002

Table 3 shows the results of the regression analyses. I first examine the income model to determine whether or not the measures used here produce findings similar to those obtained by Sicherman (1991) and Cohn and Kahn (1995) using self-report measures. Indeed, the analysis reveals that the magnitudes of the coefficients are similar and that the relationships between the three education coefficients are identical (compare Cohn and Kahn 1995:73, Table 2). A test of the equality of the education coefficients shows that the returns to required education are larger than the returns to overqualification ($\alpha > \tau$) for both men and women (F = 119.8 and 160.4, respectively). The same test confirms that the required education effect is also stronger in absolute terms that the underqualification effect ($|\alpha| > |\delta|$) for both men and women (F = 72.3 and 74.4). The strong consistency between these findings and those obtained using different data and measures provides additional evidence that the GED-based measure of required schooling is similar to subjective reports, and that these measures should yield equally valid results in each of the following analyses.

Table 3: Results from the Regression of Educational Variables and Controls on Selected Outcomes, GSS Full-time Workers, 1972-2002

MEN	Income (ln)	Job Satisfaction	Liberal	Achievement
Required education (α)	.116**	.069**	.088**	-.007
	(36.73)	(6.48)	(7.42)	(.50)
Overqualification (τ)	.056**	.001	.127**	-.081**
	(10.51)	(.08)	(6.89)	(3.90)
Underqualification (δ)	-.065**	.072**	-.026	.005
	(12.72)	(4.40)	(1.32)	(0.22)
Age (25 yrs. = 0)	.036**	-.015	.002	-.021*
	(14.18)	(1.88)	(.20)	(2.02)
Age^2	-.001**	.001**	.000	.000
	(9.77)	(3.64)	(1.18)	(1.66)
Year (1972 = 0)	-.005**	-.005	-.020**	.019**
	(5.85)	(1.82)	(6.23)	(5.32)
Non-white	-.158**	-.153*	.546**	-.364**
	(7.41)	(2.25)	(7.54)	(4.41)
Never married	-.210**	-.263**	.413**	-.100
	(10.23)	(4.04)	(5.93)	(1.23)
Income (ln)	—	.203**	-.141**	.114**
		(5.86)	(3.69)	(2.67)
Constant	1.920**	-1.680**	-1.541**	.301
	(39.15)	(9.96)	(8.20)	(1.45)
R^2	.21	—	—	—
χ^2	—	294.95	251.91	70.62
N	8756	8419	8179	5630
WOMEN				
Required education (α)	.144**	.098**	.139**	-.067**
	(36.85)	(7.27)	(9.33)	(3.96)
Overqualification (τ)	.062**	-.039*	.178**	-.074**
	(10.18)	(2.00)	(8.58)	(3.00)
Underqualification (δ)	-.075**	.000	-.036	.084**
	(10.13)	(.01)	(1.27)	(2.66)
Age (25 yrs. = 0)	.030**	.006	-.003	-.007
	(10.61)	(.67)	(.32)	(.62)

Table 3 ***continued***

Age^2	-.001**	.000	.000	.000
	(7.14)	(1.38)	(.96)	(.84)
Year (1972 = 0)	.003**	-.016**	-.001	.021**
	(3.31)	(4.76)	(.20)	(4.98)
Non-white	-.021	-.447**	.238**	-.154
	(1.01)	(6.66)	(3.27)	(1.82)
Never married	.065**	-.246**	.162*	-.210*
	(2.87)	(3.45)	(2.14)	(2.28)
Income (ln)	—	.152**	.020	.029
		(3.97)	(.47)	(.59)
Constant	.902**	-1.537**	-2.871**	1.445**
	(15.68)	(8.36)	(13.81)	(6.13)
R^2	.20	—	—	—
χ^2	—	312.14	227.86	70.92
N	6941	6648	6295	4388

Notes: Absolute value of t or z statistics in parentheses. Education measured in years.
*$p < .05$ **$p < .01$

The model for job satisfaction yields the predicted mismatch effect for both men and women. Unexpectedly, however, the mismatch effect appears as an overqualification effect for women and an underqualification effect for men. For women, each year of non-required education is associated with around a 4 percent decrease in the probability of full satisfaction. Yet, this is the one analysis for which the "grace period" measurement strategy yielded a somewhat different result; using this specification, there were no mismatch effects for women. Thus this finding should be interpreted cautiously. For men, on the other hand, each year *below* the education requirement is linked to a 7 percent *increase* in the probability of full satisfaction. This finding is consistent with the predictions of the "classic" overeducation thesis (i.e., of rising expectations) but is harder to explain theoretically. Perhaps just as failing to meet one's expectations seems to produce a negative reaction, surpassing one's expectations may produce a positive reaction. Perhaps succeeding in an occupation where most other workers are more highly educated can lead to pleasingly high estimates of one's "native" ability. Without different kinds of data, however, this is pure speculation.

The next model, which predicts liberal political attitudes, also shows some evidence of an overqualification effect. Unsurprisingly, both the required education parameter (α) and the overqualification parameter (τ) are positive and significant, confirming that education tends to lead to liberalism whether required by one's job or not. However, a Wald test of the equality of the α and τ coefficients confirms the "acceleration effect" of overqualification, at least

for men ($\chi^2 = 4.43$, $p = .04$). The effect of a non-required year of schooling on the odds of liberal political ideology is approximately 45 percent greater than that of a required year of schooling for male workers. For women, however, the difference between the coefficients falls short of the conventional benchmark for statistical significance ($\chi^2 = 3.08$, $p = .08$). It seems that more educated women are more liberal regardless of their occupation. In any case, the effects here are not massive; while an overeducation effect on political views appears to be supported, there is little evidence of the revolutionary consequences predicted by early theorists.

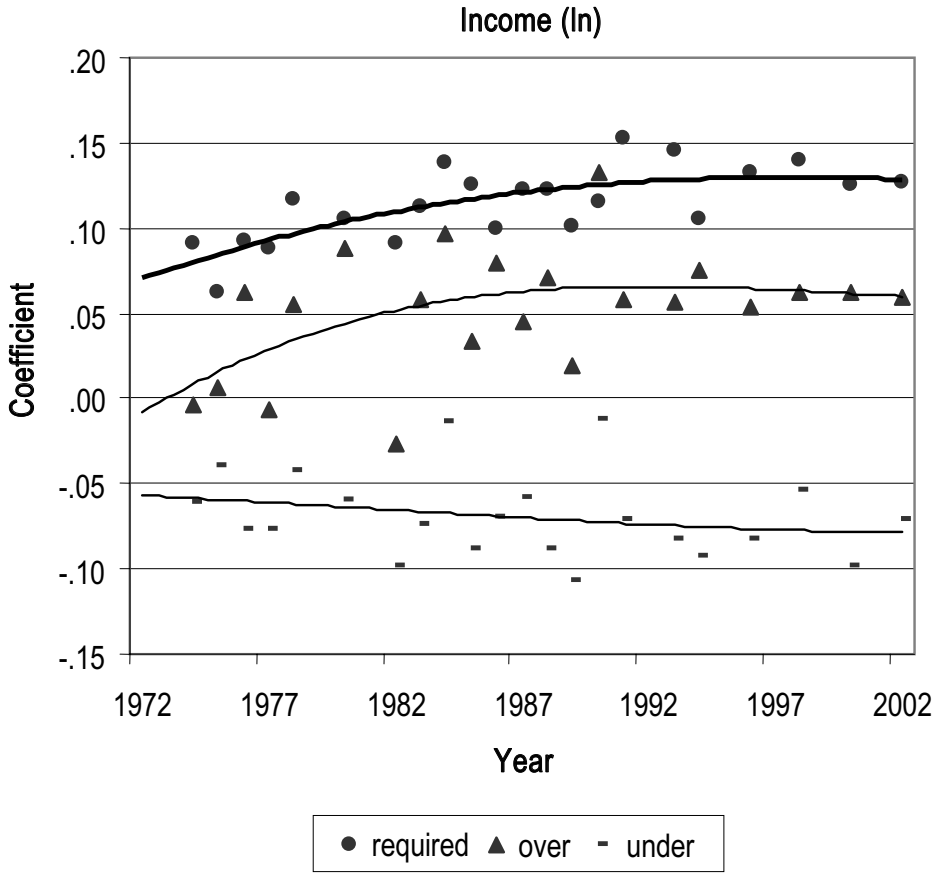
The final model, which predicts the belief that hard work is the key to success, yields differing results by gender. For men, the overqualification parameter (τ) is the only one of the educational parameters to be statistically significant. Each year of overqualification leads to an 8 percent decrease in male workers' probability of endorsing a hard work-only achievement ideology. For female workers, the picture is different – it appears that education, required or not, has a negative impact on this achievement ideology. A Wald test confirms that $\alpha = \tau = -\delta$ in this model ($\chi^2 = .37$, $p = .83$).

Changes in Effects over Time

The models above represent the average effect of mismatch over time, controlling for the effect of a possible yearly trend in the dependent variables. This does not tell us, however, if the *effect* of mismatch has varied over the 1972-2002 period. Knowing whether and how mismatch effects have changed serves two related goals. The first and most obvious is that there is a simple descriptive value in discerning trends in mismatch effects – if such trends exist, they represent an interesting phenomenon to examine and explain. The second, more important, reason to examine these trends is that looking at effects over time can help confirm my measurement strategy. Skeptics may argue that the measures used here do not adequately track changes in educational requirements on the demand side (especially since 1991). I cannot completely refute this objection through data and measurement. However, if it is true that educational requirements have in fact been increasing more quickly than my measures reflect, we should observe an attenuation of overqualification effects over time; this is because the "overqualification" measured here would be decreasingly "real" (i.e., more and more upwardly biased). I therefore argue that evidence of stable mismatch effects over time serve as evidence of the reliability of the measures used in this research and the reality of overqualification as a legitimate social phenomenon.

One way to measure such trends would be to add interactions between the survey year and the education measures (required, over and under). One limitation of this strategy, however, would be that only linear changes in effects could be modeled. What is more, entering three interactions with the survey year would introduce unacceptably high levels of collinearity and difficulties in interpretation. As an alternative strategy, I conduct separate regressions for each year of the survey (using all controls from Table 3) and create a new dataset containing the education parameters. I graph the results by year and use both visual and statistical means to spot trends. This approach has been used

Figure 3a. Trends in Educational Effects for Male Workers, 1974-2002



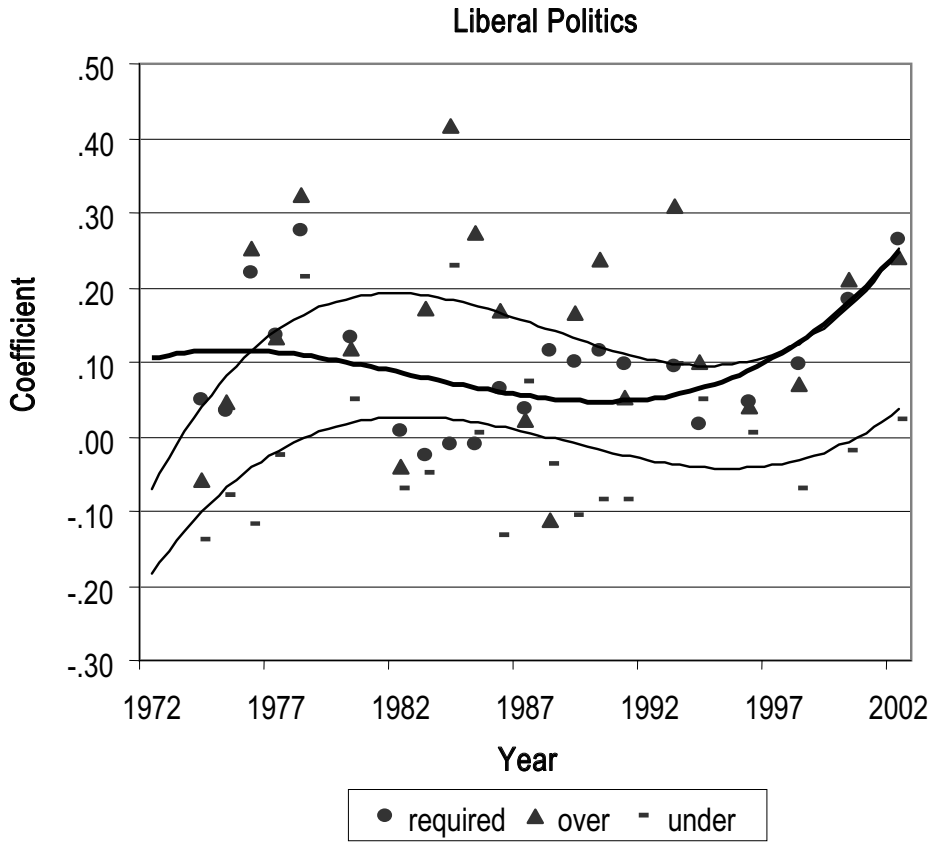
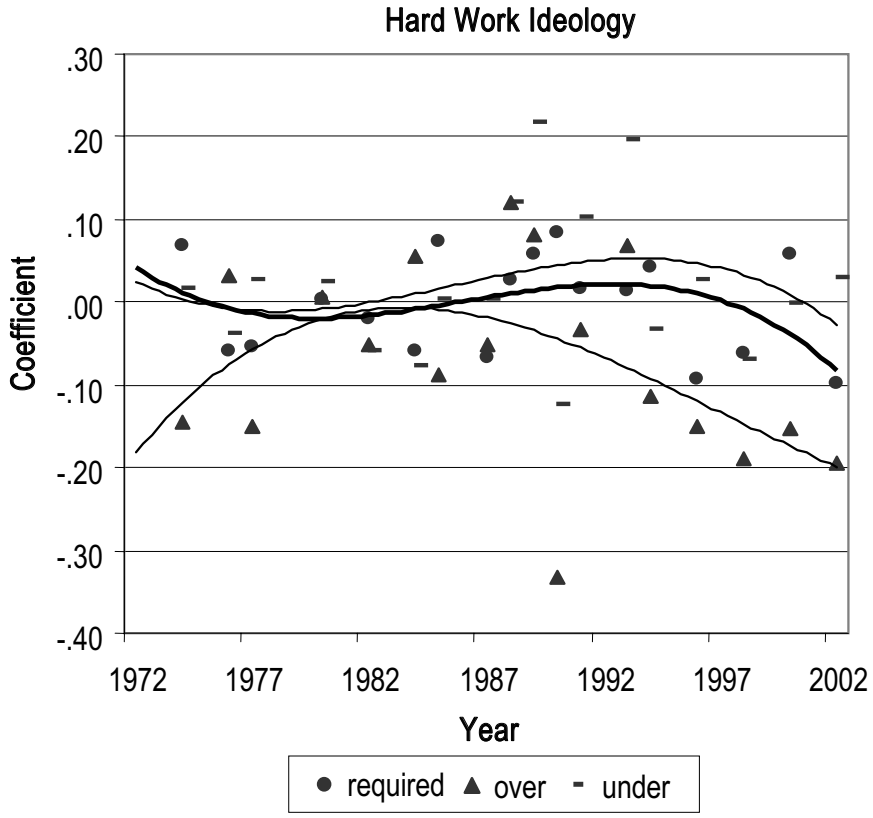


Figure 3a. ***continued***



profitably in previous research (e.g., DiMaggio, Evans and Bryson 1996). I present these graphs separately for men (Figure 3a) and women (Figure 3b). To facilitate discerning changes, I use a third order polynomial to draw trend lines. To test whether there are meaningful trends, I conduct regression analyses, modeling trends in the estimated coefficients with both linear and quadratic specifications. While significance tests of these trends are not strictly applicable, they are used here for heuristic value (see DiMaggio, Evans and Bryson 1996).

Figure 3a shows that, despite some yearly volatility, there are few systematic changes in the value of the three educational coefficients or the relationship between these coefficients over time for male workers. There are a few exceptions that are noteworthy. First, the income premiums for required education and overqualification increased in parallel from the earliest time point, before leveling off around 1990. Over the same period, the underqualification penalty increased as well. These polar trends most likely reflect the increasing size of the college/non-college wage gap (Smith 1986). Again, the fact that the required education and overqualification coefficients are increasing in parallel implies that the measure of required schooling used is fairly accurate. If I were dramatically underestimating the increase in educational requirements over time (because of computerization or other factors), one would expect a gradual convergence between the overqualification coefficient and the required education coefficient. Again, this is because an increasing amount of "overqualification" would have become *required education* over this period. The results are consistent with the contention that these measures are sufficiently accurate and unbiased.

There are two other significant trends in Table 3a. The effect of required education on job satisfaction has followed a U-shaped path from 1974-2002. Why this should be the case is unclear. The relationship between required education and liberal political views has followed the same U-shaped trajectory; there has been a strong increase in the educational effect on liberalism in recent years. In fact, it appears that the required education effect has caught up with the overqualification effect. This means that the acceleration effect of overqualification – though true on average between 1972 and 2002 – has largely been absorbed by the effect of education in general. Apart from this change, however, there is no evidence that the mismatch effects reported in Table 3 have systematically decreased over time. In fact, in the case of hard work ideology, the overeducation effect is actually *increasing* in strength. That is, overqualified workers are increasingly likely to disavow the "American Dream" of success through hard work alone.

Figure 3b presents the same analyses for female workers. In contrast to the models for male workers, the coefficients here seem to be much more stable over time, with both fewer trends and fewer fluctuations. In fact, the only significant trend in any of the four models is the increasing income penalty for underqualification. Again, this probably reflects the growth of the college/non-college wage gap between the early 1970s and the early 1990s. Unlike for men, however, this has manifested itself only as a decrease for underqualified female workers rather than the polarization seen for male workers. In sum, Figures 3a and 3b provide some evidence that the effects of mismatch are relatively stable over time, with no clear strengthening or weakening trends across the board.

Figure 3b. Trends in Educational Effects for Female Workers, 1974-2002

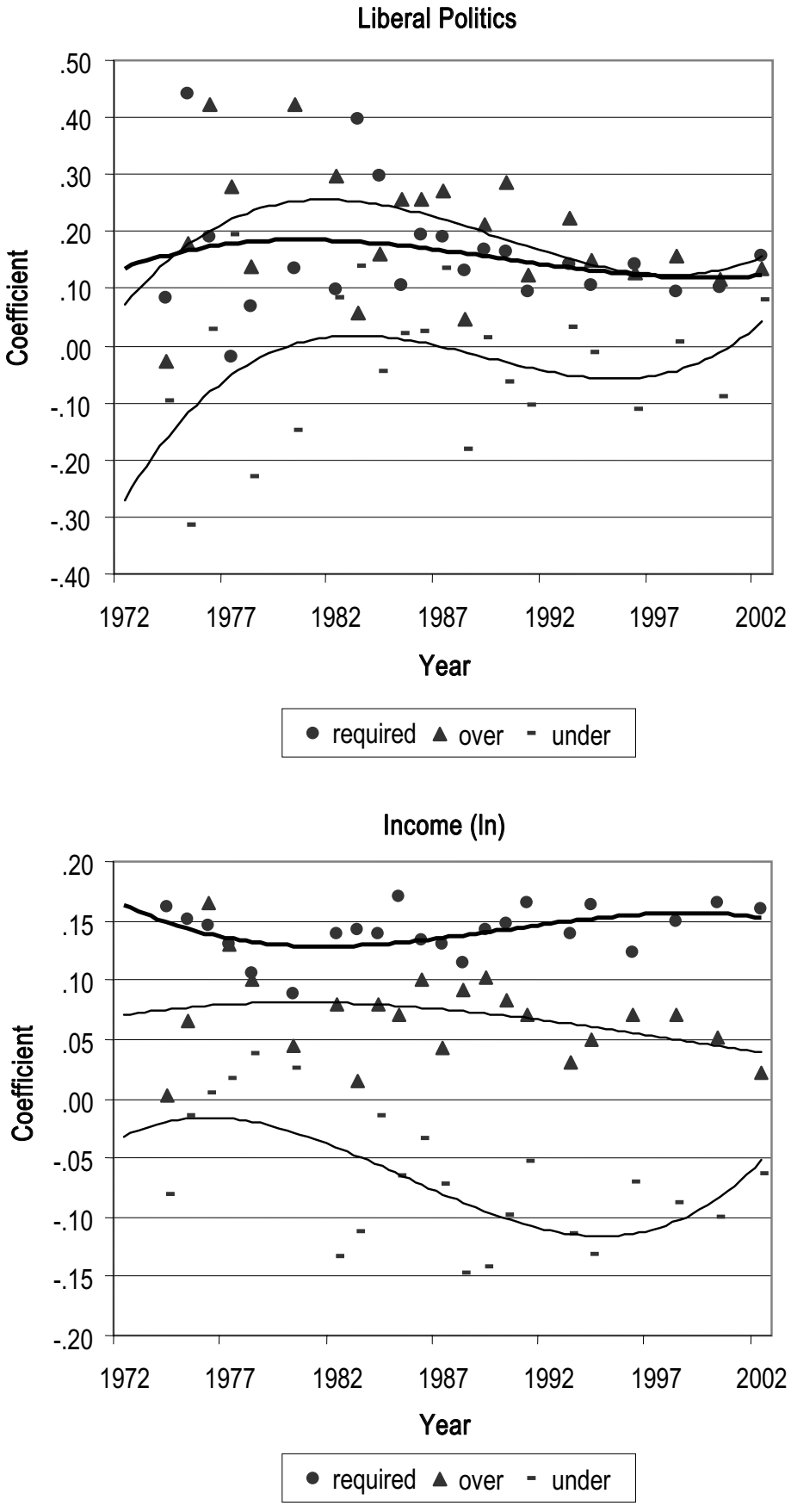
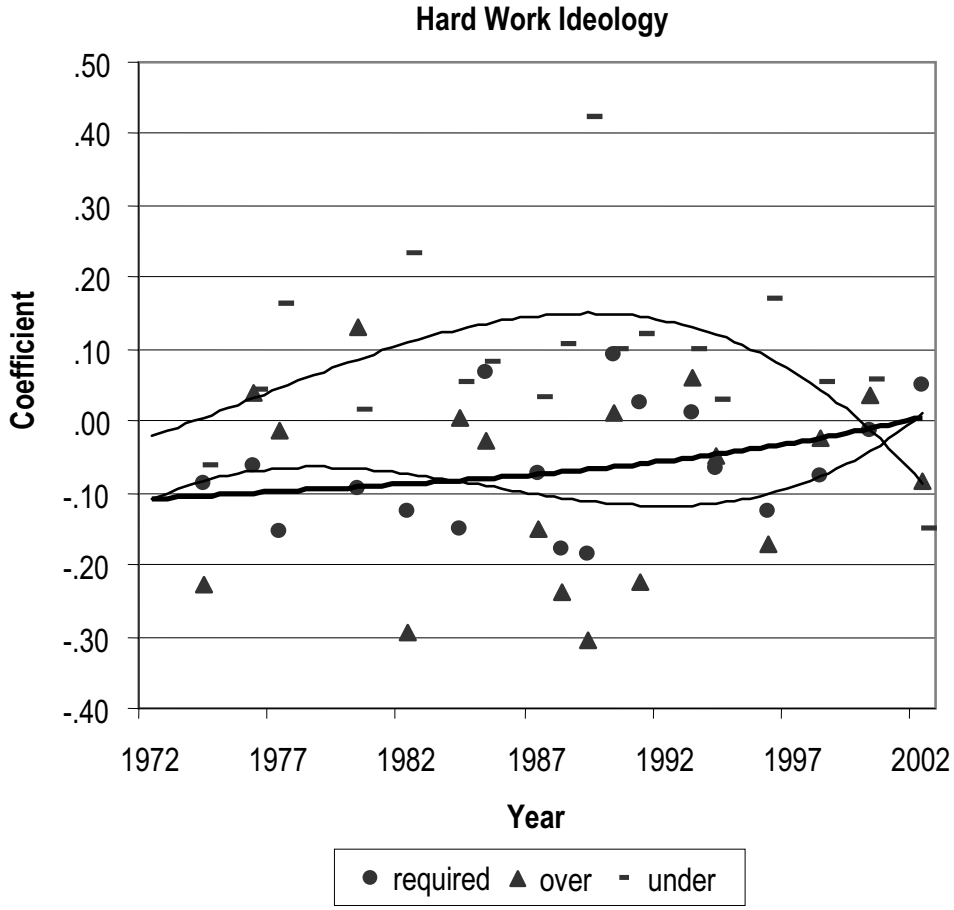


Figure 3b. ***continued***



The results of these trend analyses lend credence to the measure of required education employed here. If the measure were systematically underestimating recent increases in educational requirements on the demand side, the results over time would be less stable than they are. The GED-based measure undoubtedly contains substantial error, if for no other reason than it was constructed at the level of three-digit census occupational codes, which are somewhat heterogeneous with respect to actual jobs. Yet the trend analyses, coupled with the steps taken in the previous sections, strongly suggest that the measure is capturing real overqualification – a phenomenon with stable and significant effects.

Discussion and Conclusion

Within the limitations imposed by the data, the analyses undertaken here yield several significant findings. First, there is evidence that a substantial – and growing – number of American workers are overqualified for their jobs. Though economic cycles do have some impact on this phenomenon, the principal time-trend is positive and linear, and appears to be the result of the widening gap between a large expansion in educational attainment and only modest increases in job educational requirements over the past three decades. Second, overqualification is not limited to the young nor to any other particular demographic group, though disadvantaged groups do appear to be at higher risk. Third, as predicted by the theoretical mechanism of the habitus and the corresponding drive toward achieving structurally-conditioned expectations, overqualification is linked to psychological, social and political outcomes for full-time workers, although the effects are generally stronger for men than for women. Finally, the regression and trend analyses show no evidence that the effects of overqualification are diminishing, despite its increased prevalence. Within the limitations imposed by the data, these findings paint a picture of a growing phenomenon with substantively significant consequences.

The discovery of gender differences in the mismatch effects seems to provide further support for the theoretical mechanisms I have argued are related to overqualification. Because the structural position of women may diminish the occupational aspirations associated with a given level of education (due, for example, to a greater anticipated likelihood of leaving the labor force, at least for a time), the subjective impact of a social discrepancy may be smaller for women. Indeed, work in social psychology (Thoits 1992) has shown that, on average, work is a less identity-salient role for women than it is for men, reinforcing the plausibility of this explanation. This is not primarily because of individual or natural dispositions, but because of our culture's disproportionate focus on paid work in men's lives. This difference can perhaps explain the fewer discrepancy effects for female as opposed to male workers. (As one reviewer suggested, this fits the general pattern of men's greater rebelliousness in crime, violence, political extremism and overall alienation from society.) It would clearly take more (and different) research to be sure if these mechanisms were indeed responsible for the effects outlined here, but the results are at least consistent with such an explanation.

One interesting question that deserves to be addressed regards changes in the levels of the dependent variables over time. If overqualification is increasingly

common and it affects such outcomes as job satisfaction and political views, how have the aggregate levels of these outcomes changed? In the case of job satisfaction, separate analyses show that the percentage of full time workers who are fully satisfied has declined from a high of near 60 percent in the mid-1970s to around 45 percent in the year 2000 (though this figure rebounded to over 50 percent in 2002). In this case, overqualification is likely a causal factor in this decline. More perplexing are levels of political liberalism, which have remained essentially unchanged, and belief in the "American Dream" of hard work, which has substantially *increased* since 1972. Clearly there must be other mechanisms at work to offset the changes produced by overqualification. These mechanisms are certainly worthy of further research.

At this point, I return to the two objections to my measurement strategy that I mentioned above but did not address. The first objection – purported decreases in levels of school quality and/or student performance – should be considered in light of the recent review by Michael Handel (2003), which explicitly addresses the widespread belief that decreases in school quality and student performance have rendered the U.S. labor force unprepared for today's "high tech" labor market. Handel demonstrates that there is no evidence of systematic decreases in either educational quality or in students' cognitive abilities. For example, reading and math test scores from the U.S. Department of Education's National Assessment of Educational Progress (NAEP) have been almost entirely stable since 1971 to the late 1990s. Further, data from the National Adult Literacy Survey (NALS) provide no evidence for the common stereotype that younger workers are less literate than older workers are. In addition to these empirical refutations, Handel relies on historical analysis to show how the skills shortage paradigm emerged as labor economists' response to growing wage inequality; that is, they chose to focus on individual workers rather than management strategies or patterns of job creation. Little attention has been paid, for example, to dramatic increases in "bad jobs" over the same period of increasing wage inequality (Kalleberg, Reskin and Hudson 2000). In the same vein, Handel (2003) shows that fear about school failure emerged as a "moral panic" in response to the economic threat posed by Japan rather than as a rational response to empirical analysis. In summary, then, the available evidence does not support the argument that decreases in school quality or student performance can explain away observed overqualification.

The second possible objection is that since the latest job skill estimates in these data are from 1991, the measure of required education used here cannot capture the impact of widespread personal computing or other technological advancements in the workplace. This is a valid criticism, as far as it goes. Those who invoke it should remember that computers and other technology had significantly entered the workplace between 1977 and 1991 (Autor et al. 2003), and that the Department of Labor skill estimates only show modest increases in skill requirements during this period (see above). The linear interpolation of the skill trajectories should pick up some of any post-1991 increase in any case. Ultimately, though, this objection would only be fatal to the argument here if increases in skill requirements since 1991 could *completely* make up for the large increase in observed overqualification over the same period. This seems unlikely. Moreover,

any actual upgrading of skills that went undetected by the data would only serve to diminish the effect of overqualification on the outcome variables used in the regression analyses. Yet the supplemental analyses showed that the effects of overqualification are stable over time. Taken as a whole, then, these findings provide a compelling case for the reality of overqualification as a social fact.

What are the implications of overqualification for processes of social stratification? For one thing, if the level of education is increasing faster than the education required to do available jobs, this may have serious consequences for individuals, especially in lower income families or among other disadvantaged groups. While the cost of higher education has increased dramatically over the past 30 years, available funding resources – state and federal grants, family incomes – have not kept pace in real terms (National Center for Public Policy and Higher Education 2002). Though the relative wage returns to post-secondary education remain high, this is in large part due to the *falling* real wages of those without it (Livingstone 1998). Because the same jobs now cost more in educational terms than ever before, the families and individuals least able to pay the costs are the most likely to be squeezed out of the best sectors of the labor market, thus maintaining existing stratification hierarchies (at a higher absolute level). As others have argued, post-secondary education is not necessarily the key to upward mobility; rather, for many, it may be the only effective safeguard against downward mobility (Collins 1979). This explanation of increasing inequality can be viewed as a complement (and/or competitor) to the skills-biased economic change theory that is currently orthodox, which holds that a skills shortage among the disadvantaged is primarily responsible for increasing inequality (Handel 2003). It does seem that further research would be warranted to investigate this claim.

The economic costs of overqualification do not only involve workers, however – they also affect employers. Overqualification implies underemployment – the underutilization of available skills and abilities. Given that an increasing proportion of workers may be underemployed, it is in employers' interests to find new ways to make use of these untapped human resources. While there are material and other limits to work design, employers can certainly play a role in closing the gap between educational attainment and occupational content; in fact, it may be beneficial for them to do so. For example, Tsang and colleagues (1991) argue (citing several other studies) that decreases in job satisfaction brought about by overqualification can have serious negative effects on productivity (see also Petty et al. 1984).

Finally, I do not mean to imply that education does not confer non-workplace benefits or that it is not desirable for its own sake (e.g., as a consumption good). Nor am I arguing that education should be made less available or portrayed as less valuable than it is today. My goal has been to point out that increases in educational attainments have unintended consequences that affect society and the economy in ways that are not often recognized. The limitations imposed by the available data necessarily restrict the precision of these results, but it does seem reasonable to conclude this: that the number of overqualified workers in the United States is substantial, that it has grown over the past 30 years, and that this phenomenon has social, psychological and political consequences that

are both substantively significant and theoretically consistent. Though more work is needed, this article provides a needed step in reviving the study of an underexamined social phenomenon.

Notes

1. This phenomena is also known as "overeducation," "overschooling," "overtraining" and "underemployment." I see the terms as interchangeable, but generally favor "overqualification" as the least normatively charged.

2. These data were generously provided by David Autor.

3. Another possibility would be to use a linear interpolation, but only for the years between 1977 and 1991, assigning the 1977 values to the years 1972-1977 and the 1991 values to the years between 1991-2002. These different measurement strategies did not substantively change the results.

4. Bayesian Information Criterion (BIC) comparisons are an appropriate method for choosing between nested and non-nested models (Raftery 1995).

5. However, since this first part of the trend analysis deals primarily with *aggregate* levels of overqualification rather than with the classification of individual worker, the random measurement error should cancel out, yielding an unbiased point estimate of the level of overqualification in a given year.

6. Descriptive statistics for all variables are given in Appendix 2. Alternative specifications of the models using quadratic forms of the survey year variable were estimated, but they were rejected in nearly all cases by BIC comparisons. In no case were the substantive results affected.

7. Even if we only consider the change between 1977 and 1991 (the years for which specific DOT data exists), the average educational attainment increases by 1.14 years and the average educational requirement increases by only .26 years.

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Appendix 1

Actual Education	Required Education: 6	7	8	9	10	11	12	13	14	15	16	17	18	19	20	Total
⩽6	.0	1.0	2.7	10.9	20.8	27.6	7.8	19.4	6.5	1.0	1.0	.0	.3	.0	1.0	100
	.0	7.0	7.0	5.7	5.0	3.1	1.4	1.4	.6	.3	.1	.0	1.6	.0	.8	1.62
7	.0	2.5	1.9	9.3	22.2	29.0	10.5	13.0	10.5	.6	.6	.0	.0	.0	.0	100
	.0	9.3	2.6	2.7	3.0	1.8	1.0	.5	.5	.1	.0	.0	.0	.0	.0	.89
8	.0	.2	3.1	9.4	13.8	25.6	10.1	25.2	10.1	.9	1.8	.0	.0	.0	.0	100
	.0	2.3	12.2	7.7	5.2	4.5	2.8	2.7	1.4	.4	.3	.0	.0	.0	.0	2.51
9	.0	.7	1.1	8.6	17.1	28.2	11.1	22.5	9.3	.5	.5	.2	.0	.0	.2	100
	.0	7.0	4.4	6.8	6.1	4.8	3.0	2.4	1.2	.2	.1	.6	.0	.0	.3	2.42
10	.0	.5	1.6	7.3	17.4	28.6	10.6	22.2	8.6	1.6	1.0	.7	.0	.0	.0	100
	.0	7.0	8.7	8.1	8.8	6.8	3.9	3.3	1.6	.9	.2	2.4	.0	.0	.0	3.39
11	.1	.7	1.3	7.8	13.2	26.5	12.0	22.3	12.7	1.3	1.8	.1	.0	.3	.0	100
	25.0	11.6	8.7	10.7	8.3	7.8	5.5	4.1	2.9	.9	.6	.6	.0	1.0	.0	4.22
12	.0	.3	.9	4.2	9.2	19.3	11.4	31.2	15.8	4.0	3.1	.2	.0	.1	.2	100
	50.0	41.9	41.7	42.4	41.9	41.4	38.5	41.5	26.7	19.4	6.9	7.9	3.2	3.6	2.3	30.73
13	.0	.3	.3	1.7	5.3	17.1	12.6	32.3	18.8	5.7	5.5	.4	.1	.0	.1	100
	.0	9.3	3.5	4.7	6.6	10.1	11.6	11.8	8.8	7.7	3.3	3.7	1.6	.0	.5	8.45
14	.0	.1	.2	1.4	4.7	12.1	9.7	28.5	24.3	7.8	9.8	.7	.1	.3	.2	100
	.0	4.7	3.5	5.4	8.4	10.3	13.0	15.1	16.3	15.2	8.6	9.8	3.2	3.6	1.3	12.19
15	.1	.0	.4	.7	4.0	10.2	7.4	26.0	22.3	8.5	17.8	.9	.2	.7	.7	100
	25.0	.0	2.6	1.1	2.7	3.2	3.7	5.1	5.5	6.2	5.8	4.3	3.2	3.1	1.6	4.52
16	.0	.0	.2	.7	1.3	4.3	6.4	13.3	26.3	11.4	30.1	2.1	.8	1.4	1.8	100
	.0	.0	4.4	3.2	3.0	4.6	10.6	8.8	22.1	27.9	33.1	34.8	35.5	20.4	12.8	15.25

17	.0	.0	.0	.7	.5	2.8	4.6	8.3	21.2	10.7	41.8	2.0	.8	3.5	3.2	100
	.0	.0	.0	.9	.3	.8	2.1	1.5	4.8	7.0	12.4	9.2	9.7	13.3	6.3	4.11
18	.0	.0	.1	.2	.8	1.6	4.1	6.5	17.0	9.8	48.3	3.5	1.3	3.1	3.7	100
	.0	.0	.9	.4	.6	.5	2.1	1.3	4.5	7.5	16.6	18.3	17.7	13.8	8.3	4.76
19	.0	.0	.0	.0	.5	1.1	2.4	3.5	11.6	7.6	38.4	1.4	.5	7.0	26.0	100
	.0	.0	.0	.0	.2	.2	.5	.3	1.3	2.5	5.6	3.1	3.2	13.3	25.0	2.04
20	.0	.0	.0	.4	.0	1.1	1.1	2.1	11.6	8.8	30.4	1.7	2.5	10.5	29.9	100
	.0	.0	.0	.4	.0	.2	.4	.3	1.9	4.1	6.3	5.5	21.0	28.1	40.9	2.89
Total	.0	.2	.6	3.1	6.7	14.4	9.1	23.1	18.2	6.3	13.9	.9	.3	1.1	2.1	100
	100	100	100	100	100	100	100	100	100	100	100	100	100	100	100	100

Note: figures are percentages; row and column modes shaded

Appendix 2
Descriptive Statistics for All Variables

Variable	N	Mean	SD	Min	Max
Required education	18172	13.24	2.30	6	20
Education	18172	13.49	2.96	0	20
Overqualification (years)	18172	1.07	1.50	0	11
Underqualification (years)	18172	.82	1.49	0	20
Income (ln)	15697	3.44	.79	-.69	5.93
Full job satisfaction	17463	.50	.50	0	1
Liberal	15658	.28	.45	0	1
Hard work ideology	11348	.66	.48	0	1
Female	18172	.43	.50	0	1
Age	18172	41.05	10.51	25	65
Non-white	18172	.16	.36	0	1
Never married	18172	.17	.37	0	1