Laid-Off Workers in a Time of Structural Change

Yolanda K. Kodrzycki

Senior Economist, Federal Reserve Bank of Boston. *The author is grateful to Suzanne Teegarden and the staff of the Massachusetts Industrial Services Program for providing the data used in this study and for encouraging research on displaced workers; to Katharine Bradbury, Lynn Browne, Kathryn Lasch, and Robert Triest for detailed comments on an earlier draft; and to participants in a meeting of the New England Study Group for additional suggestions. Margaret Enis provided excellent research assistance throughout the study, Karen Therien was instrumental in the initial preparation of the data while at the Boston Fed, and Ann Kinsella of MIT was responsible for estimating distance to the new job.*

Labor markets have undergone considerable change in recent years. A rising share of unemployment is accounted for by workers who have been permanently laid off. When laid-off workers search for a new job, they often find that job opportunities are different from when they were last hired. Some observers have asserted that the available openings are increasingly for bad jobs. Regardless of whether this is generally true, some types of workers seem destined to settle for jobs that are worse than what they had before. Manufacturing positions are shrinking, especially in blue-collar occupations, and real wages for workers with little education are declining.

This article examines the experiences of workers laid off in this recent period of structural change. Aside from examining the characteristics of individual workers, the research addresses industry effects. That is, do former manufacturing workers have a tougher readjustment following layoff than workers from growing industries such as services? And within manufacturing, do those formerly employed by defense contractors suffer disproportionately as a result of a lack of experience in nondefense work? The article uses data on displaced workers from Massachusetts who sought government-provided reemployment assistance in the early 1990s. Massachusetts experienced a very sharp recession starting around 1989, combined with noticeable shifts in the importance of different industries.

The article starts with a review of previous research on displaced workers. It then analyzes displaced workers from Massachusetts, first examining their duration of unemployment and then, for those who found work, their new wages and other job attributes. The evidence shows that workers from declining industries suffer especially sharp wage cuts, in large part because they tend to have extensive experience at their previous employer which is not highly valued by their new employer. Former defense workers have the most severe adjustment costs of all, in terms of both above-average difficulty in finding work and in

very severe wage losses. The research also indicates that early sign-up for adjustment services tends to reduce the duration of joblessness. However, laid-off workers who participate in education and training programs do not necessarily find better jobs than those who avail themselves only of more basic forms of assistance such as counseling sessions and information on job postings.

I. The Costs of Structural Change

When an economy undergoes structural change, laid-off workers have difficulty finding new jobs comparable to their previous positions. Because of changes in job requirements and job opportunities, laid-off workers are likely to be offered lower pay by new employers. This has two effects. First, absent any change in workers' willingness to accept pay cuts, workers experience a longer duration of unemployment as they search for the best possible job. Second, reemployed workers are likely to incur a pay cut, as even the most attractive job offered within a reasonable time pays less than their previous job. Presumably, adjustment costs should be most severe for workers laid off from declining industries, since they appear least likely to find a new job that is similar to their previous job.[1]

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An extensive literature has investigated which types of workers tend to experience longer spells of joblessness and sharper pay cuts. A longer work history, weaker education, poorer prospects in one's former occupation or industry, and abnormally high previous pay have all been found to lead to greater earnings reductions. The findings are less conclusive with respect to the duration of joblessness, as search techniques appear to vary considerably among otherwise comparable workers. That is, some workers with poor prospects for reemployment at comparable pay do take longer to find a job. Others quickly lower their expectations or broaden their search, so as to minimize their unemployment spell. The duration of joblessness depends both on employers' willingness to offer jobs to different types of workers and on differences in these workers' job search strategies.

Inapplicability of Previous Work Experience

As a result of changes in the structure of jobs, laid-off workers' experiences are not fully transferable to their new job. This has implications for expected earnings at their new job, for several reasons. The pay of employees with a long work history at their previous employer may have reflected the development of job-specific skills, or simply the returns to seniority.[2] When they switch employers, their wage losses are likely to be particularly large. In addition, older or more experienced workers may be at a disadvantage, regardless of whether they spent many years at a single employer or had experiences at many employers. Employers may perceive them as being less malleable in adjusting to a new work environment or more costly in terms of benefit costs per year of expected service.[3] Previous studies overwhelmingly indicate that displaced workers with longer job tenure, and those who change their line of work, face larger wage reductions (or smaller wage gains) at their new job, holding all other factors constant.[4] These workers also experience a longer duration of unemployment.

The Changing Value of Education and Skills

Another aspect of recent structural change is that formal education is becoming more highly valued in the job market relative to "learning by doing." As employers' demand for educated workers has increased, college graduates have been paid a higher wage premium compared to workers with only a high

[1] These workers may be more negatively affected by structural change than those in other industries, although Browne (1985) correctly cautions against ascribing layoffs solely to structural, as opposed to cyclical, changes.

[2] Long tenure may also be a sign that the employee is particularly well suited to the job.

[3] Studies usually are not able to distinguish the separate effects of age and experience. Lacking information on employees' complete work history, they examine potential work experience, assuming that workers participate in the labor force without interruption upon completion of their schooling.

[4] See, for example, the surveys by Hamermesh (1987) and Fallick (1995).

school education.[5] This trend has implications for displaced workers. One study concluded as follows: "[T]he good jobs being created in growth sectors today are concentrated in white-collar work to a degree that they were not previously. The economy is trading decent blue-collar jobs for a range of service sector ones. It is the education and training gap between the high-wage, less educated displaced and their comparable-wage, growth-sector counterparts that is creating difficulties for job-losers" (Seitchik and Zornitsky 1989, p. 82). Indeed, a number of studies have found smaller wage losses for displaced workers who are highly educated and who have been in white-collar jobs. Findings on the duration of unemployment vary.[6] Highly educated workers have skills that are increasingly demanded by employers and they are thought to be more knowledgeable about how to search for work, both of which would tend to shorten their spell of unemployment. On the other hand, they may have at their disposal greater financial resources, which would enable them to stay out of work longer.

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The Disappearance of "Good Jobs"

Some observers believe that displaced workers face a tough adjustment for a broader reason; they allege that "good jobs" are being eliminated and a greater proportion of new openings are for "bad jobs." It has been well documented that certain industries, including transportation and public utilities and most manufacturing industries, pay higher-than-average wages given the experience, occupation, and education profiles of their work forces. Other industries, most notably trade, pay less than expected.[7] When workers lose "good jobs," they are unlikely to find new opportunities that are as attractive, especially if the "good jobs" have been concentrated in industries that are not growing. Indeed, previous studies have found that displacement from an industry that pays high wages increases the likely duration of unemployment and the pay cut at the new job.[8]

Some Unanswered Questions

In addition to providing new evidence on the issues enumerated above, the current study addresses some questions that have been raised in the literature on worker displacements, but that have received little empirical attention. The available studies confirm that workers from declining industries are more likely to switch to a different type of work, but they leave unanswered some other questions pertaining to industry differences. Little evidence exists on whether workers from declining industries have longer-than-average job tenure or more overall work experience. Nor do previous studies indicate whether potential employers judge former defense workers to be particularly ill-suited for nondefense jobs.[9] Anecdotally, it has been pointed out that prime defense contractors serve one customer, the Pentagon, while nondefense firms must market their products more broadly. Also, defense firms specialize in high-precision products that are manufactured in small quantities. By contrast, the typical nondefense firm is engaged in mass production of goods or services that are not as intricate in terms of their technological content or engineering standards. Accordingly, defense employees may be regarded as less able to identify market trends, satisfy a diverse set of customers, or control production costs.

[5] According to the U.S. Department of Labor, in 1992 male college graduates earned 74 percent more than high school graduates and 133 percent more than high school dropouts. In 1979, these differentials had been only 36 and 70 percent, respectively. For further statistics and discussion, see Kodrzycki (1996).

[6] See, for example, the surveys by Hamermesh (1987) and Fallick (1995).

[7] See, for example, Gittleman and Wolff (1993). High levels of concentration may enable firms in some industries to earn above-average profits which they share with workers. In other cases, well-paying jobs may be the result of a high degree of unionization or high productivity. High pay may also serve to compensate employees for unpleasant or dangerous working conditions, as in construction. Economists are continuing to study the reasons for pay differentials by industry.

[8] See, for example, Kletzer (1991), Jacobson, LaLonde and Sullivan (1993), and Carrington and Zaman (1994).

[9] Kodrzycki (1995) studied adjustment costs for New England defense workers, and compared them with findings for national samples of displaced workers from a variety of industries. New England defense workers appear to have been at a disadvantage, especially judging by their likelihood of replacing their former earnings. The study did not examine whether the defense workers did worse because they had been employed in defense industries, as opposed to other characteristics of workers or of the New England economy.

Previous research has looked at the duration of joblessness without distinguishing between time that workers actively spend searching for work and time spent in education and training programs that prepare them for work. This distinction is important, as workers who are most severely affected by structural change may have the greatest need for augmenting their skills. To the extent they participate in education and training programs that enable them to find better-paying jobs, longer spells of nonemployment may not be undesirable.

Another under-researched topic is worker adjustment costs apart from the duration of joblessness and wage changes. New jobs may be less attractive to the extent they are located far away (and therefore entail greater commuting costs or require a move) or offer less generous benefits.[10]

II. Experiences of Displaced Workers

To examine the experiences of displaced workers, this article uses a large sample of Massachusetts workers who were laid off in the early 1990s. This section describes the sample and the economic backdrop, and then turns to examining worker adjustments to layoffs.

The Dislocated Worker Sample

To respond to the needs of displaced workers, the federal government established assistance programs under the Economic Dislocation and Worker Assistance Act (EDWAA), a 1988 amendment to Title III of the Job Training Partnership Act (JTPA). These programs are available for workers who lose their jobs in mass layoffs or plant closures, as well as others who have been laid off and are unlikely to return to their jobs. States apply for worker assistance grants under the auspices of these federal programs and design services within the guidelines set by the federal government. The data on displaced workers in this study are drawn from the administrative records of the 23 assistance centers that served over 20,000 Massachusetts residents laid off between January 1991 and September 1994. (For further information on the sample, see the appendix.) These centers provide various forms of employment assistance to laid-off workers. In addition to offering basic readjustment services such as counseling and job market information to all users, the centers often fund enrollment in education and training programs on a case-by-case basis.

The data base provides considerable information on demographic and job characteristics for the displaced workers, as well as on the assistance services they used while out of work. For those individuals who found new employment through a center, information is available on the duration of joblessness and the characteristics of the new job. Thus, the data can be used to measure the economic costs of job loss and the influences of factors such as the worker's age, educational background, occupation, pay and length of experience at the previous employer, industry, and reemployment services used, as well as local economic conditions.[11]

The Massachusetts Economy in the Early 1990s

In the early 1990s, the Massachusetts economy experienced a severe reduction in employment and a pronounced shift in its composition away from manufacturing and towards services. Between 1989 and 1992, employment in Massachusetts fell 10 percent (Table 1).[12] The losses were disproportionately concentrated in manufacturing and construction, while

[10] New jobs could also entail psychological costs, if, for example, they are considered by the worker to be less prestigious. More fundamentally, being laid off could lead to health problems that might impair a worker's productivity in a new job.

[11] Most of the studies reviewed in Section I were based on the national biennial survey of dislocated workers. The Massachusetts Industrial Services Program (ISP) data have several advantages and disadvantages relative to the Displaced Worker Survey (DWS). First and most obviously, the ISP data are limited to Massachusetts. However, the ISP sample is much larger, as recent samples for the DWS cover only about 3,800 workers. The ISP data are limited to those dislocated workers who chose to seek government assistance in finding a new job, while the DWS is based on representative samples of households across the nation and therefore is thought to produce representative cross sections of displaced workers. The DWS asks questions of individuals about job experiences over the past several years; this retrospective aspect has been shown to result in errors, especially with respect to recalling information concerning the more distant past. By contrast, most of the ISP data are recorded on a current basis, and therefore are less subject to errors due to individual forgetfulness. A somewhat offsetting disadvantage is that they are maintained for administrative rather than statistical purposes, and therefore may contain more data entry errors. Finally, the ISP data set provides more information on employers than is available in the DWS. (Over 5,000 employers are represented in the ISP data set, but almost half of the sample come from 57 employers that laid off at least 50 workers each. Athough many of large layoffs took place at companies with a national or even international reputation, some of them involved employers with only a local presence—such as community hospitals or municipal governments.)

[12] These statistics are based on annual averages, and therefore may disagree slightly from employment changes based on monthly or quarterly data.

Table 1
Composition of Massachusetts Nonagricultural Employment, 1989 to 1994, and of the Displaced Worker Sample

	Massachusetts Employment (Thousands)			Percent of Massachusetts Employment Level		Percent of Massachusetts Employment Loss, 1989–92	Percent of Massachusetts Employment Gain, 1992–94	Number in Displaced Worker Sample	Percent of Displaced Worker Sample
	1989	1992	1994	1989	1994				
All Industries	3,108.4	2,795.0	2,903.6	100.0	100.0	100.0	100.0	20,624	100.0
Manufacturing	561.1	465.7	447.2	18.1	15.4	30.4	−17.1	10,913	52.9
Durables	372.0	299.6	278.2	12.0	9.6	23.1	−19.7	8,363	40.5
Nondurables	189.1	166.1	168.9	6.1	5.8	7.3	2.6	2,550	12.4
Nonmanufacturing	2,136.9	1,945.6	2,065.2	68.7	71.1	61.0	110.1	7,093	34.4
Construction	126.8	73.7	86.0	4.1	3.0	17.0	11.3	341	1.7
Transportation and Public Utilities	128.3	121.4	127.3	4.1	4.4	2.2	5.4	473	2.3
Trade	740.5	640.4	669.4	23.8	23.1	31.9	26.6	2,398	11.6
Finance, Insurance, and Real Estate	217.3	196.7	206.9	7.0	7.1	6.6	9.5	842	4.1
Services	924.0	913.5	975.6	29.7	33.6	3.4	57.2	3,039	14.7
Government	408.8	382.5	390.0	13.2	13.4	8.4	6.8	1,188	5.8
Other and Not Known[a]								1,431	6.9
Memo:									
Computer Manufacturing	36.5	27.6	25.4	1.2	.9	2.8	−2.0	1,357	6.6
Defense-Related									
Private Industry	168.9	143.4	127.8	5.4	4.4	8.1	−14.4	3,089[b]	15.0
Government	21.8	18.2	16.0	.7	.6	1.1	−2.0	477	2.3

[a]"Other" includes agricultural workers.

[b]Includes 2,960 workers in defense-related manufacturing and 129 in other industries.

Source: Except for the memo items, Massachusetts employment from New England Economic Indicators machine readable data; computer manufacturing from New England Economic Project machine readable data; defense-related from Kodrzycki (1995). Displaced worker sample are author's calculations.

job losses in the services industry were minimal. The recovery was uneven. Durables manufacturing firms continued to shed jobs. The services industry, which accounts for about one-third of all jobs, was responsible for over one-half the job growth between 1992 and 1994. Construction employment also rebounded, although the number of construction jobs remained far below the pre-recession level.

Reflecting the extensive job losses in manufacturing, about half of the displaced worker sample consists of former manufacturing workers. The bottom part of Table 1 shows employment in two prominent durables manufacturing industries in the state. Computer industry employment shrank sharply in the early 1990s as demand fell for minicomputers and other products in which Massachusetts manufacturers had specialized. Over 1,300 displaced workers from this industry are included in the sample. Many defense-related jobs disappeared as a result of decreases in the federal defense budget; about 3,000 former employees of prime defense contractors in manufacturing industries are in the sample.[13] The computer and defense industries are of interest not only because of the magnitude of their job cuts, but also because they had been a source of many well-paying jobs. Furthermore, the firms in these industries have been among the state's largest employers and have contributed significantly to the state's reputation as a center for high technology. As a further indication of defense downsizing, the government category includes close to 500 laid-off civilians, mostly from the Fort Devens army base.

[13] Undoubtedly additional laid-off workers in the sample had defense-related private-sector jobs at manufacturing subcontractors or outside the manufacturing sector.

Table 2
Employment Status of Displaced Workers after 12 and 24 Months, and Length of Unemployment Spells
Percent

Year of Layoff and Number of Months	Total Reemployed	Of which: Recalled	Of which: Found New Job	No Longer Enrolled	Not Employed	Average Unemployment Spell for Workers Who Found a New Job (months)[a]	Sample Size
1991							
12 months	14.8	.9	14.0	4.4	80.8		
24 months	53.0	2.0	50.9	29.7	17.4	15.1	3,741
1992							
12 months	45.4	4.0	41.4	15.6	39.0		
24 months	64.8	4.4	60.4	27.7	7.5	8.7	7,394
1993[b]							
12 months	46.1	4.2	42.0	16.1	37.8	6.8	5,897
1994[c]							
12 months	21.7	5.3	16.4	6.7	71.6	4.1	1,332

[a]Calculated for all workers who found a job, regardless of the length of job search.
[b]Includes some workers laid off less than 12 months prior to the end of the sample.
[c]All workers laid off less than 12 months prior to the end of the sample.
Source: Author's calculations based on a sample of displaced workers in Massachusetts.

Jobless Spells

One important measure of the cost of adjusting to a layoff is the length of time workers remain without a job. The displaced worker sample includes people who were laid off between January 1991 and September 1994, but the complete jobless spell is known only for those who found a job by the end of this period. A valuable alternative measure is the percentage of workers who have found work as of a given number of months after being laid off, as opposed to those who are still without work.

Reemployment rates vary, depending on when a worker was laid off. Those losing their jobs later in the period, when the Massachusetts economy was recovering, had more success in their first year of job search than those who were laid off while the state was still in recession. Of those laid off in 1991, only about 15 percent found a job through a worker assistance center within 12 months. By contrast, 45 percent of those laid off in 1992 and 46 percent of those laid off in 1993 had found a job within a year (Table 2).[14]

The probability of reemployment also rises as more time passes since the date of layoff. Consider those laid off in 1992. In the second year after layoff, the reemployment rate rose from 45 percent to 65 percent, as another 19 percent of the sample found new jobs and a few more were recalled. The percentage looking for a job or preparing for a new job through enrollment in vocational or general education classes fell from 39 percent to less than 8 percent.[15]

The industry results show considerable variation (Table 3). Combining the 12-month results across all four years, workers previously employed in the transportation and public utility industry, by defense contractors, and by government had the lowest rates of reemployment. The reemployment percentage for the former defense manufacturing workers would look substantially worse were it not for the relatively high

[14] The 1991 reemployment rate is understated as a result of sample selection bias. The sample includes workers in the assistance center data base as of July 1992. Thus, the data do not reflect the experiences of workers laid off in 1991 who found jobs promptly. (This sample selection bias also results in an understatement of the percentage who are "no longer enrolled.") But the 12-month results for those laid off in the second half of 1991 (which should be less subject to sample selection bias than those from the first half of the year), as well as the 24-month results in general, nevertheless suggest that the recession had a negative effect on job-finding rates for laid-off workers. The 1993 reemployment rate also is somewhat understated, since workers laid off after September were not observed for the full 12 months.

[15] The remaining workers, almost 28 percent, had stopped using the services of the center. Relatively little is known about this last group. Some of them may have found a job on their own, while others may have remained unemployed or were no longer actively looking for a job.

Table 3
Percent of Displaced Workers Reemployed after 12 Months, by Former Industry[a]

Former Industry	Total Reemployed	Recalled to Old Job	Found New Job	Sample Size
All Industries	37.7	3.5	34.2	18,364
Manufacturing	37.7	4.4	33.3	9,778
Defense-Related	34.7	7.6	27.1	2,879
Computers	39.2	2.3	36.9	1,294
Other	38.9	3.2	35.7	5,605
Nonmanufacturing	39.0	2.8	36.1	6,512
Construction	40.7	3.0	37.7	337
Transportation and Public Utilities	30.2	2.4	27.8	417
Trade	39.5	3.7	35.8	2,244
Finance, Insurance, and Real Estate	33.7	1.0	32.8	818
Services	41.2	2.7	38.5	2,629
Government	28.9	3.1	25.7	766
Defense	26.6	1.3	25.3	79
Nondefense	29.1	3.3	25.7	688
Other and Not Known	36.6	1.0	35.6	1,309

[a]Excludes workers whose layoff date or termination date at the assistance center was not known. Includes some workers laid off less than 12 months before the end of the sample.
Soruce: Author's calculations based on sample of displaced workers in Massachusetts.

recall rate for this industry. At the other extreme, laid-off services and construction workers had the highest rates of reemployment. These patterns are fairly consistent with the hypothesis that workers laid off from declining industries have greater difficulty finding a new job than those laid off from growing industries. But some puzzles remain. For example, computer manufacturing employment fell precipitously during the early 1990s, but computer manufacturing workers did not appear to have unusual difficulty finding new jobs.

Probability of Reemployment

To further explore the role of the former industry versus other factors that make reemployment more or less difficult, regressions were run to explain the likelihood of finding work (Table 4). These regressions take into account information on employment outcomes at all intervals, instead of looking at progress after an arbitrary number of months, as did Tables 2 and 3.

The Cox proportional hazards model was used to estimate the coefficients, which represent the relative likelihood of finding employment in any given time period, for a unit increase in the value of the explanatory variable.[16] (Similarly, for dummy variables, the coefficients represent the difference in the likelihood of finding employment when the variable equals one rather than zero.) The effect of a two-unit change in the value of an explanatory variable is obtained by taking the square of the estimated coefficient, and, similarly, the effects of larger changes are measured through exponentiation to the appropriate power.

The two regressions differ only in the measurement of the duration of participation in education and training programs. In the first regression, the actual duration is used, but the number of observations is reduced because duration could not be calculated for over half the sample. The second regression uses estimated duration for these workers, based on an auxiliary regression (Appendix Table 2). The coefficients indicate that spending an additional month in an education or training program lowers the probability of reemployment to about 90 percent of what it would

[16] For each individual in the sample, define h(t) as the probability of becoming reemployed in month t divided by the probability of becoming reemployed after time t. In the Cox model, $h(t) = h_0(t)e^{b_1x_1+\ldots+b_nx_n}$, where $x_1 \ldots x_n$ are the explanatory variables and $h_0(t)$ is the so-called baseline hazard function—that is, the value of h(t) if all the explanatory variables equal zero. The parameters $b_1 \ldots b_n$ are estimated using maximum likelihood. Note that the change in the relative likelihood of becoming reemployed if the value of variable x_k changes by one unit equals $(e^{b_1x_1+\ldots+b_k(x_k+1)+\ldots+b_nx_n})/(e^{b_1x_1+\ldots+b_kx_k+\ldots b_nx_n}) = e^{b_k}$. These are the values reported in the table under the heading "Hazard Ratio." The model takes into account time censoring—that is, some workers sever their relationship with the assistance center prior to taking a job, while in other cases, the sample period ends before an employment outcome is observed. The Cox technique is efficient, in that the lack of employment for such workers during the time period in which they were observed is taken into account in estimating parameters. The term "hazard" reflects the original use of the Cox technique to analyze the probability of an undesired outcome; however, the technique is equally applicable to analyzing the probability of a desired outcome.

Table 4

Reemployment Probabilities—Estimates Using Cox Proportional Hazards Model

Independent Variable	(1) Hazard Ratio	(1) Standard Error	(2) Hazard Ratio	(2) Standard Error
Experience				
Potential Work Experience Squared	.99978**	.00004	.99971**	.00002
Job Tenure	.99389*	.00243	.99668*	.00153
Schooling (Omitted = Less than High School)				
High School	.89*	.04	.95	.03
Some College	.80**	.04	.93*	.04
College Degree	.72**	.05	.81**	.04
More than College	.55**	.08	.76**	.06
Previous Occupation (Omitted = Services)				
Professional, Technical and Managerial	1.06	.09	1.04	.06
Clerical and Sales	1.12	.09	1.03	.06
Production	1.15	.09	1.10*	.06
Other	1.04	.12	1.01	.08
Not Known	1.36**	.12	1.25**	.07
Previous Industry				
12-Month Employment Growth Rate	1.02**	.005	1.01**	.003
Wage Premium	1.0002	.003	.9951*	.0019
County Unemployment Rate at Time of Layoff				
Level	.98	.01	1.01	.01
12-Month Change	.94	.04	.96	.02
Difference between Statewide and County Unemployment Rate at Time of Layoff				
Level	.97	.02	1.00	.01
12-Month Change	.95	.05	.91**	.03
Pre July 1992 Dummy	.94**	.005	.94**	.003
Demographic Characteristics				
Gender and Marital Status (Omitted = Unmarried Male)				
Married Male	1.37**	.07	1.23**	.04
Married Female	1.20**	.06	1.10**	.04
Unmarried Female	1.10	.05	1.07*	.03
Nonwhite	.85**	.04	.84**	.03
Number of Dependents	.97*	.01	1.01	.01
Recalled	2.73**	.24	4.11**	.16
Enrolled Prior to Layoff	1.70**	.14	1.58**	.07
Duration of Employment and Training	.90**	.004	.92**	.003
Proxied Duration?	No		Yes	
Pseudo R^2	.023		.018	
Number of Observations	7,122		16,723	

*Significantly different from one at 5 percent level.
**Significantly different from one at 1 percent level.

be otherwise, presumably because these workers are less likely to be actively searching for a job. In most other findings, the two regressions also are quite similar.

Added years of work experience, either generally or in the previous job, lower the probability of reemployment. The coefficients are close to one, which means that small increments in overall work experience or tenure at the previous job have little impact. But extensive experience can be a substantial impediment to finding a job. The estimates imply that having 30 years' general work experience lowers the probability of reemployment by about one-third.[17]

Workers with less education were more likely to be reemployed. Controlling for education, previous occupation was not a reliable predictor of reemployment, except that former production workers were more

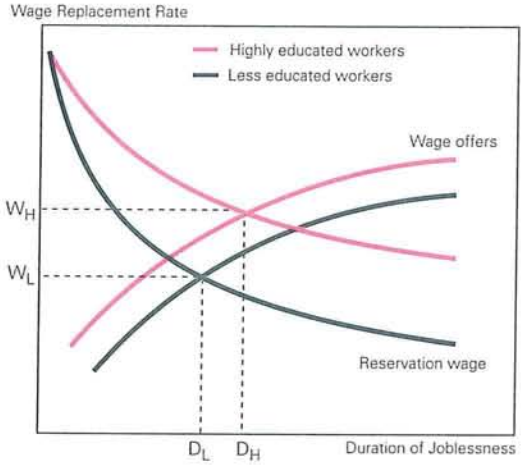
[17] Generalizing the formula in the previous footnote, the effect of an additional n years of experience for someone who has E years of experience is computed as $[e^{-.00022(E+n)**2}]/[e^{-.00022E**2}] = e^{-.00022E(E+2n)}$, where −.00022 is the natural logarithm of the hazard ratio, .99978. This formula is most appropriate for small values of n; that is, the effect of the last year of experience should be rather similar to the effect of the next year of experience. Nevertheless, calculations with larger values of n are suggestive. At the margin, 10 years' experience decreases the probability of reemployment by 6 percent for someone with 10 years' experience, 16 percent for someone with 20 years' experience, and 28 percent for someone with 30 years' experience. Thirty years' experience for someone with E = 30 is calculated to decrease the chance of reemployment by 33 percent. Another version, not shown, estimated the Cox model using age categories instead of experience. That version indicated that the probability of reemployment falls with age, and the results for older workers are fairly consistent with the version with experience (assuming that age and experience are closely related). Relative to the reference category of workers under 25 years of age, the probabilities were .77 for 25- to 44-year-olds, .73 for 45- to 55-year-olds, and .54 for those 55 years old and above.

Job Search for Workers with Different Levels of Education

Figure 1 depicts a simplified economic model of job search. The wage offer curve indicates that displaced workers most readily find jobs with low wages. (The wage replacement rate, measured on the vertical axis, refers to the new wage in relation to the worker's previous wage.) As they extend their job search, however, they are offered jobs with successively higher wages. The reservation wage is the lowest wage workers are willing to accept. At first, workers are not willing to accept much, if any, reduction in their wage. But their reservation wage falls as they become more familiar with job opportunities and as their financial situation deteriorates. The intersection of these two curves indicates how long they will search for a job and the wage replacement rate at their new job. The evidence in this study points to two differences between highly educated and less educated workers. First, as a result of changing labor market demands for skill, less educated workers typically generate inferior wage offers relative to their old pay than highly educated workers. Second, less educated workers are less patient in their search. That is, their reservation wage falls relatively quickly with time.

Figure 1

Simplified Model of Job Search



likely to be reemployed (compared to those in the omitted occupation, service workers).[18] Regressions presented below indicate that less educated and production workers on average accepted job offers with low pay relative to their old job; more highly educated workers searched longer, but their patience paid off in terms of obtaining jobs with a smaller reduction (or greater increase) in earnings. Thus the reemployment results appear to reflect the greater willingness of less educated and production workers to accept job offers, as opposed to the alternative explanation that employers have greater demand for these types of workers. (See the box.)

Workers laid off from faster-growing industries were quicker to find employment. Table 5 shows the growth rate in each industry in the 12 months subsequent to when workers were laid off.[19] According to the estimates in Table 4, all else equal, workers from the construction industry, which was growing at a 5 percent rate, were 7 percent more likely to be reemployed than the average laid-off worker. Services industry workers' likelihood of reemployment were about 5 percent above average.[20]

Industry wage premia varied considerably and appear to have had a modest effect on the probability of reemployment. The premia shown in Table 5 were derived from

[18] Examples of service workers include cooks and chefs, caterers, medical attendants, security guards, police officers, janitors, waiters, hairdressers, and amusement and recreation service occupations. Production workers include occupations such as mechanics, machinists, metalworkers, food processors, welders, benchwork occupations, construction occupations, tailors, and landscapers. Production workers may be employed in service-producing industries (such as a mechanic in an auto repair shop), and service workers may be employed in goods-producing industries (such as a security guard at a factory). The professional, technical, and managerial category includes general managerial and administrative occupations, as well as specialized positions such as drafters, educators, librarians, computer programmers, scientists, and health care professionals. Bookkeepers, secretaries, sales and stock clerks, cashiers, tellers, and billing clerks are examples of sales and clerical workers. The "other" category includes truck drivers and materials handlers; graphics specialists included under this occupational category were reassigned to the professional, technical, and managerial category.

[19] That is, the growth rate depends on both the industry and the date of layoff. The value shown by industry is the average for all workers laid off from that industry.

[20] These findings refer to the second regression; the first regression indicates somewhat larger effects.

Table 5
12-Month Employment Growth and Wage Premium, by Industry
Percent

Industry	Employment Growth[a]	Wage Premium[b]
Manufacturing		
Defense-Related	−3.2	12.4
Computers	−14.1	8.8
Other	.2	−2.4
Construction	5.1	12.7
Transportation and Public Utilities	1.4	5.7
Trade	1.3	−8.0
Finance, Insurance and Real Estate	1.7	−.0
Services	3.2	−1.9
Government		
Defense	−2.3	0
Nondefense	−.1	0
Other and Not Known	1.1	−.1

[a]Based on the 12 months following layoff. Sample average is −0.05 percent.

[b]Relative to government. Sample average is 0.7 percent.

Source: Author's calculations based on machine readable data from the New England Economic Project, the sample of displaced workers from Massachusetts, and Appendix Table 3.

a regression of the worker's previous wage as a function of individual qualifications and industry dummies (Appendix Table 3). Defense-related manufacturing and construction paid over 20 percent above trade after adjusting for the education, occupation, and experience profiles of their workers and, as a consequence, their reemployment probabilities were 8 percent lower.[21]

Local economic conditions influenced employment somewhat. For the period July 1992 onward, a worker laid off in a county where the unemployment rate rose by 1 percentage point in the 12-month period following layoff was about 5 percent less likely to be reemployed after a given number of months than a worker laid off in a county where the unemployment rate was stable.[22] Moreover, an increase in the statewide unemployment rate in the 12-month period following layoff or increases in either the county or state unemployment rate in the 12-month period following layoff also reduced the chances of finding work. The experiences for those laid off before July 1992 are captured by a dummy variable. Workers laid off relatively early, when total employment in Massachusetts still had comparatively far to fall, had notably lower success in finding reemployment. The coefficient suggests, for example, a one-third lower reemployment likelihood for those laid off at the beginning of 1992 than six months later. However, the interpretation of this result is somewhat unclear (especially for the very early layoffs), as the dummy variable picks up the effects of both economic conditions and sample selection bias.[23]

Married and white workers had higher reemployment probabilities than unmarried and nonwhite workers, after adjusting for other qualifications. Those recalled to their old job were reemployed much more quickly than those who looked for a new job.

Finally, those who registered for services prior to being laid off were far less likely to remain without a job for an extended period of time. This last result suggests that employers can mitigate adjustment costs for laid-off workers. By announcing layoff plans in advance, they can help assure that state reemployment services are in place promptly in the local area.[24] Workers then have the opportunity, if they choose, to receive counselling about new job opportunities and to develop search strategies prior to losing their old job.

Explaining Reemployment Differences by Industry

To explore the reasons behind industry patterns further, the estimated coefficients from the Cox regression were used to calculate the reemployment probability for a worker with the average characteristics (job tenure, experience, and so forth) for selected industries, relative to a worker with the average characteristics for the sample as a whole. Then the calculation was redone, assuming sequentially that in each respect the "average" worker in the industry was the same as the "average" worker for the sample.[25] The

[21] This result refers to the second regression shown in Table 4; the coefficient in the first regression indicates an unexpected, though very small, association between the wage premium and reemployment.

[22] There is little correlation between the number of sample layoffs in a given county and the change in that county's unemployment rate over the subsequent six- to twelve-month period. Thus, the regression results appear to reflect the effect of local labor market conditions on the fortunes of displaced workers, not the reverse.

[23] Total state employment reached its trough in August 1992. The "pre-July-1992" dummy is equal to the number of months between the layoff and July 1992, or zero if the layoff occurred after July 1992. The sample includes those workers for whom the center had enrollment records in July 1992 or later. Thus, workers laid off prior to this date were included in the sample only if they had not found work by around the middle of 1992.

[24] Anecdotal evidence in Kodrzycki (1995) further suggests that laid-off workers are more likely to use government-provided reemployment services if they are available without delay.

Table 6

Contribution of Key Regression Variables to Differences in Reemployment Probabilities for Selected Industries

	Mean Value of Regression Variable					Contribution to Difference in Reemployment Probability[a]			
	All Industries	Defense-Related Manufac-turing	Computer Manufac-turing	Transpor-tation and Public Utilities	Services	Defense-Related Manufac-turing	Computer Manufac-turing	Transpor-tation and Public Utilities	Services
Recalled to Previous Job	.04	.09	.03	.03	.04	7.1	−1.3	−1.2	−.2
Early Notification of Layoff	.06	.07	.24	.03	.05	.2	6.8	−1.9	−.8
Length of Education and Training	5.3	4.6	5.9	4.5	5.0	5.6	−4.6	6.4	2.6
Unemployment Rates at Time of Layoff[b]						3.2	10.6	−.9	1.1
County Unemployment Rate	7.3	6.7	6.3	7.3	7.6				
12-month Difference in the County UR	−1.1	−.9	−.9	−1.1	−1.2				
State minus County Unemployment Rate	.0	.2	.8	−.1	−.2				
12-month Difference in State minus County UR	.2	.2	.0	.1	.1				
Wage Premium for Previous Job	.7	12.4	8.8	5.8	−1.9	−6.2	−3.6	−2.6	1.3
Percent Change in Employment of Previous Industry	−.4	−2.8	−14.3	1.3	3.2	−3.6	−18.6	2.4	5.6

[a]Calculations based on the second set of regression results in Table 4.

[b]Unemployment rates for layoff dates after June 1992. Contributions to reemployment probability based on pre-July-1992 dummy in addition to unemployment rates.

Source: Author's calculations based on sample of displaced workers from Massachusetts who were reemployed at a new job.

results are shown in Table 6 for those variables that accounted for large differences across industries. The left side of the table shows the average values of key characteristics by industry, while the right side of the table shows their estimated contributions to the probability of reemployment.

The roles of industry wage premia and employment growth rates, as well as the high recall rate for defense manufacturing workers, have already been highlighted. For computer workers, the table indicates two reasons why reemployment was less problematic than might be expected based on the extensive layoffs in this industry. Twenty-four percent of the computer sample started receiving reemployment assistance prior to being laid off, far above the 6 percent rate overall. These were mostly employees of a single large computer manufacturer.[26] Second, computer layoffs took place in areas where the economy was relatively strong: The county unemployment rate averaged a full percentage point below that for the sample as a whole.[27] Unlike the case for computer industry workers, the regressions were not able to shed much light on the reasons behind lengthy jobless spells for workers from the transportation and public utilities industry.

[25] For education, occupation, and demographic mix, the shares of workers in each category were used. The recall rate and the use of assistance services prior to layoff were industry averages. In interpreting the results, it is useful to bear in mind that, because of the nonlinear specification of the Cox model, the average probability of reemployment for workers in an industry is not the same as the probability of reemployment for a worker with average characteristics for that industry.

[26] Of those laid off by this employer, more than one-third registered for services in advance of their layoff.

[27] The study did not investigate another potential explanation: the extent to which computer manufacturing workers were able to be reemployed in related nonmanufacturing industries such as computer software, which were expanding.

Table 7
Earnings and Job Characteristics for Workers Employed in a New Job

	Nominal Hourly Wage at Previous Job (Mean)	Nominal Hourly Wage at New Job (Mean)	Percent Difference Between Real Hourly Wage at New and Old Jobs (Mean)	No Medical Insurance at New Job (Percent)
All Industries	13.12	11.34	−12.7	24.7
Manufacturing	13.90	11.69	−16.4	22.7
Defense-Related	16.66	13.31	−21.6	24.9
Computers	17.50	15.12	−14.4	16.9
Other	12.27	10.40	−15.0	23.2
Nonmanufacturing	12.36	11.02	−8.8	26.4
Construction	13.88	12.20	−10.6	30.7
Transportation and Public Utilities	14.51	12.55	−9.9	22.2
Trade	11.53	9.68	−10.7	28.8
Finance, Insurance, and Real Estate	11.52	10.09	−11.1	27.7
Services	12.82	12.05	−6.2	24.2
Government	12.58	11.48	−7.1	21.8
Defense	12.53	12.34	3.3	8.9
Nondefense	12.59	11.26	−9.7	27.6
Other and Not Known	11.86	10.54	−9.7	32.1

Pay and Other Job Characteristics for Reemployed Workers

Table 7 examines earnings and other job characteristics for workers employed in a new job. Average earnings at the new job were $11.34 an hour, $1.78 less than at the previous job. Adjusted for inflation, the mean wage loss was 12.7 percent.

Earnings fell for displaced workers in every private industry, but the losses were most severe for manufacturing workers. On the whole, displaced defense and computer manufacturing workers found new jobs with relatively high pay—17 percent and 33 percent, respectively, above the displaced-worker sample average. But these workers had earned even greater premia in their old job. The real wage loss averaged 21.6 percent for former defense workers and 14.4 percent for former computer workers. Other manufacturing workers had been earning less than the sample average, and they slipped further behind at their new job. Their real wage losses averaged 15 percent. The lowest rate of real earnings decline, 6.2 percent, was experienced by service industry workers. At their old jobs, they had been earning about 50 cents less per hour than other displaced workers; but at reemployment, their pay was about 70 cents higher than average. The only group to experience a wage increase on average were government defense workers. The results may convey too positive an impression, however, because the bulk of the layoffs took place at the end of 1993, so the statistics refer mostly to workers who accepted a new job rather quickly after being laid off.[28]

A more comprehensive measure of the income loss at the new job would take into account losses of benefits, but the data set includes information only for the new job. One-quarter of reemployed workers had no medical insurance at their new job, and 38.5 percent had no pension benefits.[29] Altogether, 8.5

[28] Only 78 government defense workers are represented in the statistics on new wages, compared to a total of 477 laid off.

[29] The information on medical insurance indicates only whether the employer offered a group plan, not the fraction of the insurance premium paid by the employer. It is not clear how the responses regarding pension plan coverage treat retirement plans funded entirely by the employee (such as individual retirement accounts with investment options set by the employer). To provide a context for the estimates for reemployed workers, 13 percent of all employed New Englanders were estimated to lack medical insurance in 1994 (Sum et al. 1996). This rate ranged from a low of 5 percent in government to a high of 32 percent in construction. About 8 to 9 percent of manufacturing, transportation and public utilities, finance-insurance-real estate, and professional services workers lacked health coverage. Greater percentages of trade and nonpro-

Table 7 continued
Earnings and Job Characteristics for Workers Employed in a New Job

No Pension at New Job (Percent)	Real Hourly Wage Reduced by One-Third or More and No Benefits at New Job (Percent)	Miles to New Job (Median)	New Job Between 30 and 100 Miles Away (Percent)	New Job More Than 100 Miles Away (Percent)	Observations in Sample
38.2	8.5	9.2	7.0	2.7	10,374
37.6	8.4	9.9	7.6	2.7	5,281
37.1	13.0	11.4	8.4	3.3	1,136
35.0	5.9	11.1	9.0	3.5	696
38.3	7.4	9.1	6.9	2.2	3,449
38.0	8.2	8.2	5.9	2.3	3,821
44.9	9.6	11.1	7.8	3.4	194
37.9	8.7	9.6	11.4	6.4	225
42.0	9.6	8.7	6.5	1.8	1,358
40.3	10.3	7.0	3.5	1.6	466
33.3	6.3	8.2	5.2	2.3	1,578
34.6	4.0	11.4	9.5	9.5	418
12.5	3.9	19.3	18.9	20.8	78
44.7	4.1	8.3	5.2	4.3	340
46.0	12.5	7.9	7.9	2.8	854

Source: Author's calculations based on sample of displaced workers from Massachusetts who were reemployed at a new job.

percent of the reemployed workers had their real hourly wage reduced by at least one-third, while not receiving medical or pension benefits. By this measure, severe income impacts were most common among former defense manufacturing and finance, insurance, and real estate workers. Interestingly, even though former employees of computer manufacturers had high average wage losses, relatively few ended up in new jobs that lacked medical insurance and pension benefits. Service industry and government workers, the groups with the lowest average wage declines, also had a low incidence of severe wage-and-benefit losses.

Another indicator of adjustment costs is the relative convenience of the new job. Some workers may accept greater wage losses in return for being able to get a job locally. Local jobs mean that workers are able to avoid either the expenses of a long commute or the financial and psychological costs of moving. Again, the data set is imperfect, as it permits only the calculation of the distance between the worker assistance center (not the worker's home) and the new job, and it does not specifically indicate whether a move took place. Nevertheless, the data suggest that manufacturing workers' sharp wage losses were not offset by shorter commutes. At 11-plus miles, the average commutes for former defense and computer manufacturing workers were more than 20 percent longer than the sample average; for other manufacturing workers, typical commutes were comparable to the sample average. On the other hand, the data suggest that federal government defense workers were able to replace their former wages in part because many moved away from the area, presumably to places where the local economy was stronger.[30]

Determinants of Earnings Losses

Regressions were used to measure the effects of individual factors on real earnings replacement rates. Table 8 indicates the results of regressions that in-

fessional services workers were without insurance. Sum et al. estimate that 57 percent of New England workers lacked pension plan coverage, but this ratio is skewed by very low coverage among young workers. The ratio also varies by industry; only about one-third of the employees in durables manufacturing and transportation and public utilities were without pension benefits.

[30] As was the case with the Loring Air Force Base sample discussed in Kodrzycki (1995), some of the civilian employees at Fort Devens probably were married to military personnel who were transferred when the base closed.

Table 8

Real Hourly Wage Replacement Rate: Regression Results

Dependent Variable = New Real Hourly Wage as a Percent of Previous Real Hourly Wage

	(1)		(2)	
Independent Variable	Coefficient	Standard Error	Coefficient	Standard Error
Experience				
Potential Work Experience	−.22**	.04	−.23**	.04
Job Tenure	−1.77**	.13	−1.71**	.13
Job Tenure Squared	.04**	.00	.04**	.01
Education and Skills				
Schooling (Omitted = Less than High School)				
High School	−.59	1.49	−.56	1.49
Some College	1.15	1.58	1.14	1.59
College Degree	4.22**	1.75	4.07**	1.76
More than College	4.79*	2.48	4.98**	2.48
Reading Test Score	.23	.17	.25	.17
Previous Occupation (Omitted = Services)				
Professional, Technical, and Managerial	−2.47	1.98	−2.41	1.99
Clerical and Sales	1.67	2.01	2.21	2.03
Production	−7.01**	2.05	−6.18**	2.10
Other	−2.69	2.80	−2.58	2.82
Not Known	−8.76**	2.14	−7.51**	2.21
Switched Occupation	−5.43**	.99	−5.18**	.99
Duration of Unemployment	−.76**	.06	−.74**	.07
Location of New Job				
Distance	.02**	.01	.02**	.01
Distance Squared	−8.1e-06**	3.0e-06	−8.0e-06**	3.0e-06
Work Effort				
Full-Time at Previous Job	−1.46	2.40	−1.07	2.41
Switched to Part-Time	−1.82	1.19	−1.90	1.19
Switched to Full-Time	8.09**	3.11	8.39**	3.11
Demographic Characteristics				
Gender and Marital Status (Omitted = Unmarried Male)				
Married Male	−1.42	.99	−1.37	.98
Married Female	1.97	1.20	1.71	1.20
Unmarried Female	1.44	1.04	1.10	1.04
Nonwhite	2.36*	1.30	2.23*	1.30
Previous Industry				
12-Month Employment Growth Rate	−.14	.09	−.03	.19
Wage Premium	−.29**	.07	.03	.29
Dummies (Omitted = Government)				
Defense-Related Manufacturing			−9.73**	3.15
Computer Manufacturing			−1.96	4.37
Other Manufacturing			−2.17	3.00
Construction			[a]	
Transportation and Public Utilities			−5.94*	2.91
Trade			−2.30	4.38
Finance, Insurance, and Real Estate			−2.40	2.88
Services			−.95	2.98
Other and Not Known			−4.45	2.97
Switched Industry	−3.32**	.77	−3.03**	.86
Constant	109.68**	3.75	111.27**	4.46
Adjusted R^2	.140		.143	
Number of Observations	5,492		5,492	

*Significantly different from zero at 5 percent level.

**Significantly different from zero at 1 percent level.

[a]Dropped because of collinearity.

Source: Author's calculations based on sample of displaced workers from Massachusetts who were reemployed but not recalled to their previous job.

cluded as explanatory factors the worker's experience and skills, education, occupation, industry characteristics, work effort, duration of unemployment, distance from the new job, and demographic characteristics such as gender, race, and marital status. The second regression also includes industry dummies.[31]

Previous experience was discounted. Potential work experience measures the maximum number of years a worker could have spent in paid employment, and is measured as age less years of education less six. The estimated coefficient implies that, assuming equal years of education and other characteristics, 10 years of added age results in a 2 percent greater wage loss.[32] The job tenure coefficients imply that specific work experience at the previous employer generally was discounted even more heavily. For example, someone who had spent 20 years in his or her last job would expect to have a real wage loss about 6 percentage points greater than someone employed in the same job for only 10 years.[33]

The greater wage losses for defense manufacturing workers suggest that employers discounted their previous experience more than that of other job applicants.

Employers valued a college education. Graduating from college or pursuing postgraduate studies reduced the average wage loss (or increased the average wage gain) by 3 to 5 percentage points, compared to receiving only a high school education or less. The worker's score on a reading test administered by the worker assistance center had only a tiny effect on the wage outcome (as well as being statistically insignificant).[34]

Previous production workers had the worst wage outcomes. Their losses were 4 to 7 percentage points greater than those in clerical, sales, or services positions. Regardless of initial occupation, however, those who switched occupations in their new job lost out.

On average, those who searched longer for a job and those who ended up reducing their work hours tended to accept greater pay cuts. The coefficient on the duration of unemployment indicates that, for each additional year of looking for a job, the real replacement rate falls by 9 percentage points.[35] A switch from part-time to full-time work increased the hourly replacement rate substantially. Expanding one's job search to faraway locales boosted wages, but not by much. Being willing to commute 100 miles was estimated to raise pay by only 2 percent. Gender, marital status, and race had little effect on wage outcomes, controlling for other factors.

The remaining variables address industry effects, apart from inter-industry differences in years of experience, education, and so forth. In the absence of industry dummy variables, the average wage premium enters with a negative sign, as expected under the hypothesis that wage losses are the result of the disappearance of good jobs. But with the addition of industry dummies, the industry wage premium has no remaining effect on the wage replacement rate. *The wage losses for defense manufacturing workers are 4 to 10 percentage points greater than for workers from other industries, holding all else constant.* This suggests that employers discounted the previous experience of defense workers more than that of other job applicants. Finally, the regressions indicate that those who switch industries had about a 3 percent greater wage loss on average.

[31] Unemployment rate measures, which helped to explain reemployment probabilities, were never significant in the wage regressions. Thus, the improvement in economic conditions from the early to the mid 1990s was not found to raise the quality of jobs obtained by displaced workers, after adjusting for other factors. In addition to the specifications shown, another version used the two-step Heckman procedure to adjust for sample selection bias. Results with the Heckman correction were indistinguishable from those using ordinary least squares.

[32] These estimates were quite similar when previous occupation was excluded from the regression. Thus the occupation results, mentioned below, are not picking up some of the effects of education, even though average education levels vary somewhat across occupations.

[33] The effect of additional years of job tenure diminishes as tenure increases, but remains negative up to 41 years. The average job tenure for all workers in the sample was about 7 years.

[34] The effect of scoring in the 75th as opposed to the 25th percentile was estimated to increase the wage replacement rate by only one-half of a percentage point. This small contribution may be due to the fact that the test does a poor job measuring differences in reading ability across individuals; on a scale of 1 to 13, only 2 percent of the sample scored below 7 and 24 percent scored above 9.

[35] This result reflects workers' falling reservation wage (that is, the wage required for them to accept a job) as time passes. However, it also reflects a negative trend in real wages over this period. Another version used the unemployment spell predicted on the basis of a regression, in order to correct for the fact that, all else equal, those workers who accept a job offer have lower reservation wages than workers who reject job offers. This had very little effect on the estimates.

Table 9
Contribution of Regression Variables to Differences in Real Wage Replacement Rates for Selected Industries

	Mean Value of Regression Variable					Contribution to Difference in Real Wage Replacement Rate[a]			
	All Industries	Defense-Related Manufacturing	Computer Manufacturing	All Other Manufacturing	Services	Defense-Related Manufacturing	Computer Manufacturing	All Other Manufacturing	Services
Education	12.8	13.5	13.5	12.0	13.5	.4	.5	−.4	.4
Tenure	7.3	10.6	10.4	8.3	5.2	−3.0	−3.9	−.4	2.1
Occupational Mix						−.7	.1	−.9	1.2
Professional, Technical, and Managerial	.29	.35	.54	.19	.47				
Clerical and Sales	.21	.13	.15	.15	.21				
Services	.04	.01	.00	.01	.12				
Production	.27	.38	.15	.45	.10				
Other Occupations	.04	.01	.06	.05	.02				
Occupation Not Known	.19	.14	.16	.19	.10				
Switched Occupation	.30	.23	.29	.29	.32	−.3	.0	.2	−.1
Industrial Mix	[b]					−6.6	1.2	.9	2.2
Switched Industry	.65	.97	.86	.64	.52	−.6	−.4	−.0	.2
Length of Unemployment	7.9	7.8	7.1	8.1	7.1	.1	.1	−.2	.5
All Other Variables						−.1	.5	−.8	.6
Total Difference in Real Wage Replacement Rate									
Explained						−8.1	−1.8	−1.6	7.0
Actual						−9.1	−1.5	−2.0	6.6

[a]Calculations based on the second set of regression results in Table 8.
[b]See Appendix Table 1.
Source: Author's calculations based on sample of displaced workers from Massachusetts who were reemployed at a new job.

Additional regressions explored whether education and training courses helped to boost wage replacement rates. The results were mixed, and are discussed below in the section on "Education and Training."

Explaining Wage Changes by Industry

The regressions were used to examine why workers from declining industries did worse than others (Table 9). At 21.8 percent, the average real wage loss of reemployed defense manufacturing workers was 9 percentage points greater than the sample-wide average. Roughly two-thirds of this gap is explained by the defense dummy. That is, prospective employers discount the skills and experience of former defense workers more than those of other workers. Almost all defense workers who were not recalled switched industries. In addition, many laid-off defense workers had been in production jobs, the occupational category with the largest earnings declines across all industries. Finally, on average, defense workers had been employed at their previous employer for over 10 years, compared to about 7 years for the full sample. As found in the regressions, each added year of experience at the past employer is valued less by the new employer.

Service industry worker wages fell only 6.1 percent, 6.6 percentage points less than average. Low job tenure, favorable occupational mix, and better-than-average education were especially significant in explaining their relatively high wage replacement rates.

Service industry workers also did not have to look as long, so they were not as desperate. Interestingly, however, the fraction of service workers remaining in service work was only 48 percent, even though on the whole service industries were increasing their employment during this time period. This is one indication of the pervasiveness of structural change in labor markets.

III. Education and Training

All displaced workers in the sample were offered basic readjustment assistance. This consisted of group workshops designed to help them cope with unemployment and undertake a job search, as well as individual meetings with job counselors. The centers also made available resources that could be used directly in their job search, such as phone banks and job listings.

In addition to receiving basic assistance, 42 percent of the sample enrolled in education or training classes approved and funded by the centers (Table 10). The most common course of study was occupational training. This consisted of preparing for a new job by taking classes related to a particular employment field at a local university, community college, or specialized training facility. Typical subjects included computer programming, equipment or machinery repair, accounting, culinary arts, truck driving, and health sciences.[36] Two percent received entrepreneurial training, to help them start their own businesses. Education programs took three forms: basic education to improve reading, writing, mathematics, and computer literacy skills; English as a second language (ESL); and GED classes, to obtain a high-school equivalency diploma. Between 1 and 4 percent of the sample enrolled in at least one of these courses of study. Education and training programs on balance appear to have had mixed but generally small effects on wage outcomes (Table 11). According to the first regression, workers enrolled in education and training programs had approximately the same wage replacement rates as those who did not, holding other qualifications constant. When the effects of various types of education and training were measured separately (equation 2), occupational skills training (the most popular course of study) remained unhelpful in explaining wage outcomes. Adult basic education, ESL, and integrated training were associated with a positive effect on wages, and GED classes a negative effect. However, the standard errors were large, implying that these effects were measured with a high degree of uncertainty. Those receiving entrepreneurial training had substantially lower wage replacement rates than others, confirming the findings of Bradbury (1994) that self-employment was a useful, but not very lucrative, stopgap in the early 1990s. It is possible, however, that entrepreneurs received some measure of satisfaction from being their own boss, or that, at least for some, earnings grew rapidly as their business became more established. The third regression includes a dummy variable equal to one for those workers whose new job was related to the training they received; this variable indicated a small positive effect on wages.

The final two regressions include length of enrollment in education and training programs. The first of these specifications indicates that longer enrollment was associated with a small reduction in the new wage: Displaced workers enrolled for six months had wage replacement rates 2.5 percentage points lower than those who did not enroll at all, assuming all other characteristics were similar. Workers who got very

Table 10
Summary of Education and Training Programs

	Participation Rate (percent)	Median Duration (months)	Maximum Duration (months)
Education			
Adult Basic Education	4.2	4	23
English as a Second Language	3.0	6	32
GED Class	1.4	5	33
Training			
Occupational Skills	33.6	4	44
Entrepreneurial Training	2.0	2	18
Integrated Training[a]	.4	6	30
On-the-Job Training	.0	5	5
All Education and Training Programs	42.0	4	44

Note: Long maximum duration of education and training reflects workers who enrolled in training prior to being laid off.

[a]Integrated Training combines occupational classroom training with ESL, Basic Education, or GED training. This single integrated training program is not the same as occupational and basic education courses taken concurrently or sequentially.

[36] This list is drawn from the anecdotal evidence used in Kodrzycki (1995). The subject matter of the occupational training is not available in the computerized records obtained for the current sample.

Table 11
Real Hourly Wage Replacement Rate: Regression Results
Including Various Measures of Education and Training

Independent Variable	(1)	(2)	(3)	(4)	(5)
Experience					
Potential Work Experience	−.23**	−.22**	−.22**	−.18**	−.18**
Job Tenure	−1.71**	−1.73**	−1.72**	−2.43**	−2.43**
Job Tenure Squared	.04**	.04**	.04**	.06**	.06**
Education and Skills					
Schooling (Omitted = Less than High School)					
High School	−.59	−.09	−.66	2.49	2.42
Some College	1.11	1.89	.95	2.87	2.77
College Degree	4.02**	5.07**	3.98**	3.35	3.25**
More than College	4.90**	6.11**	4.97**	9.69	9.52**
Reading Test Score	.24	.30	.27	−.00	.08**
Previous Occupation (Omitted = Services)					
Professional, Technical, and Managerial	−2.40	−1.96	−2.42	−8.18**	−8.15**
Clerical and Sales	2.23	2.28	2.11	.76	.73
Production	−6.15**	−6.18**	−6.12**	−7.78**	−8.01**
Other	−2.57	−2.62	−2.51	−1.87	−1.86
Not Known	−7.53**	−7.80**	−7.38**	−5.55	−5.74
Switched Occupation	−5.15**	−5.10**	−5.20**	−2.04	−2.16
Duration of Unemployment	−.75**	−.72**	−.74**	−.53**	−.45**
Location of New Job					
Distance	.03**	.02**	.02**	.04**	.04**
Distance Squared	−8.08e-06**	−8.48e-06**	−7.63e-06**	−1.73e-05**	−1.72e-05**
Work Effort					
Full-Time at Previous Job	−1.08	−1.20	−1.07	−3.43	−3.42
Switched to Part-Time	−1.86	−2.23	−2.07	.17	−.05
Switched to Full-Time	8.41**	8.53**	8.36**	11.77**	11.69**
Demographic Characteristics					
Gender and Marital Status (Omitted = Unmarried Male)					
Married Male	−1.34	−1.17	−1.42	−.48	−.49
Married Female	1.81	1.71	1.38	1.64	1.48
Unmarried Female	1.17	1.19	.89	.34	.22
Nonwhite	2.22	2.09	2.24	4.22	3.97
Previous Industry					
12-Month Employment Growth Rate	−.04	−.04	−.02	.11	.24
Wage Premium	.02	.05	.02	−.40	−.44
Dummies (Omitted = Government)					
Defense-Related Manufacturing	−9.84**	−9.76**	−9.39**	−4.86	−3.75
Computer Manufacturing	−2.19	−1.99	−1.45	5.94	8.54
Other Manufacturing	−2.21	−1.75	−2.30	−2.98	−2.90
Construction	[a]	[a]	[a]	[a]	[a]
Transportation and Public Utilities	−5.95*	−5.70*	−6.06**	−7.17	−6.98
Trade	−2.34	−1.60	−2.43	−5.06	−5.48
Finance, Insurance and Real Estate	−2.43	−2.10	−2.58	−4.32	−4.34
Services	−.98	−.42	−1.08	−4.18	−4.33
Other and Not Known	−4.25	−2.42	−5.20	−7.17	−7.21
Switched Industry	−2.83**	−2.91**	−3.59**	−1.71	−1.61

continued

Table 11 continued

Real Hourly Wage Replacement Rate: Regression Results

Including Various Measures of Education and Training

Independent Variable	(1)	(2)	(3)	(4)	(5)
Education and Training					
Received Education or Training	−.65				
Attended Adult Basic Education Classes		4.17			
Attended ESL Classes		6.44			
Attended GED Classes		−2.21			
Received Occupational Skills Training		.18			
Received Entrepreneurial Training		−13.55**			
Received Integrated Training		10.51			
New Job Related to Training			−2.11**		
Total Duration of Education and Training				−.41**	
Adjusted Total Duration of Education and Training					−.23
Constant	111.54**	109.51**	111.06**	115.82**	113.21**
Adjusted R^2	.143	.148	.144	.137	.136
Number of Observations	5,492	5,492	5,492	2,256	2,256

*Significantly different from zero at 5 percent level.

**Significantly different from zero at 1 percent level.

[a]Dropped because of collinearity.

Note: Standard errors available from the author upon request.

Source: Author's calculations based on sample of displaced workers from Massachusetts who were reemployed but not recalled to their previous job.

attractive job offers while enrolled in classes would be likely to cut short their course of study, however, producing a negative correlation between the wage replacement rate and the duration of education and training. The last regression attempts to correct for this bias. The duration of education and training programs was measured only for those enrollees who continued to be registered at the worker assistance center after their last class had ended. Presumably this omits anyone who interrupted his or her coursework as a result of an attractive job offer. The coefficient on the duration variable becomes less negative in this specification; it still does not indicate a positive association between longer education and training and the new wage.

In summary, the regressions tend to indicate little overall wage effect from education and training, while not closing the door on the possibility of a larger effect for some programs. Two explanations for the finding of a minimal wage gain seem plausible while being consistent with the view that education and training is beneficial to displaced workers. The first explanation is that most workers are enrolled for a rather short time. For all individuals receiving education and training programs, the median duration was only four months, and only two programs (English as a second language and integrated training) had median durations as long as six months. The added skills acquired may simply have been too minor to matter much. They might have enabled some to find an entry-level position in another field, but they could not compensate for the lack of a college degree or detailed knowledge of a particular field.[37]

Another possibility is that those who decided to enroll in education and training were, in one way or another, at a greater disadvantage than was apparent in the regressions. For example, their math skills may have been more deficient than indicated by their educational attainment. GED classes may have helped high school dropouts with poor math skills get better jobs than they would otherwise have gotten, but not better jobs than other high school dropouts who had

[37] A similar conclusion was reached by Jacobson, LaLonde, and Sullivan (1994). In a study of education programs for displaced workers in Pennsylvania, they estimated that a year of schooling raised long-term earnings by 6 to 7 percent for male participants, and 3 to 4 percent for female participants. But most participants acquired less than one year of education, despite the fact that the program was subsidized. Therefore the average wage effect was smaller than the estimated annual rate of return.

learned math on the job. Or to take another example, a laid-off physical education teacher who took courses to enter the health care field may have remained a "professional, technical, and managerial" worker in the data. This would be better than settling for a job at a health club, but other professionals with more marketable skills may have gotten still better positions without undergoing occupational training. The regressions can be used to compare wage outcomes of those receiving education and training to the outcomes of those who did not—but not to measure the hypothetical outcomes in the absence of education and training.

IV. Conclusions

This study has examined the experiences of a large group of workers from Massachusetts who were laid off in the early 1990s and who sought government assistance in finding a job. It provides evidence on their difficulty in finding reemployment, the extent to which they were able to obtain new jobs that were as attractive as their former positions, and on the ways in which government services were able to help.

In general, displaced workers experienced noticeable wage losses. Many, though not a majority, ended up finding jobs either without medical insurance or without pension benefits. Most displaced workers experienced an extended period of joblessness, although the duration depended somewhat on general economic conditions. Jobs were easier to find toward the end of the sample period than the beginning, as total Massachusetts employment was increasing rather than decreasing, and as the state unemployment rate fell from 9 percent in 1991 to 6 percent in 1994.

Experiences differed across categories of workers. Older, more experienced workers had longer durations of joblessness and lower wage replacement rates. Educational background had mixed effects. On the one hand, less educated workers tended to be reemployed more quickly than college-educated workers. On the other hand, their new wages tended to be considerably lower relative to their old job. Thus, the study points out that reemployment outcomes depend not only on changes in the relative demand for different types of workers (such as the increasing value placed on education), but also on differences in how long job seekers feel they can hold out without a paycheck.

Being laid off from a declining industry tended to result in a longer duration of joblessness; thus workers from the fastest-growing industries, services and construction, were quicker to find a new job than many former manufacturing workers, including those from defense industries. But other factors sometimes offset the influence of industry trends. Despite extensive layoffs in their industry, computer manufacturing workers did not experience abnormal difficulty finding work, in part because they tended to be located in areas of the state with relatively low unemployment and because many registered for reemployment assistance before losing their old job.

Workers from declining industries tended to suffer sharper earnings cuts than others. Steep earnings losses in part related to long tenure at their former job and the prevalence of production (rather than sales or services) skills. Defense manufacturing workers' large wage cuts upon reemployment appeared also to reflect their new employers' belief that experience at a defense contractor firm was particularly inapplicable to other industries.

Early sign-up at a worker assistance center was found to reduce the period of joblessness. This result indicates that employers can mitigate the costs of layoff, as workers are able to register for government services prior to being laid off only if they receive advance notification of impending layoffs.

Many displaced workers received government funding for education and training, in addition to the counseling and job market information services that were available to all workers in the sample. Participation in education and training tended to lengthen joblessness, as workers were less likely to be actively looking for a job while taking classes. On the whole, however, workers who enrolled in education or training obtained jobs that paid about the same as those who received only basic services, after adjusting for other measurable differences in qualifications. One explanation for this finding is that education and training services were used disproportionately by job seekers who faced particularly large difficulties recouping their former wage or who decided to make more dramatic changes in their line of work, in ways that the available data could not detect. Another explanation lies in resource constraints, which limited the number of classes workers could take, as well as their willingness to turn down job offers in order to train for better opportunities.

Appendix, by Margaret E. Enis

Appendix Table 1 provides statistics for a data base of 20,624 displaced workers provided by the Massachusetts Industrial Services Program. Potential work experience was computed as age minus years of education minus six. Most observations in the data set included a reading test score in the form of a grade level equivalent, but for a few only a raw score on a standardized reading test was given. For those few observations, the actual number of years of education was substituted for the raw score. College graduates, who were not tested, were assigned the highest possible grade level equivalent in reading (13).

The observed length of nonemployment was calculated as the number of months between the date of layoff and the date of termination from the center. The date of layoff was taken as the ending date of work at the former employer. The date of application at the worker assistance center was used as the layoff date if the end date was unknown. In cases where the date of termination was not available because the sample period ended, the observed nonemployment spell was calculated as the number of months between the layoff and the end of the sample (September 1994). The period of unemployment is measured as the months of nonemployment not spent in education and training.

The length of education and training, the total months spent in training programs sponsored by the Industrial Services Program, is measured as the sum of the number of months between the enrollment and completion dates for each of the activities in which the worker participated. If a worker was enrolled in training prior to displacement, the length of this advance training was calculated as the number of months between the first day of training and the layoff date. The adjusted length of education and training is the length of education and training for workers who did not terminate at the center on the same day that they ended training.

The job tenure variable, measuring the years of employment at the former employer, is the length of time between the start date and the end date at the former employer. The worker assistance centers included a code for those

Appendix Table 1
Displaced Worker Sample Variables

Variable	Mean	Standard Deviation	Number of Observations
Worker			
Potential Work Experience (years)	22.2	10.5	20,495
Age (years)	41.0	10.3	20,624
Education (years)	12.8	2.4	20,495
Number of Dependents	.9	1.2	20,624
Nonwhite (proportion)	.13	.34	20,624
Male (proportion)	.53	.50	20,624
Married (proportion)	.50	.50	20,624
Adjusted Reading Score	8.4	2.4	19106
Employment Status			
Observed Length of Nonemployment (months)	11.5	7.4	20,412
Observed Length of Unemployment (months)	9.2	6.8	20,370
Education and Training			
Observed Duration of Education and Training (months)	5.3	4.4	8,662
Enrollment in Training Prior to Displacement (proportion)	.06	.25	20,624
Duration of Training Prior to Displacement (months)	.17	1.33	20,447
Adjusted Duration of Education and Training (months)	2.7	3.5	4,786
Former Job			
Hourly Wage (dollars)	13.2	6.0	20,371
Hours Per Week	39.4	4.6	20,421
Job Tenure (years)	7.7	7.9	18,479
Recalled to Former Job (proportion)	.04	.20	20,624
Employed Full-Time at Former Job (proportion)	.95	.22	20,624
Proportion formerly employed in:			
Defense Manufacturing	.14	.35	20,624
Computer Manufacturing	.07	.25	20,624
Other Manufacturing	.32	.47	20,624
Construction	.02	.13	20,624
Transportation, Communications, and Public Utilities	.02	.15	20,624
Wholesale and Retail Trade	.12	.32	20,624
Finance, Insurance, and Real Estate	.04	.20	20,624
Services	.15	.35	20,624
Government, Defense-Related	.02	.15	20,624
Government, Not Defense-Related	.03	.18	20,624
Other and Not Available	.07	.25	20,624
Former Industry			
12-Month Employment Growth Rate (percent)	−.4	4.6	20,624
Wage Premium (percent)	.7	6.3	20,624

continued

Appendix Table 1 continued
Displaced Worker Sample Variables

Variable	Mean	Standard Deviation	Number of Observations
Former Occupation (proportion):			
Professional, Technical and Managerial	.30	.46	20,624
Clerical and Sales	.21	.40	20,624
Production	.29	.45	20,624
Service	.05	.22	20,624
Miscellaneous Occupations	.04	.19	20,624
Not Available	.12	.32	20,624
New Job			
Hourly Wage (dollars)	11.5	5.7	9,371
Hours Per Week	38.0	5.3	9,062
Distance to New Job (miles)	33.9	182.6	7,401
Comparison between Old and New Job			
Percent Difference between Former and New Real Wage	−11.7	28.0	9,276
Hourly Replacement Wage (percent)	88.1	28.0	8,811
Switched Occupation (proportion)	.54	.50	20,624
Switched Industry (proportion)	.76	.42	20,624
Hours Decreased from Full-Time to Part-Time (proportion)	.04	.20	20,624
Hours Increased from Part-Time to Full-Time (proportion)	.04	.21	20,624
Unemployment			
Country Unemployment Rate at Time of Displacement (percent)	7.9	1.9	18,717
12-month Change in County Unemployment Rate (percentage points)	−1.0	1.0	18,717
Difference between State and County Unemployment Rates at Time of Displacement (percentage points)	−.1	1.5	18,717
Difference between 12-month Change in State and County Unemployment Rates (percentage points)	.1	.8	18,717

Source: See Appendix text.

recalled to their former job. An individual was also considered to be recalled to his or her old job if the name and location of the former and current employer were the same. For categorization purposes, any individual working 35 or more hours per week was determined to be employed full-time.

The industries of the former and new jobs were grouped using the Standard Industrial Classification codes. SIC codes for the former employers were included in the data base, but they appeared in the form of 2-digit, 3-digit, and 4-digit codes. The 2- and 3-digit SIC codes were changed into 4-digit codes by adding zeros. Observations that did not have an SIC code for the former employer were supplied with one if it could be determined from the name of the employer. Miscoded SICs were corrected using the employer name. Manufacturing jobs were determined to be defense-related if the employer appeared on the 1993 list of "Prime Contractors Plants with Awards Totaling $5 Million of More During FY 1992," from the Department of Defense, or was known to have appeared on previous lists. Defense-related employers whose SIC codes fall in the computer manufacturing category were classified as computer manufacturers. Government jobs were determined to be defense-related if the employer was a military base or otherwise known to be defense-related. The defense-related government workers in this sample are civilians, as military employees have access to separate reemployment services.

The 12-month employment growth rate was computed as the percent change in Massachusetts employment in the industry one year after the time of layoff. The industry employment levels were categorized by SIC code except for defense- and non-defense-related government, which were taken as federal government employment (not including postal workers) and state and local government employment plus postal workers, respectively. The wage premium was computed as the percent difference relative to government and is estimated in Appendix Table 3.

The former and current occupations were categorized using the classification codes from the Dictionary of Occupational Titles. Observations without a DOT code or with a code that does not correspond to any occupation in the Dictionary of Occupational Titles were classified as "occupation not available." Although the Dictionary of Occupational Titles classifies graphic designers as "miscellaneous," they are classified here in the "Professional, Technical, and Managerial" category.

To calculate the distance to new job, the zip codes for the worker assistance center and the new employer were matched to their latitude and longitude centroids using ATLAS GIS for Windows, version 2.0. This software includes data for all U.S. zip codes. The distance between these two centroids was converted to miles using Geodist, a C program written by Philip Thompson at MIT's Computer Resource Lab. The distance to new job was not calculated for the relatively few workers who moved overseas.

The hourly replacement wage was constructed as a ratio of the hourly wage on the new job to the hourly wage on the former job. A worker was deemed to have switched occupation if the former occupation was different from the new occupation, and to have switched industries if the former

Appendix Table 2

Duration of Education and Training: Regression Results

Independent Variable	Coefficient	Standard Error
Experience		
Potential Work Experience	−.04**	.01
Job Tenure	.04**	.01
Schooling (Omitted = Less than High School)		
High School	−.41**	.16
Some College	−.46**	.18
College Degree	−1.32**	.22
More than College	−.96*	.44
Previous Real Wage	−.05**	.02
Full-Time at Previous Job	.50*	.22
Previous Occupation (Omitted = Services)		
Professional, Technical, and Managerial	−.73**	.27
Clerical and Sales	−.57*	.27
Production	−.22	.27
Other	−.25	.36
Not Known	−.09	.29
Previous Industry (Omitted = Government)		
Manufacturing		
Defense-Related	.41	.30
Computers	1.65**	.32
Other	.28	.27
Construction	−.10	.42
Transportation and Public Utilities	.63	.41
Trade	.43	.28
Finance, Insurance, and Real Estate	.35	.33
Services	.40	.28
Other and Not Known	.86**	.30
Unemployment Rate at Time of Layoff		
County	.05	.04
State	1.04**	.07
Demographic Characteristics		
Male	−.21	.15
Married	.04	.14
Married Male	−.50*	.21
Nonwhite	.91**	.15
Constant	−2.36**	.62
Adjusted R^2 = .096		
Number of Observations = 7,079		

*Significantly different from zero at 5 percent level.
**Significantly different from zero at 1 percent level.
Source: Author's calculations based on sample of displaced workers from Massachusetts.

Appendix Table 3

Log of Previous Real Wage: Regression Results

Independent Variable	Coefficient	Standard Error
Experience		
Potential Work Experience	.02**	.0009
Potential Work Experience Squared	−.0003**	.00002
Job Tenure	.02**	.0008
Job Tenure Squared	−.0004**	.00003
Education and Skills		
Schooling (Omitted = Less than High School)		
High School	.10**	.01
Some College	.18**	.01
College Degree	.35**	.01
More than College	.49**	.02
Reading Test Score	.01**	.001
Occupation (Omitted = Services)		
Professional, Technical, and Managerial	.30**	.01
Clerical and Sales	.07**	.01
Production	.14**	.01
Other	−.002**	.02
Not Known	.13	.01
Full-Time	.13**	.01
County Unemployment Rate at Time of Layoff	−.02**	.001
Year of Layoff (Omitted = 1991)		
1992	−.03**	.01
1993	−.09**	.01
1994	−.11**	.01
Demographic Characteristics		
Male	.12**	.01
Married	−.02**	.01
Married Male	.10**	.01
Nonwhite	−.06**	.01
Number of Dependents	.005*	.002
Industry (Omitted = Government)		
Manufacturing		
Defense-Related	.12**	.01
Computers	.08**	.01
Other	−.04**	.01
Construction	.12**	.02
Transportation and Public Utilities	.06**	.02
Trade	−.08**	.01
Finance, Insurance, and Real Estate	−.01	.02
Services	−.02	.01
Other and Not Known	−.02	.01
Constant	1.39**	.03
Adjusted R^2 = .532		
Number of Observations = 15,365		

*Significantly different from zero at 5 percent level.
**Significantly different from zero at 1 percent level.
Source: Author's calculations based on sample of displaced workers from Massachusetts.

industry was not the same as the new industry. Workers whose former or new occupation or industry is not known are considered not to have switched.

County unemployment rates were assigned based on the location of the former employer. If the former employer had locations in more than one county, and the particular location was unknown, the county unemployment rates for all of the possible locations in Massachusetts were averaged. The 12-month change in the county unemployment rate was computed as the difference between the county unemployment rate 12 months after layoff and the county unemployment rate at the time of layoff. The same was done for the state unemployment rate.

Appendix Table 2 shows the results of the regression estimating the duration of education and training.

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Risk and the Capital of Insurance Companies

Richard W. Kopcke

Vice President and Economist, Federal Reserve Bank of Boston. *Kathryn Cosgrove provided valuable research assistance. A version of this article was presented as a paper at a conference sponsored by the Competitive Enterprise Institute, "Rethinking Insurance Regulation," January 12, 1996, Washington, D.C.*

Insurance companies, like other financial institutions, have been evolving from specialized businesses to enterprises offering a variety of financial services. Rising interest rates impelled this evolution during much of the past three decades as most insurers tried to remain competitive. As insurers' profit margins subsided and they attracted new business, their assets generally grew more rapidly than their capital. This erosion of capital per dollar of assets for insurance companies concerned their regulators, especially as more insurers increased their investments in assets commonly regarded as risky.

To maintain the safety and soundness of insurance companies, regulators increasingly are adopting risk-based capital requirements instead of rules that limit insurers' investments and contracts. The prompt enforcement of capital requirements linked to the risks assumed by each company may reassure policyholders of the integrity of their investments in their companies without imposing excessive costs on insurers, which could diminish their capacity to serve as efficient financial intermediaries. The consequences of such policies, however, depend greatly on the design and enforcement of these requirements. These policies work best when capital requirements properly reflect the risks assumed by each insurer, when the assets and liabilities of insurance companies are priced fairly in financial markets, and when insurers may sell their risky assets, if necessary, without incurring a significant penalty. Otherwise, these policies can weaken the insurance industry by pricing inaccurately the risks assumed by insurance companies.

Existing risk-based capital regulations are not so much a new way of measuring and controlling insurers' risks as they are a new way of managing those controls. The current regulations essentially define an insurer's risk by the properties of its assets and obligations considered in isolation, not by the blend of its assets and liabilities. Accordingly, these regulations give too little credit to those companies that mitigate their risks by diversifying their investments or matching the terms of their

assets closely to the terms of their liabilities. Furthermore, prevailing measures of insurers' capital mark some assets according to their market values while marking other assets and most liabilities according to their book values. Consequently, these measures of capital can substantially misrepresent a company's capacity for bearing risk.

These apparent deficiencies in existing regulations may reflect more than problems with technical details. If markets for financial instruments are not perfect, as assumed in the reasoning supporting risk-based capital requirements, then coherent measures of risk and capital may be elusive. If the assets and liabilities of insurers are not always priced efficiently

To maintain the safety and soundness of insurance companies, regulators increasingly are adopting risk-based capital requirements instead of rules that limit insurers' investments and contracts.

in liquid markets, the strategy of promptly enforcing any capital requirement at times may undermine, rather than foster, safe and sound financial institutions. In these circumstances, promising measures of risk and insurers' capacities for bearing risk rest on judgments about the odds of future economic conditions and about the implied correlations among returns on investments; yet, the prevailing regulations, striving for a degree of simplicity and objectivity, grant these judgments little force.

I. The Promise of Risk-Based Capital Requirements

The assets of insurance companies are investments made on behalf of their owners and policyholders.[1] Policyholders' claims against these assets are defined by their contracts, which typically obligate insurance companies to make specific payments on behalf of their policyholders in the event of retirement, death, illness, accident, or natural disaster. Accordingly, insurers collect premiums from their policyholders in order to accumulate assets, designated as reserves, that are sufficient to meet these claims. While basic hazard, term life, or health insurance policies may not require substantial reserves for each dollar of coverage, other policies such as popular permanent life insurance policies, annuities, and investment contracts accumulate considerable reserves.

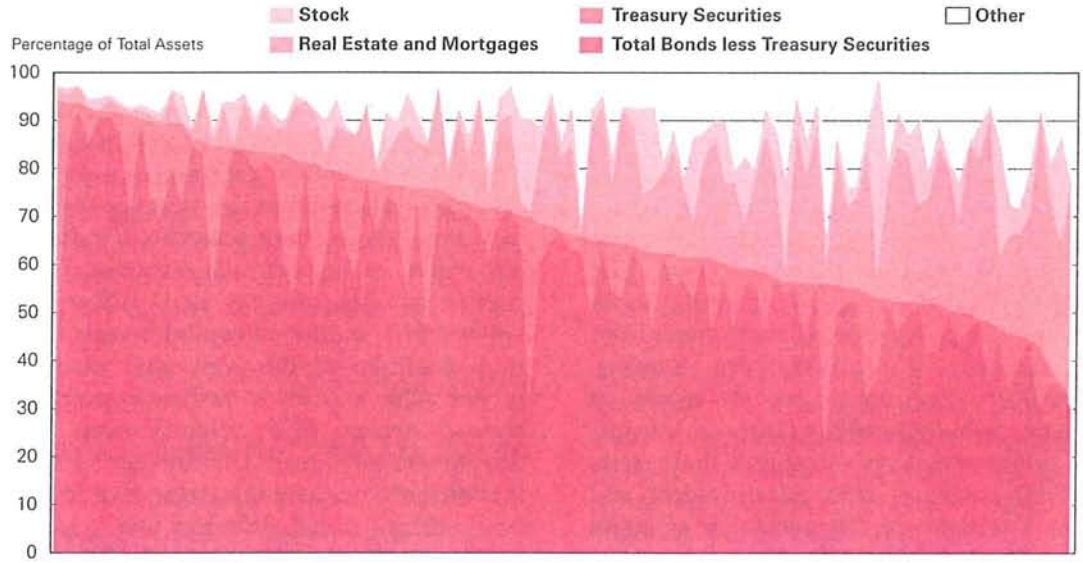
Because insurance companies continually are writing new contracts and collecting new premiums even as they are making payments as warranted by older contracts, their reserves tend to represent fairly stable portfolios of funds that they principally invest in longer-term assets such as bonds, mortgages, and equity (Figure 1). Although all invest a substantial proportion of their general accounts in bonds, the allocation of these investments among different types of bonds vary greatly. The darkest segment in each graph shows the proportion of each company's assets invested in bonds other than U. S. Treasury securities. For life insurers these bonds are most often corporate securities; property-casualty companies are more inclined to hold municipal bonds. In any case, as the graph suggests, no simple correlation exists between an insurer's commitment to bonds and the allocation of this investment among safer and riskier bonds. The graphs also show that companies that invest a smaller share of their assets in bonds tend to invest a greater share of their assets in mortgages, real estate, and equities. Accordingly, regulations that would treat the companies constituting the life or property-casualty industries equitably must weigh the consequences of their different investment strategies as well as the often considerable differences among their contracts with their policyholders.

Although not all insurance contracts are regarded as investments by policyholders, the premiums for all contracts depend on the returns companies expect to earn on their reserves. A company that earns competitive returns can afford to credit its shareholders with a competitive yield while charging a competitive premium for its contracts. When a company's return on assets is greater than expected, it can credit its shareholders with greater earnings, or charge its policyholders lower net premiums, or both. When a company's rate of return falls short of its expectation, it must reduce the yields it effectively pays to its shareholders or policyholders. If this deficiency is sufficiently great, the company also risks not being

[1] The distinction between owners and policyholders is not always sharp, especially for mutual insurance companies or for participating, experience-rated, and variable insurance policies.

Figure 1a

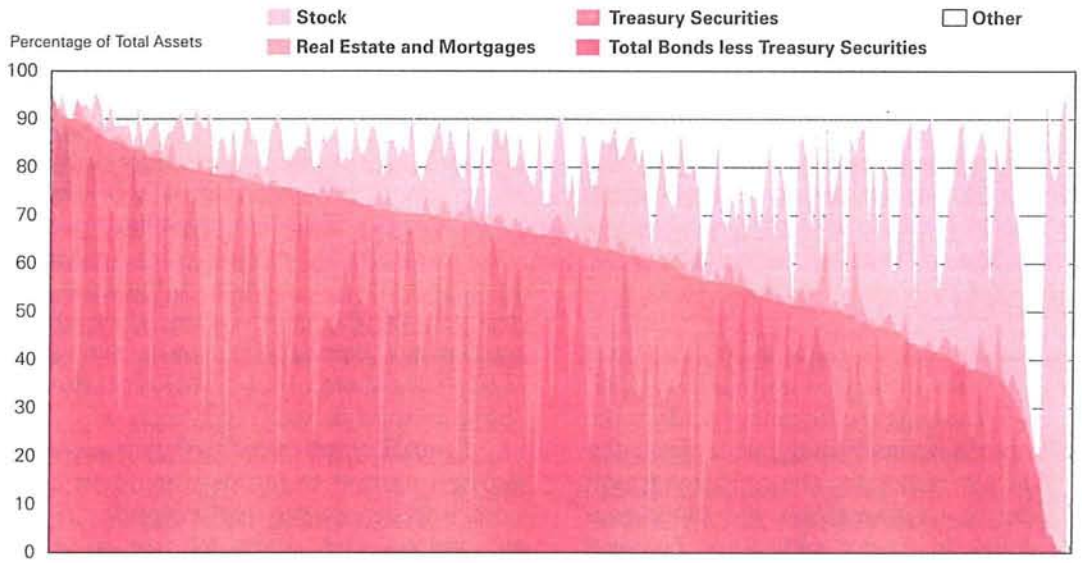
Assets of 100 Largest Life Insurance Companies
Ordered by Percentage of General Account Assets Held in Bonds, End of 1993



Note: These 100 companies hold 80 percent of the industry's general account assets.
Source: National Association of Insurance Commissioners

Figure 1b

Assets of 220 Largest Property and Casualty Insurance Companies
Ordered by Percentage of General Account Assets Held in Bonds, End of 1993



Note: These 220 companies hold 80 percent of the industry's general account assets.
Source: National Association of Insurance Commissioners

able to pay fully its policyholders' claims, especially if the company must liquidate assets at inopportune times when its disappointing returns induce its customers to shift their business out of the company.

The Role of Capital

Shareholders' earnings, which are the difference between insurers' returns from their assets and their net credits to policyholders, represent a financial shock absorber that protects policyholders' investment from the inevitable variations of insurers' return on assets. In perfect financial markets, the value of this margin of protection equals the market value of insurers' capital, the difference between the market value of their assets and that of their contracts with policyholders and other creditors. The more capital per dollar of assets and, consequently, per dollar of reserves, the more secure are policyholders' investments and claims, other things equal.

The increasing diversity of insurers' portfolios poses a challenge for regulators.

The regulation of insurance companies has been shifting away from attempting to control insurers' risks toward setting capital requirements commensurate with insurers' risks in order to protect policyholders' interests. As financial markets have evolved and as economic conditions have changed, fixed rules governing the investments of insurers and the designs of their contracts with policyholders have become dated. Ironically, rules that once made insurance companies safer and sounder now might compromise their security by limiting their ability to diversify adequately in accordance with changing market conditions. At the same time, rules that impeded financial innovations at insurance companies might undermine their role as financial intermediaries and increase the price of insurance. Consequently, the rules governing the activities of insurance companies, like those governing the activities of banks and other financial institutions, have been relaxed since 1950 as regulators increasingly audit risks rather than enforce regulations that delimit the activities of financial institutions. The increasing diversity of insurers' portfolios poses a challenge for regulators: Rules for measuring risk that do not comprehend fully the risks created by the various blends of assets and liabilities often unintentionally subsidize certain types of risk-taking while taxing certain financial strategies that diminish risk.

When insurance companies assume more risk, they must maintain more capital per dollar of assets in order to shelter their policyholders from bearing the consequences. For example, suppose an insurer with a current liability (reserves) of $1 billion to policyholders and $100 million of capital invests this $1.1 billion in a portfolio of relatively safe assets, a portfolio whose value is likely to neither appreciate nor depreciate 5 percent ($55 million) more than expected during the next year. This insurer's capital probably is sufficient to protect the interests of its policyholders, other things equal, through two years of adverse returns. If, on the other hand, the insurer invests in a portfolio of assets whose value is likely to appreciate or depreciate 10 percent ($110 million) more than expected, then its capital probably will not protect policyholders much beyond one year of adverse returns. In the first case, the longer interval of protection not only gives the insurer more time to take defensive actions, it also diminishes the odds of a "fatal draw" —a single year of very low returns occurs more frequently than several years of such returns one after another. The insurer in the second case must maintain at least twice as much capital in order to provide the same protection for policyholders as the insurer in the first case.

An insurance company's risks and, therefore, its need for capital depend on its blend of assets. The risk in a diversified portfolio of assets typically is less than the average risk for each of its assets (Sharpe and Alexander 1990; footnote 3, below). Accordingly, an insurer that purchases risky assets whose returns do not rise or fall rigidly in unison dilutes, to a degree, the risks inherent in each of these assets. When one investment falters, others falter less or may even prosper, thereby diversifying the insurer's risk. The greater the correlations among the returns on an insurance company's investments, other things equal, the greater is its need for capital.

Capital requirements, of course, should take into account more than the risks inherent in an insurance company's assets by considering the risks entailed by its policies and contracts. Just as assets may be blended to reduce risk, so the matching of assets with liabilities also may reduce risk. Insurers expose themselves to substantial risks by financing even safe assets with permanent life insurance or annuity contracts

that guarantee their policyholders a specific rate of return and contracts that allow policyholders to "withdraw" their cash values with little penalty.[2] If yields on alternative investments rise above those implicitly offered in insurers' outstanding contracts, then insurers who invested in long-term bonds run the risk of substantial losses whether they sell assets to meet their customers' withdrawals or they pay their customers a competitive rate of return in order to deter these withdrawals. On the other hand, if insurers invest in short-term securities, they run the risk of substantial losses if interest rates should fall below those either guaranteed in their contracts or offered by their competitors. Insurers assume much less risk either by issuing contracts that impose appropriate penalties on customers who withdraw funds prematurely or by financing shorter-term assets with contracts whose yields vary with market returns.

An insurer's need for capital does not depend simply on the risks inherent in its assets and obligations; it also depends on the frequency of its supervisors' audits, the liquidity of its risky assets, and the power of its supervisors to enforce minimum standards for capital.

Insurers also can manage their risk through financial contracts such as derivatives (financial reinsurance contracts), which do not necessarily appear on their balance sheets, in order to hedge the risks in their balance sheets. For example, insurers holding long-term bonds financed by permanent life or annuity contracts that grant policyholders valuable guarantees can diminish their potential losses by purchasing put options on bond contracts, thereby offsetting the put options they have sold to their policyholders.

An insurer's need for capital does not depend simply on the risks inherent in its assets and obligations; it also depends on the frequency of its supervisors' audits, the liquidity of its risky assets, and the power of its supervisors to enforce minimum standards for capital. If regulators seldom audited insurance companies, then policyholders would require sizable capital-asset ratios to protect their interests. An extremely conservative policy might require that capital equal 100 percent of the value of risky assets. Whenever a company's investment in risky assets exceeds its capital, policyholders' investment potentially is at risk. Although a 100 percent capital requirement certainly would guarantee policyholders' claims, the need for such a severe standard could be relaxed with periodic monitoring and intervention. If, for example, regulators appraised the values of assets quarterly or annually, if insurers could sell their risky assets when necessary with little penalty, and if regulators could require insurers to sell their risky assets when their capital falls below specific bench marks, then this policy of enforcing prompt remedies would allow insurers to maintain much less capital per dollar of risky assets without compromising the interests of policyholders. The more frequent these appraisals by regulators and the more liquid the markets for risky assets, the lower prudent standards for capital may be set.

Balancing Capital Requirements against the Cost of Capital and Regulation

Capital requirements and the implicit, if not explicit, "assurance" resulting from regulators' "seal of approval" allow insurance companies to sell their contracts at more favorable and more stable terms. The challenge for regulators is to set capital requirements that are commensurate with companies' risks, so that the price of this assurance is neither too cheap nor too expensive.

Policyholders' need for the protection provided by capital depends on their ability to assess properly the degree of risk inherent in an insurance company's assets. If all customers understood the risks they were assuming by purchasing an insurer's contract, the need for capital would be moot; policyholders would require returns implicit in their contracts that would compensate them for the risks they bear. Policyholders would assess the appropriate "deposit insurance" premium themselves. Customers, however, rarely understand insurers' risks adequately, and expecting all to assess these risks for themselves would be inordi-

[2] Fixed contract loan rates and guarantees of cash values give policyholders valuable options. Likewise, policyholders also have the option not to renew short-term contracts with life and property and casualty insurers.

nately costly, even if it were feasible. Consequently, regulations have long imposed both minimum standards for the capital of insurers and rules that restrict the types of assets insurers might purchase, in order to protect the investment of policyholders.

To the degree these regulations reassured policyholders, they also benefited insurers by allowing them to sell insurance contracts at more favorable and more stable terms. If, for example, policyholders as relatively uninformed investors generally were too wary of the risks inherent in an insurer's portfolio, then they would require better terms of their insurance contracts, terms that would appear too expensive to insurers. These requirements would be no less volatile than customers' confidence in insurers' investments. Regulations that enforced adequate standards for capital and limited insurers' risks also comforted customers who purchased longer-term contracts: Insurance companies would not assume substantial risks sometime in the future, thereby diminishing the value of their contracts. Accordingly, longer-term contracts could be sold on better terms. Finally, regulations benefited most insurance companies by preventing those companies most inclined to take risks (perhaps to gain a competitive advantage) from undermining policyholders' confidence in the entire insurance industry should these substantial risks produce substantial losses. The success of this regulation, of course, rests on regulators' assessing each company's risks more accurately than the typical policyholder.

Because the managers of insurance companies typically possess better information than other investors about the risks inherent in their investments, the cost of funds for insurers typically rises with capital requirements. Savers not privy to information available to insurance companies generally are less certain than the companies' managers about the potential returns on insurers' assets. For this reason insurance companies, like other financial intermediaries, have profited by transforming the obligations of investors into financial instruments that appeal to savers: Policyholders generally value the guarantees and options embedded in insurers' contracts more than management believes they cost. This advantage, however, becomes a disadvantage when insurers must sell equity to "outsiders," who require a rate of return that "insiders" regard as excessive (Myers 1984; Myers and Majluf 1984).

If regulators assess the risks of insurance companies accurately, then the diligent enforcement of capital requirements that vary with the risks assumed by insurers may allow regulators to strike a good balance between promoting sound intermediaries and fostering efficient intermediation. Capital requirements that are not linked to each company's risks would impose excessive costs on most insurance companies if these requirements were set high enough to protect the interests of the policyholders of companies that assume above-average risks. By linking each company's requirement to its risks, insurers would avoid much of the expense of holding excessive capital, while policyholders and regulators would avoid much of the expense of bearing excessive risk.

Risk-Based Capital Requirements

The prompt enforcement of risk-based capital requirements is tantamount to portfolio insurance for policyholders (Fortune 1995). As the value of an insurer's assets falls relative to that of its liabilities, thereby reducing its capital, regulations compel the insurer either to raise new capital or to reduce its risks commensurately. Should an insurer's capital per dollar of assets fall below a minimum control level, its regulators may take control of the company. For this portfolio insurance to be effective, the risk-based capital requirements ought to take into account the likely costs of selling risky assets in weak markets, and the rules governing regulatory actions ought to allow intervention before a company's capital is likely to be exhausted. For this portfolio insurance to be efficient, both the assets and the contractual obligations of companies ought to be marked according to their "market values"; otherwise, regulators would overestimate the capital for companies whose obligations correspond poorly with their assets and underestimate the capital for companies whose obligations correspond well with their assets.

The prompt enforcement of risk-based capital requirements is tantamount to portfolio insurance for policyholders.

The risk-based capital requirements (RBCR) proposed by the National Association of Insurance Com-

missioners assess the risks inherent in the assets, liabilities, and lines of business of insurance companies (Webb and Lilly 1995; Barth 1995 and 1996; Cummins, Harrington, and Niehaus 1994; Cummins, Harrington, and Klein 1995). The NAIC's proposals also recommend intervention by regulators when insurers' capital does not exceed these requirements.

The RBCR for life insurance companies comprise four components (Table 1). The NAIC's RBCR implicitly assume that the elements of risk within any of these four components are perfectly, positively correlated.[3] The largest of the four components is the asset charge (C1), about two-thirds of risk-based capital (RBC), which comprises assessments for life companies' holdings of bonds, stocks, mortgages, and other investments. For example, assessments for bonds range from no assessment for U. S. Treasury debt to a 30 percent assessment for bonds near or in default; the assessment for the stock of businesses not engaged in insurance is 30 percent; and assessments for mortgages range from 0.1 percent for insured mortgages in good standing to 6 percent for farm and commercial mortgages at least 90 days past due to 20 percent for mortgages in foreclosure.[4] Almost one-fifth of the RBCR for life companies may be attributed to the risks of underwriting various lines of business (C2), risks that arise from inaccurately pricing or estimating morbidity and mortality. These assessments generally are specific proportions of the premiums or net reserves in each of a life company's lines of business. The third component of life insurers' RBCR levies additional assessments on their obligations, assessments that depend on the interest rate risk in their contracts (C3). For example, reserves backing "low risk" contracts (those with cash values that either policyholders cannot withdraw or are subject to market value adjustments) entail a 0.5 to 0.75 percent assessment; reserves against "high risk" contracts (those with guaranteed cash values that policyholders can withdraw without penalty) entail a 2 to 3 percent

Table 1
Capital of Life Insurance Companies

Composition of Risk-Based Capital (RBC)	1994 RBC for 1,540 Life Companies (Millions of Dollars)	Percent of Total 1994 RBC
Total Asset Risk (C1)	55,671	65.9
Bonds After Size Factor	10,342	12.2
Mortgages	6,977	8.3
Preferred/Common Stock	25,205	29.9
Separate Accounts with Guarantees	432	.5
Real Estate	4,966	5.9
Schedule BA Assets	4,855	5.7
Asset Concentration Factor	1,885	2.2
Total Underwriting Risk (C2)	15,788	18.7
Individual & Industrial Life Insurance	4,715	5.6
Group & Credit Life Insurance	2,931	3.5
Individual Health Insurance	3,672	4.3
Group & Credit Health Insurance	7,494	8.9
Premium Stabilization Credit	−3,024	−3.6
Total Liability Risk (C3)	9,970	11.8
Interest Rate Risk-Low	3,643	4.3
Interest Rate Risk-Medium	2,171	2.6
Interest Rate Risk-High	4,157	4.9
Total Business Risk (C4)	3,002	3.6
Total Risk-Based Capital Assessments	84,431	100.0
Risk-Based Capital After Covariance	74,577	88.3

Total Adjusted Capital (TAC)	1994 TAC for 1,540 Life Companies (Millions of Dollars)	Percent of Total 1994 TAC
Capital and Surplus	142,109	79.5
Asset Valuation Reserve	25,200	14.1
Voluntary Investment Reserves	1,329	.7
Dividend Liability	6,518	3.6
Life Subsidiaries' Asset Valuation Reserve	3,444	1.9
Total Adjusted Capital	178,855	100.0

Source: Barth (1995).

[3] The assessment for each asset held by insurers, for example, reflects the volatility (standard deviation, σ_i) of its returns multiplied by the proportion of the portfolio invested in the asset (s_i). The volatility of the return on a portfolio comprising two risky assets is

$$\sqrt{(\sigma_1 s_1)^2 + (\sigma_2 s_2)^2 + 2\rho\sigma_1 s_1 \sigma_2 s_2}.$$

This expression equals the weighted sum of the volatilities of the two assets ($\sigma_1 s_1 + \sigma_2 s_2$) only when the correlation coefficient (ρ) between the assets' risks equals its maximal value, one. For all other values of ρ, the risk of the portfolio is less than the simple sum of the assets' risks.

[4] Some of these assessments can be more or less, depending on the concentration of investments, the number of an insurer's investments (bonds), or an insurer's previous losses (mortgages). The assessment for investments in bonds, for example, depends on the number of different bonds held by an insurer. The assessment for companies holding only 50 bonds is about 2.5 times that for companies holding 1,300 bonds. Yet, this size factor recognizes neither the diversification of credit risks over industries or types of issuer nor the diversification of market risk resulting from the conversion, put, or call options and other features of the bonds.

assessment. The last component of RBCR represents other business risks (C4).

The four components of life insurers' RBCR are combined to obtain the authorized control level risk-based capital (ACRBC). A simple sum would treat these four different types of risk as though they were perfectly, positively correlated. Instead, the formula assumes that the asset and interest rate components are so correlated, while the underwriting component is not correlated with either the asset or the interest rate components of RBC:[5]

$$ACRBC = .5 \times (\sqrt{(C1 + C3)^2 + C2^2} + C4).$$

According to the NAIC's model, when the total adjusted capital (TAC) of a life company (Table 1) is more than 2.5 times its ACRBC, the company is not threatened with regulatory action. When TAC is between 1.5 and 2 times ACRBC, the company must present a plan to increase this ratio, to be approved and monitored by its insurance commissioner. When TAC is between 1.5 and 1 times ACRBC, the commissioner also may issue corrective orders to the company. When TAC is less than ACRBC, the commissioner may take the necessary actions to rehabilitate the company, including seizure or liquidation.

Because the underwriting risks of property and casualty companies are commensurately greater than those of life companies, the assessments for these risks represent almost two-thirds of the RBCR for property and casualty companies (Table 2). These underwriting charges for each company—which comprise assessments for net premiums received by line of business, for reserves against future claims, and for loss adjustments—depend on the industry's previous experience in pricing its claims and the company's previous experience relative to that of the industry. Charges for property and casualty insurers' investments in bonds and equity, which are similar to those for life insurers, represent only about one-seventh of the RBCR for property and casualty insurers. The remaining components of RBCR for property and casualty insurers include charges for reinsurance and other receivables as well as for guarantees and other liabilities that do not appear on their balance sheets.

According to the NAIC's formula, the ACRBC for property and casualty companies assumes that the risks for investments in equity and fixed-income securities, credit risks, and underwriting risks are not correlated, while the sum of these risks is perfectly, positively correlated with the asset risk for affiliates:

$$ACRBC = .5 \times (R0 + \sqrt{R1^2 + R2^2 + R3^2 + R4^2 + R5^2}).$$

The NAIC's proposal requires the ratio of the total adjusted capital to ACRBC for property and casualty companies to pass the same tests that are applied to life insurers.

Table 2
Capital of Property and Casualty Insurance Companies

Composition of Risk-Based Capital (RBC)	1994 RBC for 2,244 P/C Companies (Millions of Dollars)	Percent of Total 1994 RBC
Asset Risk—Affiliates (R0)	24,038	18.1
Asset Risk—Fixed Income (R1)	2,633	2.0
Bonds of Affiliated Insurers	273	.2
Other Bonds	1,363	1.0
Bond Size Factor	560	.4
Asset Concentration Factor	210	.2
Asset Risk—Equity (R2)	16,944	12.8
Stock of Affiliates	2,327	1.8
Other Stock	8,852	6.7
Asset Concentration Factor	3,342	2.5
Asset Risk—Credit (R3)	5,632	4.2
Underwriting Risk—Reserves (R4)	52,955	39.8
Underwriting Risk—Written Premiums (R5)	30,693	23.1
Total Risk-Based Capital Assessments	132,894	100.0
Risk-Based Capital After Covariance	94,907	71.4

Total Adjusted Capital (TAC)	1994 TAC for 2,244 P/C Companies (Millions of Dollars)	Percent of Total 1994 TAC
Capital and Surplus	204,808	98.4
Life Subsidiaries' Asset Valuation Reserve	2,971	1.4
Life Subsidiaries' Voluntary Investment Reserves	859	.4
Life Subsidiaries' Dividend Liability	367	.2
Non-Tabular Discount	−864	(.4)
Total Adjusted Capital	208,133	100.0

Source: Barth (1996).

[5] Business risk, according to the formula, is perfectly, positively correlated with the sum of the first three components of risk.

II. Problems with Current Risk-Based Capital Requirements

Risk-based capital requirements for insurance companies, like those for banks and other intermediaries, are new and still experimental. As techniques for measuring risk and regulating intermediaries according to their risks continue to evolve, future capital requirements should satisfy several deficiencies apparent in current regulations. Some of these deficiencies may be remedied by improving the design of prevailing rules but, whatever the design, some may be intrinsic to the strategy itself (Merton 1995).

Risk-based capital requirements for insurance companies, like those for banks and other intermediaries, are new and still experimental.

The current regulations take a narrow view of an insurance company's risks and capital. Risk, according to these regulations, essentially is defined by the properties of each class of asset and each class of obligation. Because the regulations admit only the correlation coefficients of zero or one among these risks, RBCR essentially make little allowance for the diversification of investments or for matched books. Moreover, measures of capital, the gauge of insurers' capacity for bearing risk, are defined by combinations of market and book values, not by a consistent accounting framework.

These apparent deficiencies in existing regulations may reflect more than problems with technical details. If markets for financial instruments are not perfect, as assumed in the financial theory behind RBCR, then coherent measures of risk and capital may be elusive. Therefore, if the assets and liabilities of insurers are not always priced efficiently in liquid markets, the strategy of promptly enforcing any capital requirement at times may undermine, rather than foster, safe and sound financial institutions.

The Concept of Risk in RBC

Prevailing RBC assessments for the risk in assets (C1) depend little on the diversification of an insurance company's portfolio of assets. The risk of financial instruments is defined by the share of the volatility of their returns that cannot be offset or fully diluted when they are combined with other investments. Accordingly, the risk inherent in any stock, bond, or loan depends on the division of an investor's assets among other investments. Yet, the RBC assessment for an investment in IBM shares, for example, is much the same whether an insurer's equity portfolio comprises the S&P 500 or only technology stocks, whether an insurer "overweights" or "underweights" equity among its assets.[6] Furthermore, by adding the assessments for investments in stocks, bonds, mortgages, and other assets, the RBC rules that define asset charges essentially assume that the returns on all assets are perfectly, positively correlated. The rules give insurers comparatively little credit for hedging their investments by holding assets whose returns either may tend to move in opposite directions or, at least, may not tend to rise and fall together very strongly.

The RBC assessments also do not change as the risks inherent in and among the assets change. The variances and covariances of returns in the past have varied with the phases of the business cycle, the rate of inflation, changes in relative prices (oil shocks, changing exchange rates, monetary policy), or the magnitude and composition of technological innovations in the economy. The variances and covariances among returns also depend on the length of time that insurers hold their assets and obligations.

As a consequence of this inflexible pricing of risk, prevailing RBC regulations are not so much a new way of measuring and controlling insurers' risks as they are an adjustment of the prices embedded in those controls. Regulations formerly proscribed certain investments by imposing prohibitive costs on insurers who might have considered buying these "risky" assets. RBC regulations reduce, but do not eliminate, these costs. For example, RBC rules assess an insurer purchasing IBM shares a 30 percent asset charge, regardless of the insurer's efforts to hedge its investment in IBM or diversify its portfolio; consequently, this charge encourages the insurer to shun

[6] Because RBCR include an asset concentration factor, the assessments are lower for portfolios that spread their investments in equity more evenly among more stocks or for portfolios that hold more assets other than stocks. But, neither this concentration factor nor the size factor (see footnote 4) takes into account the correlations among the returns on these assets in order to measure properly the degree to which a portfolio has reduced its risk by diversifying its investments.

equity in favor of other investments that entail smaller costs.

Contrary to the strategy incorporated in RBC regulations, the risks inherent in liabilities cannot be measured and controlled apart from asset risk. Interest rate assessments do not depend on the composition of an insurer's assets or the nature of its other obligations. The assessment for a company that issues many short-term policies or contracts is less than the assessment for a company that issues many long-term contracts that make specific guarantees to policyholders. If, however, the first company invested the proceeds of its contracts entirely in long-term bonds (a strategy of leverage similar to that of Orange County recently or much of the savings and loan industry before the mid-1980s), its risk could be many times that of the second company. The second company's contracts, on the other hand, might bear little risk if it had purchased suitable structured notes, swaps, or interest rate options. Ironically, this financial reinsurance entails additional RBC assessments even if companies use these instruments to reduce their risk. Despite the differences in their liabilities, both companies essentially could eliminate their interest rate risk by matching closely the terms and features of their assets to the terms and features of their liabilities. Such matched books also would reduce substantially their asset risk.

The Measurement of Capital

When financial markets are perfect, an insurance company's capital, the difference between the market values of its assets and its liabilities, measures its capacity for protecting the investments of its policyholders and other creditors. Yet, prevailing accounting standards value some assets of insurers according to their market values, other assets according to their book values (reflecting their face values or acquisition costs plus any necessary adjustments), and most liabilities according to their guaranteed face values. This mixture of accounting techniques can produce biases in the measurement of an insurer's capital, biases that would undermine the value of RBC standards even if these standards properly reflected the company's risks (Carey 1995). Companies that actually lack sufficient capital, for example, might meet their standards if book values overstated the values of certain risky investments; conversely, other companies possessing sufficient capital might fail to meet their standards if book values overstated the values of their liabilities.

Although RBC standards, at least in principle, recognize that the values of the assets and liabilities of insurance companies can vary with economic conditions, the measure of the capital that is to be compared to these standards may not reflect the changing values of these financial instruments. For example, current regulations do not recognize the "capital" that insurers carry by matching their assets with their obligations. Should interest rates rise abruptly, the prices of equities and bonds would fall, thereby depressing the value of assets and the capital of insurers. Yet, for the same reason the prices of these assets fall, the "prices" of longer-term, market-priced insurance or annuity contracts fall as well. Insurers that have issued these contracts are credited with lower RBC requirements, but this may be little solace if, for want of market value accounting for their liabilities, they are given no credit for preserving their capital as the value of their contracts falls in concert with that of their assets. Just as the capital of these companies would be understated when interest rates are rising, their capital would be overstated when rates are falling. Similarly, the reporting of mortgages and other investments at book values also distorts the measurement of capital.

Market-value accounting for all assets and liabilities is not a panacea, because it too may misrepresent the capital of insurance companies.

Market-value accounting for all assets and liabilities is not a panacea, however, because it too may misrepresent the capital of insurance companies. To the degree insurers hold assets that do not trade in perfect markets, they also may hold "capital" in the form of excess returns on these proprietary assets. Insurers, like other financial intermediaries, profit by investing in assets that are not very familiar within public financial markets and, therefore, are not priced efficiently by those markets. For providing this service and bearing the attendant risks, insurers may earn, on average, an extra margin in their returns over time, a margin that outside investors may not recognize consistently. If outside investors too often are wary of the value of these proprietary assets and the markets for these assets too often are shallow or illiquid, market

prices in these circumstances can be biased estimates of their values.[7] As the optimism of outsiders rises, prices of these assets may nearly meet or exceed proprietary valuations for a time, only to fall below proprietary valuations when this optimism subsequently ebbs. This potential volatility of prices for these assets induces a commensurate volatility of insurers' capital with market accounting. Insurance companies in the United States and Japan as well as banks in Texas, New England, Scandinavia, and Japan, for example, possessed more than adequate protection when the value of the enterprises and real estate backing their assets was very great, but their capital eroded quickly when the prices of these assets collapsed.

The Prompt Enforcement of RBCR Is Not the Same as Portfolio Insurance

Policies for enforcing capital requirements that promote sound insurance companies in some circumstances might fail to do so in other circumstances. For example, the prompt enforcement of RBCR is a conservative policy when the markets for financial instruments are liquid. Yet, this policy tends to weaken insurers when outsiders are most skeptical of the returns on their risky assets and the prices of these assets understate their value significantly.

If risky assets were priced efficiently, tomorrow's news would be no more likely to increase the value of these assets more than expected than to decrease their value more than expected (Cootner 1964; Merton 1990). Consequently, should substantial losses reduce a company's capital per dollar of risky assets, the chance that surprisingly high returns on these assets subsequently would increase the capital-asset ratio is little greater than the chance that surprisingly low returns would reduce this ratio. As the value of a company's assets falls relative to its obligations, the odds of insolvency increase, and prudent supervisors would require that safe assets supplant risky assets, thereby reducing the company's risk to correspond to its diminished capital.

Nonetheless, the prompt enforcement of capital requirements is not necessarily a conservative policy when markets are not liquid. If proprietary assets are not priced efficiently, their values may not follow random walks. Instead, the prices of these assets may revert to trends: Once a price falls below its proprietary valuation, the odds of its returning increase with time, while the odds of its falling further diminish. The prompt enforcement of capital requirements may even magnify the degree to which the prices of these assets diverge from trends. If, for example, an insurer must sell risky assets in order to restore its ratio of capital to risky assets after the prices of these assets subside in the opinions of outsiders, then these prices will fall further in illiquid markets.[8] After the prices of risky assets fall substantially the chance of redeeming capital gains increases with time, while the chance of commensurate losses diminishes. Therefore, when the value of an insurer's assets approaches that of its obligations and its liabilities are of sufficiently long duration, its expected losses due to insolvency may be low compared to the expected gains from retaining these assets.

Suppose an insurer attempts to maintain a ratio of capital to assets of 10 percent, while investing 40 percent of its assets in proprietary investments, 60 percent in safe assets. Because policyholders believe the insurer is regulated adequately, the yield on these accounts equals the yield on safe assets. The prices of proprietary investments follow a smoothed random walk: A below-average return on these assets creates no expectation of compensating above-average returns subsequently. When favorable earnings increase its capital per dollar of assets, the insurer sells more contracts, investing the funds as required to maintain the 3:2 ratio between its safe and risky assets. When poor earnings reduce its capital per dollar of assets, the insurer sells no new contracts and acquires no new risky debt. The capital of this insurer approaches zero, on average, nearly twice every one hundred years (Figure 2a). When the insurer practices portfolio insurance, selling risky assets as required in order to prevent the ratio of risky assets to capital from exceeding 4, then its capital approaches zero less than once every century (Figure 2b).

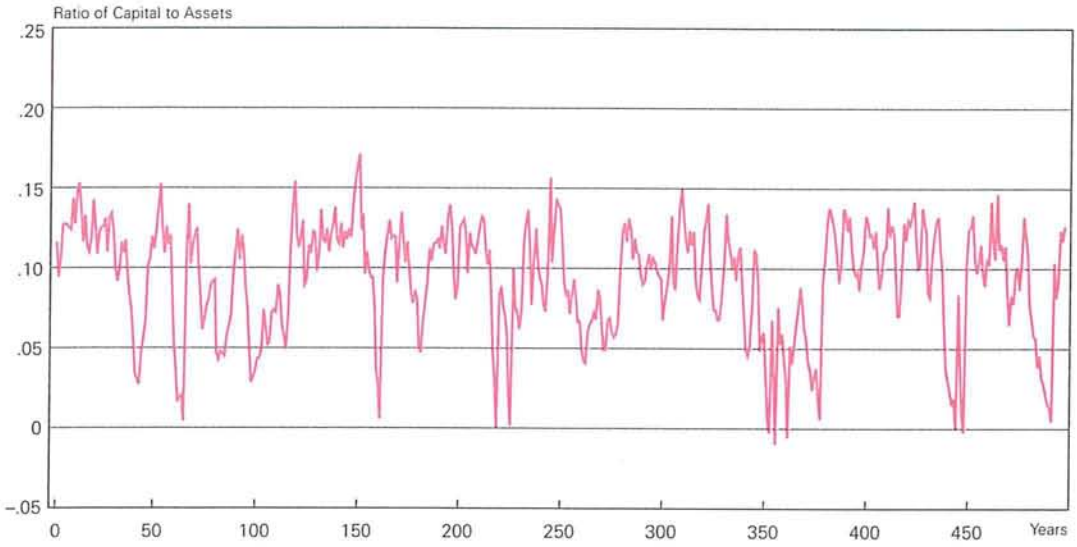
If the values of proprietary assets tend to return to trend—a run of below-average returns increases the odds of earning above-average returns—the capital-to-asset ratio almost never approaches zero with the

[7] When no true prices are quoted in markets for assets, supervisors often resort to prices of comparable assets, prices derived from models, or book values of assets net of estimates of losses (due to default or workouts).

[8] Disposing of risky assets is frequently more economical than selling new equity. When an insurer has reported losses great enough to impair its capital, wary outsiders are not likely to value its equity very greatly. If the insurer sold those assets that are most familiar to outside investors, it would only increase the proportion of its liabilities backed by questionable risky assets. If, however, outsiders discounted the value of risky assets too greatly, so that selling these assets entailed substantial losses and the insurer's capital were sufficiently near insolvency, management would need to issue new equity.

Figure 2a

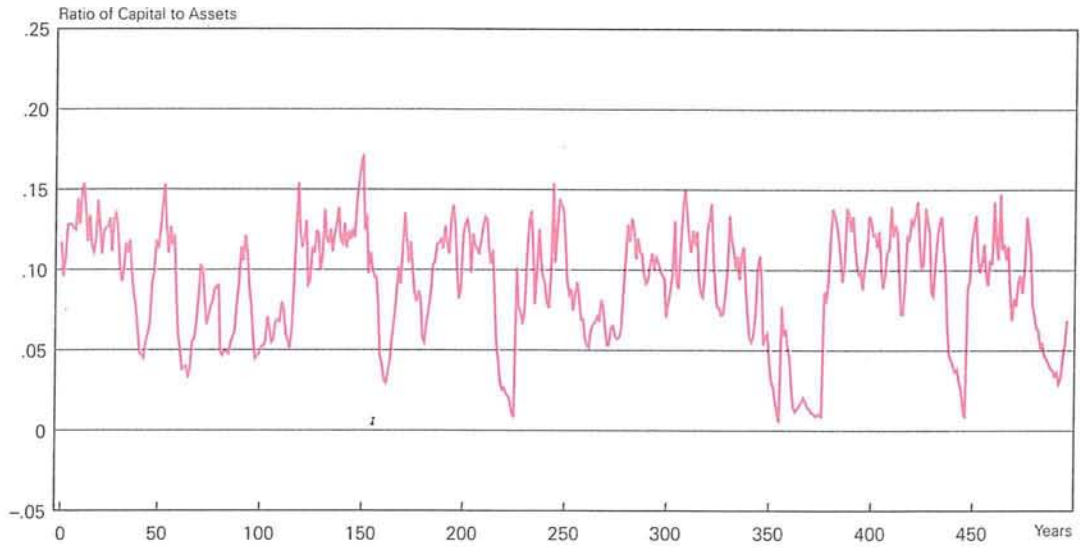
Capital Ratios When Prices of Assets Follow a Random Walk



See Appendix.

Figure 2b

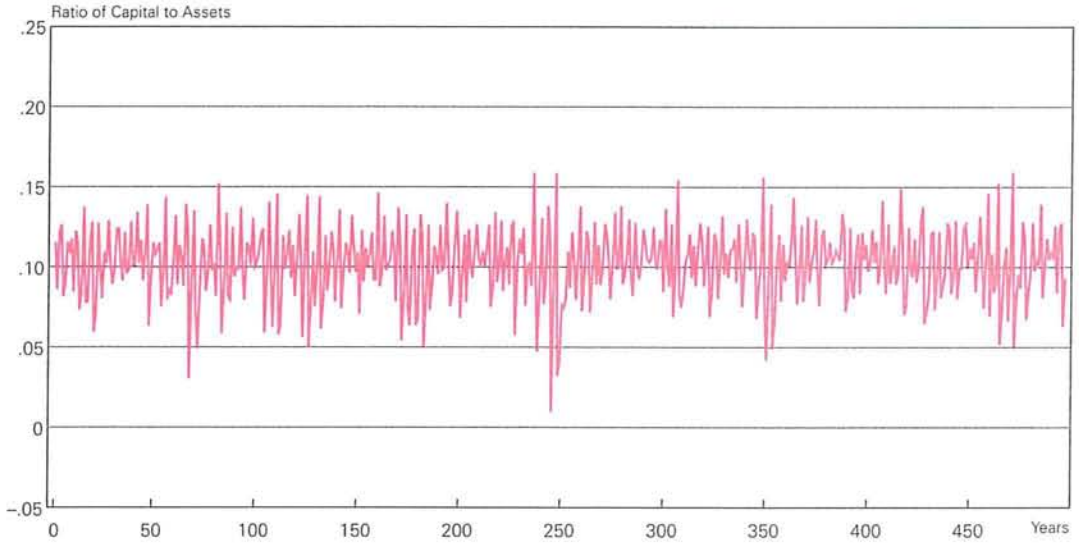
Capital Ratios When Prices of Assets Follow a Random Walk and Capital Requirements Are Promptly Enforced



See Appendix.

Figure 2c

Capital Ratios When Prices of Assets Revert to Trend



See Appendix.

Figure 2d

Capital Ratios When Prices of Assets Revert to Trend and Capital Requirements Are Promptly Enforced

See Appendix.

investment strategy described in the first simulation (compare Figure 2c to 2a), even though the annual volatility of the rate of return on proprietary assets is greater than in the first case. If, in this last instance, the insurer practices portfolio insurance by selling some of its risky assets after their values decline and if the disposal of these assets temporarily reduces their prices by an additional 10 percent, then the insurer's average capital-asset ratio (Figure 2d) falls and becomes more volatile. Consequently, the insurer's capital approaches zero more frequently, about once every century, when it sells its risky assets at distressed prices in order to meet its capital requirements. Furthermore, this policy of promptly enforcing capital requirements induces a clear credit cycle: The lending capacity of the insurer, as reflected in its capital per dollar of assets, falls further and remains depressed longer in this last case than it did in the former.

This example does not imply that financial institutions that hold illiquid assets always ought to enjoy the protection of forbearance. To be sure, the falling prices of mortgages and real estate during the commercial real estate slump of the 1990s greatly diminished the capital of many banks, a loss that often was exaggerated by their need to sell these investments. Insurance companies, on the other hand, generally survived the experience with less duress by not marking these assets according to market prices and for not having to sell them at bargain prices. Whereas patience was appropriate for insurers that financed their investments in real estate with longer-term policies that included an adequate pricing of policyholders' put options, this patience was less suitable for insurers that, like banks, had assumed more risk by financing these investments with short-term contracts or contracts that gave policyholders generous guarantees.

III. Beyond Current Risk-Based Capital Requirements

According to the current design of risk-based capital requirements, insurance companies should hold capital in proportion to their investment in assets that are designated risky, but these standards measure neither the protection for policyholders embedded in insurers' portfolios nor the rate at which this protection might change with economic conditions (Grenadier and Hall 1995). Furthermore, to the degree insurance companies hold assets that are not priced efficiently in public markets, the prompt enforcement of these capital standards might undermine, rather than foster, the safety and soundness of insurance companies. In any case, the difference between the values of an insurance company's assets and liabilities, whether these values are market or book, does not measure properly its "capital"—the protection inherent in its stream of net income—when its assets and obligations are not priced efficiently in liquid markets.

By pricing risk inaccurately, existing risk-based capital requirements may diminish the efficiency of financial markets by discouraging insurance companies from holding those assets that are not very familiar in public markets and those longer-term assets that are designated as most risky. The need to justify the valuation of assets and the potential need to sell risky assets in times of duress encourage an insurer to shun investments whose value to the company depends too greatly on the company's proprietary information. Insurers, therefore, withdraw to a degree from their role as financial intermediaries as they increasingly favor liquid, familiar assets. Insurers also cede financial intermediation to others as they alleviate their capital requirements by promoting business linked to separate accounts or mutual funds wherein policyholders bear more of the risks of the investments backing their contracts. This "mutual fund" strategy currently appeals to many customers who, as a result of the comparatively great yields generated by stocks and bonds since 1980, often expect to earn generous returns while bearing commensurately little risk.

Risk managers and regulators might use the models behind value-at-risk calculations to isolate those economic conditions that threaten the solvency of insurance companies.

Alternative standards for the capital of insurance companies ultimately might diminish such disintermediation by assessing the influence of economic conditions on insurers' earnings and cash flows, instead of "taxing" various assets and liabilities. The New York State Insurance Department, for example, requires a cash flow test for certain life insurance and

annuity contracts in order to assess the risks in these contracts. If a company's losses would threaten its solvency should interest rates rise 300 basis points or more, the company's directors and regulators might encourage the company to issue more equity or to alter the composition of its assets and liabilities in order to mitigate this threat. Such tests implicitly weigh the consequences of different portfolio strategies, including those related to: (1) the options assumed by insurers, including those embedded in their assets and liabilities; (2) the mismatches between long and short commitments at various maturities; (3) the correlation of returns among assets and liabilities; and (4) the possibility that the prices of some assets collapse and their maturities increase for want of dependable markets. These tests should be dynamic, incorporating managements' responses to changing conditions and covering intervals of time sufficiently long to encompass the full consequences of these changing conditions.

Some financial institutions currently are using models of "value-at-risk" in order to assess the adequacy of their capital. These strategies, using management's assessment of the likelihood of potential economic conditions, calculate the odds of an institution losing its capital. In principle, an insurer could avoid financial strategies for which the probability of its insolvency exceeded its tolerance. Nevertheless, the insurers still must contend with the risk that changing economic conditions entail surprisingly sharp changes in the prices of assets as well as in the customary covariances among the returns on these assets.

Risk managers and regulators might use the models behind value-at-risk calculations to isolate those economic conditions that threaten the solvency of insurance companies. A conservative policy might require that insurers adopt financial strategies that limit their maximum losses for all "feasible" conditions, a kind of minimax strategy. Each insurer's need for "capital" would vary according to the mix of its assets and liabilities. If, for example, a company is vulnerable to a specific shift of the yield curve, regulators might counsel it to alter its investments, to purchase hedges, or to sell more equity to insure that its earnings remained sufficiently great compared to its obligations should such a shift occur, even if this event were not regarded as a very likely threat.[9] This version of risk-based capital requirements might reveal best the risks that insurance companies were bearing and, when necessary, might tie their need for capital most directly to these risks, rather than to their commitments to individual assets and liabilities.

[9] Of course, the lower the odds of such an event, the more cheaply the company may purchase insurance for its potential loss. The most economical insurance may take the form of new investments or derivatives contracts designed to cover these specific risks rather than the umbrella insurance policy entailed by raising new capital.

Appendix

Figure 2

Panel a: An insurer holds risky and safe assets, financed by equity and "contracts." The expected return on risky assets, $E(r_t)$, is 10 percent annually; the standard deviation of this return, $\sigma(\varepsilon_t)$, is 6 percent annually; and the correlation coefficient between annual returns (a first-order Markov process) is 60 percent:

$$r_t = .10 + \varepsilon_t$$

$$\varepsilon_t = .6\varepsilon_{t-1} + \eta_t$$

$$\eta_t \sim N(0, .06^2(1 - .6^2)).$$

The return on the insurer's safe assets and the return that the insurer pays on its contracts is 7 percent. The values of risky and safe assets increase according to their returns and any new investments in these assets, Δ^r and Δ^s; likewise, the value of its contracts increases as a result of crediting interest and new inflows, Δ:

$$V_t^r = V_{t-1}^r(1 + r_t) + \Delta_t^r$$

$$V_t^s = V_{t-1}^s(1.07) + \Delta_t^s$$

$$L_t = L_{t-1}(1.07) + \Delta_t.$$

The capital of the insurer, C, is the difference between the value of its assets and the value of its contracts, L. When its capital per dollar of assets the previous year exceeds its target of 10 percent, the insurer issues new contracts; otherwise, Δ is zero. If the insurer's risky assets are less than 4 times its capital, the insurer purchases more risky assets in order to maintain the ratio of 2 dollars of risky assets for every 3 dollars of safe assets; otherwise, Δ^r is zero:

$$\Delta_t = \max([10C_{t-1} - (V_{t-1}^r + V_{t-1}^s)], 0)$$

$$\Delta_t^r = \max([4C_t - V_{t-1}^r(1 + r_t)], 0)$$

$$\Delta_t^s = \Delta_t - \Delta_t^r.$$

When the insurer's capital falls below 0.5 percent, it "fails," and its capital is restored to 10 percent. In the simulation

shown in the graph, the insurer fails 11 times, its mean capital-asset ratio is 9.5 percent, and the annual standard deviation of this ratio is 3.5 percent.

Panel b: The assumptions are the same as those for the previous panel, except that the insurer sells risky assets in order to maintain only 4 dollars of risky assets per dollar of capital when this ratio exceeds 4:

$$\Delta_t^r = 4C_t - V_{t-1}^r(1 + r_t).$$

In the simulation shown in the graph, the insurer fails 4 times, its mean capital-asset ratio is 9.1 percent, and the annual standard deviation of this ratio is 3.5 percent.

Panel c: The assumptions are the same as those for the first panel, except that the value of risky assets tends to revert to a trend:

$$trend_t = V_0^r(1.1)^t$$

$$r_t = .10 + \varepsilon_t$$

$$\varepsilon_t = .6\varepsilon_{t-1} - 2\log(V_{t-1}^r - trend_{t-1}) + \eta_t$$

$$\eta_t \sim N(0, 0.5^2(1 - .6^2))$$

$$V_t^r = V_{t-1}^r(1+\mathrm{r_t}).$$

The standard deviation of annual returns behind the simulation shown in the graph is 7.5 percent. Yet, because of the tendency of the value of risky assets to revert to trend, the insurer does not fail during this simulation, its mean capital-asset ratio is 10.3 percent, and the annual standard deviation of this ratio is 2.2 percent.

Panel d: The assumptions are the same as those for the previous panel, except that the insurer sells risky assets in order to maintain only 4 dollars of risky assets per dollar of capital when this ratio exceeds 4, and that this sale entails transactions costs equal to 10 percent of the value of the risky assets that are sold.

In this simulation, the insurer fails 5 times, its mean capital-asset ratio is 9.1 percent, and the annual standard deviation of this ratio is 3.6 percent.

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Can Demand Elasticities Explain Sticky Credit Card Rates?

Joanna Stavins

Economist, Federal Reserve Bank of Boston.

It has long been recognized that interest rates charged on credit card loans are sticky (that is, they remain high even when the cost of funds drops). Although some studies have blamed market power by issuing banks for the persistently high rates,[1] the credit card market is relatively unconcentrated, with hundreds of issuers nationwide. The explanation for the sticky rates is more likely, therefore, to lie on the demand side. Since consumers could minimize their cost of credit by borrowing at the lowest possible rate,[2] one would expect banks to drop their rates to attract customers in the competitive market. Yet issuing banks do not appear to be behaving in this way. Do banks maintain high rates because customers' demand for credit card loans does not respond to changes in the rates they charge (that is, because demand for credit cards is inelastic with respect to the interest rates)? Do consumers indeed borrow at high interest rates because they are irrational, as Ausubel (1991) suggested?

Several theories purport to explain credit card rate stickiness.[3] Although some studies have speculated whether demand for credit cards loans is responsive to interest rates, the only information about demand elasticities comes from consumer survey results.[4] According to evidence presented in Ausubel (1991), however, consumer survey results consistently underestimate how much consumers actually borrow. When the results of consumer surveys are compared to bank data, it turns out that consumers borrow more and repay less than they report. Therefore, evidence about demand elasticities should come from bank data, yet no study has explicitly estimated demand elasticities for credit card loans with respect to the interest rates charged. Using panel data from credit card plans offered by banks, this study estimates consumers' sensitivity to the various attributes of credit card plans: interest rates, annual fees, grace periods, finance charges, and additional enhancements. In the past, regulatory agencies and research economists have

focused their analyses of the credit card market almost exclusively on the annual percentage rate of interest (APR). However, customers may be more responsive to other characteristics of the plans. It is worthwhile to find out whether the careful scrutiny the credit card rates have received over the years should be directed at other attributes as well.

Consumers have more credit card plan options today than ever before. Most credit card plans are offered nationwide, and abundant information about them arrives in every day's mail. Each plan is composed of many attributes. Are consumers more likely to borrow at a lower interest rate, pay a lower annual fee, or choose more "bells and whistles"?[5] Consumers may opt for high-APR plans because of their inelastic demand or because those plans compensate them with other features, such as low fees. This article approaches the sticky interest rate puzzle by estimating consumers' demand responsiveness to the various features of credit card plans.

This article approaches the puzzle of sticky interest rates on credit card loans by estimating consumers' demand responsiveness to the various features of credit card plans.

[1] Calem and Mester (1995); Shaffer (1994).

[2] Approximately two-thirds of credit card holders carry debt on their cards (*American Banker*, 1/4/96, p. 12).

[3] Ausubel (1991); Calem and Mester (1995); Shaffer (1994); Woolley (1988). See also Canner and Luckett (1992) for a review of that literature.

[4] For example, a 1986 survey of cardholders by Payment Systems, Inc. found that customers would apply for a credit card with a lower annual interest rate if offered. In a recent survey by *American Banker*, 23 percent of the respondents said that they would switch to another plan for a 1 percent drop in the interest rate (*American Banker*, 11/1/95, p. 12). Canner and Luckett (1992) discuss consumer sensitivity to interest rates based on consumer survey results, but do not provide any numbers.

[5] For example, some press articles have speculated that consumers may be responsive to enhancements: "Americans are back in love with their credit cards—and not just because the economy has improved . . . by offering airline miles or rebates, [credit card issuers] are providing more incentives to use credit cards in place of cash or checks. . . . Consumers love rebate products" (*BusinessWeek*, 12/12/94, p. 42).

The first section describes the data used in the analysis. Section II addresses the question of whether credit card users are rational. Section III sets up the specification used in this paper, while the following section presents estimation results. Section V examines how a bank's size affects the credit card rates it charges and the demand elasticity it faces. The final section offers a summary and conclusions. The results show that banks face an adverse selection problem: Lowering the APR would attract risky customers or induce existing customers to borrow more than they can handle. As a result, delinquent loans rise at a significantly higher rate than that of loans in general. This induces banks to maintain high interest rates. The adverse selection hypothesis is further supported by the finding that banks' income from credit card fees and interest increases with APR.

I. The Data

This study uses data from a survey on the Terms of Credit Card Plans (TCCP), collected semiannually by the Federal Reserve Board from approximately 200 of the largest issuers of bank credit cards. The survey was conducted each January and July during the 1990–95 period. Smaller banks are not included in the sample. Although they may offer systematically different terms of credit card plans, the sampled banks issue the majority of outstanding credit.[6]

The data include characteristics of each plan, such as annual percentage rates (APR), annual fees, grace periods, minimum finance charges, late payment charges, cash advance fees, and over-the-limit fees, as well as indicators showing whether the plan had additional "enhancements," such as automobile insurance, travel discounts, extended warranty, and the like. The data set was merged with information from bank financial statements filed with the Federal Deposit Insurance Corporation. These Consolidated Reports of Condition and Income (Call Reports) include each bank's deposits and assets, as well as outstanding credit card loans and income from credit card interest and fees. The Call Report data are collected quarterly. Data from March Call Reports were merged with the January TCCP data, and data from September Call Reports were merged with the July

[6] According to the recent American Bankers Association annual survey, the outstanding credit card credit totaled $257 billion at the end of 1994 (*American Banker*, 1/4/96, p. 12). The TCCP sample issuers' outstanding credit amounted to $246 billion in January of 1995, or about 96 percent of the total.

TCCP data.[7] Panel data constructed from information on the majority of credit card banks over the period of six years permit analysis of customers' sensitivity to features of credit card plans. Table 1 (below) provides descriptive statistics and definitions of the major variables.

Although interest rates on credit card plans have remained high relative to other rates of interest (Figure 1), the average APR has declined over the past few years. Several issuers have also eliminated annual fees, although the average annual fee was approximately constant until 1994.[8] Figure 2 shows changes in the average APR and annual fee during the 1990–95 period. However, these changes do not necessarily mean that credit has gotten cheaper. As can be seen in Figure 3, issuers have been raising other charges, such as late payment fees and over-the-limit fees. The evidence also shows some regional differences among credit card plans. In particular, New England banks have been offering lower rates of interest, but charge higher annual fees than banks in the rest of the country (Figure 4).[9]

Selecting a credit card has therefore become more complicated over time: Each plan is composed of a vector of various charges and rewards, and more variation exists among them now than in the past. As Figure 5 shows, the variance in both APR and annual fees has increased, even though the sample of issuers has remained fairly stable. The increase in the variance of APR was partly caused by a higher fraction of credit card plans with variable rates of interest (Figure 6). The next two sections of the paper examine the sensitivity of customers to the various options.

II. Are Consumers Rational?

One explanation for high credit card interest rates is consumer irrationality (Ausubel 1991). According to that view, consumers typically do not intend to borrow on their credit cards but end up doing so anyway. These "irrational" consumers presumably do not take APR into account when selecting a plan, since they do not intend to carry any debt. Banks therefore have no incentive to lower their rates. On the other hand, if consumers behave consistently with their intentions, they are likely to put significant weight on APR when deciding which credit card plan to adopt.[10] Under the "rational" scenario, convenience customers (customers who repay their balance in full) would be more likely to choose a plan with a low fee, a long grace period, and many enhancements, but ignore APR. At the same time, revolvers (customers who carry a balance on their cards) would choose low-APR plans, but pay less attention to the other attributes. By the same token, under the rational scenario low-APR plans would have a relatively higher fraction of overdue loans (those on which customers failed to make minimum payments), while low-fee or high-enhancement plans would have a relatively lower fraction of overdue loans.

The difference in delinquency rates between no-fee and positive fee plans could indicate either that higher charges create high rates of delinquency or, more likely, that banks offering attractive terms are more selective in their customer approval process.

Casual observation of the data yields no support for the rational scenario. The correlation coefficient between APR and the fraction of overdue loans[11] is *positive* (0.23), contrary to the above. High-APR plans have higher rates of delinquency than low-APR plans: The average delinquency rate for plans with APR above 17 percent (the mean) is 3.4 percent, compared to 2.6 percent for plans with APR below 17 percent

[7] Quarterly flow variables were adjusted to correspond to the appropriate six-month period.

[8] Carlton and Frankel (1995) show that the average annual fee charged by Visa issuers declined throughout the 1990–95 period (p. 44). Since this study uses a much larger sample than Carlton and Frankel used, the results of this study are more likely to be accurate.

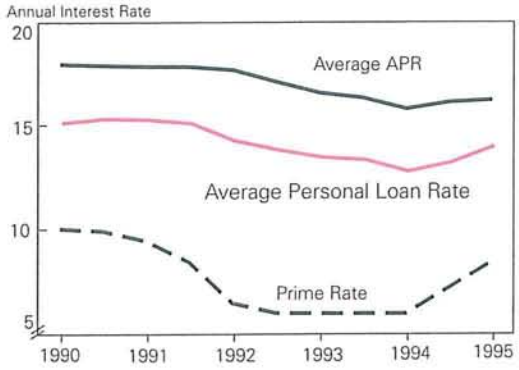
[9] There is no evidence for a trade-off between APR and annual fee for the overall sample. The correlation coefficient between the two variables is not statistically significantly different from 0.

[10] Even if the one-third of credit card users who pay their balance in full ignored the interest rates on their credit cards, the rate of interest would be a significant factor in credit card borrowing.

[11] The amount of overdue loans is from the Call Report and is defined as the amount of credit card loans on which customers have failed to make minimum payments. Although the variable does not include all the loans that accrue interest, the two measures can be expected to be correlated.

Figure 1

APR has remained above the prime rate and rates on personal loans.



(the difference is statistically significant at the 5 percent level). However, no-fee credit card plans tend to have lower delinquency rates than plans with nonzero fees, consistent with the above: The average delinquency rate for no-fee plans is 2.4 percent, as compared to 3.2 percent for plans with positive annual fees (again, the difference is statistically significant at the 5 percent level).

The differences in delinquency rates could indicate either that higher charges create high rates of delinquency or, more likely, that banks offering attractive terms are more selective in their customer approval process. Low-APR banks seem to screen their customers more carefully than high-APR banks to avoid high-risk cardholders. High-APR banks have less of an incentive to screen than low-APR banks do: Since their profit margin is higher, they can afford higher losses. Even if consumers were perfectly rational, low-APR banks might turn down revolvers who intend to borrow (because of their past credit record, for example) and force them to get high-APR cards. If banks screened their customers correctly, then low-APR banks would end up with a lower charge-off rate (a measure of losses) than high-APR banks. The data support this theory: The correlation coefficient between APR and the charge-off rate is positive, although small (0.13).

The positive correlation between APR and the delinquency rate, and between APR and the charge-off rate, provide no evidence in support of the "rational" scenario, namely that customers who do not repay their loans on time borrow at lower rates than convenience customers. The causality is not clear, however: Customers may irrationally ignore the rates in their

Figure 2

The mean APR and annual fee have declined over time.

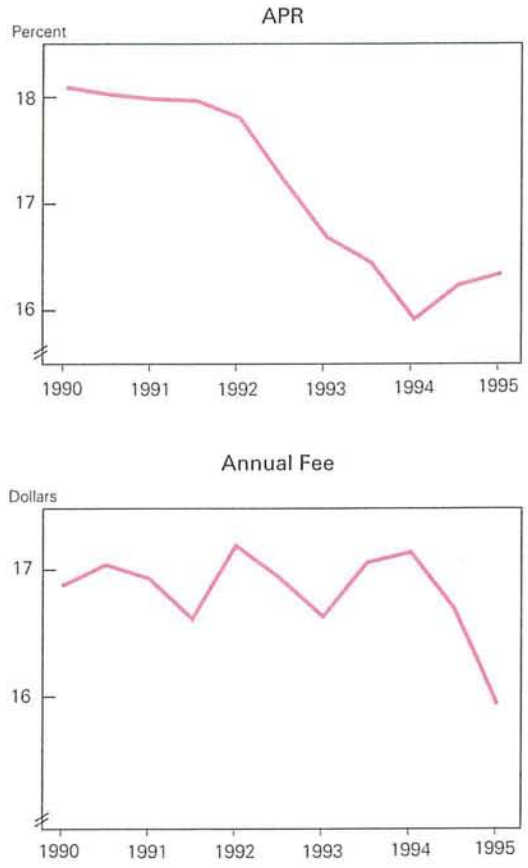


Figure 3

Mean late-payment fees and over-the-limit fees have increased over time.

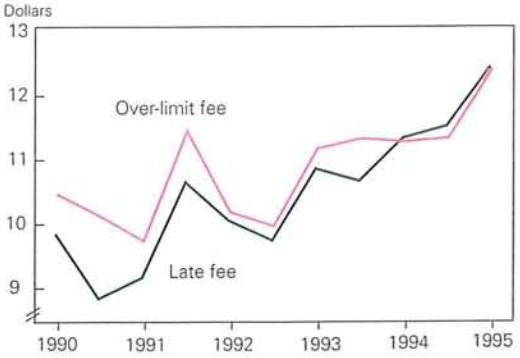


Figure 4

New England banks charge lower APR, but higher annual fee.

Figure 5

Variance of APR and variance of annual fee have increased over time.

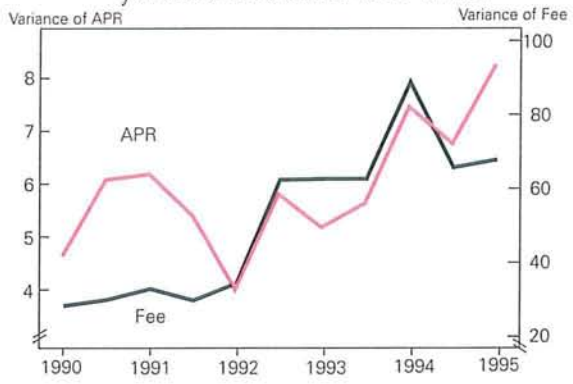
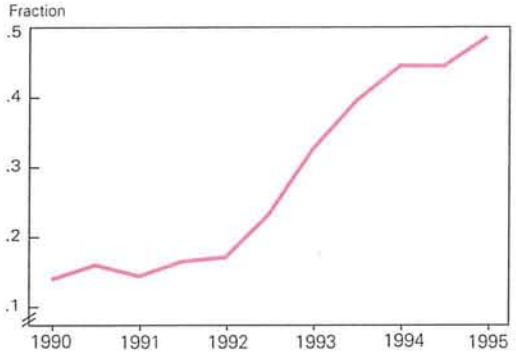


Figure 6

Fraction of plans with variable rate of interest has increased.



borrowing patterns, or banks may successfully steer them away from attractive borrowing options. It is possible that customers respond more strongly to other attributes of credit card plans. The next section uses regression analysis to determine whether consumers' demand for loans responds to the interest rates charged.

III. Consumer Demand for Credit Card Loans

Demand for credit card loans is a two-step process. Consumers first choose a credit card plan, then decide how much money to borrow. However, card selection is not an important determinant of borrowing patterns. Consumers typically own more than one credit card.[12] A consumer may carry both low- and high-APR cards. The subject of this study is consumers' actual borrowing patterns and not their credit card selections. The study therefore focuses on the amount actually borrowed on each plan and models only the second stage of the two-step demand process.[13]

[12] Estimates of the average number of charge and credit cards per household vary from 3.7 (*Fortune*, 6/27/94, p. 14) to 6 (1989 *Survey of Consumer Finances*, sponsored by the Federal Reserve).

[13] As the section above shows, banks may screen "good" and "bad" customers, making plan selection endogenous. Simultaneous supply and demand estimation will be used below to test for the endogeneity.

Overall Demand

The amount of credit card loans borrowed on a given plan is regressed on the plan's attributes, as well as on the prevailing interest rate on personal loans. Personal loans are typically the closest substitute for credit card borrowing (both types of loans are unsecured by assets). The specification is as follows:

$$CRLOAN_{it} = \beta_0 + \beta_i D_i + \beta_t D_t + \beta_1 APR_{it} + \beta_2 F_{it} + \beta_{3k} z_{kit} + \beta_4 r_t + \varepsilon_{it} \quad (1)$$

where $CRLOAN_{it}$ is the total amount outstanding of credit card loans for plan i in period t; D_i is a dummy for bank i; D_t is a dummy for period t; APR_{it} is the annual rate of interest under plan i in period t; F_{it} is the annual fee; z_{kit} is a vector of k other attributes of the plan;[14] r_t is the prevailing rate of interest for personal loans; ε_{it} is a random error term; β's are parameters to be estimated. Table 1 lists the variables used in the estimation, as well as their sources.

Table 1
Descriptive Statistics on Major Variables and Their Sources

Variable	Mean	Standard Deviation	Variable Definition
The Federal Reserve Board Survey on Terms of Credit Card Plans (TCCP):			
APR	17.21	2.43	Annual percentage rate on credit card loans
FEE	16.79	7.79	Annual fee ($)
GRACE	23.41	7.31	Grace period (days)
MINFIN	.52	1.83	Minimum finance charge ($)
CASH	2.02	1.93	Transaction fee for cash advance ($)
LATE	10.46	4.81	Late payment fee ($)
OVER	10.83	4.52	Over-the-credit-limit fee ($)
Enhancement Dummies:			
REBATE	.09	.29	Rebates on purchases
WARRANT	.19	.39	Extension of manufacturer's warranty
PROTECT	.20	.40	Purchase protection
ACCID	.62	.49	Travel accident insurance
TRADISC	.20	.40	Travel-related discounts
AUTO	.19	.39	Automobile rental insurance
BUYDISC	.04	.20	Purchase discounts (not travel-related)
REGIS	.19	.39	Credit card registration
OTHER	.28	.45	Other enhancements
Consolidated Reports of Condition and Income (Call Reports):			
CRLOAN	489.0 million	1,250 million	Credit card loans outstanding ($)
CREDINC	43.6 million	103.0 million	Interest and fee income on credit cards ($)
CHOFFS	12.5 million	34.8 million	Charge-offs on credit card loans ($)
OVERDUE	22.1 million	80.4 million	Overdue credit card loans ($)
DEPOSITS	4,140 million	7,200 million	Deposits in domestic offices ($)
Federal Reserve Bulletin:			
I_1YRTB	5.50	1.76	Interest rate on 1-year U.S. Treasury bill
I_PERSON	14.36	.90	Interest rate on personal loans
U.S. Bureau of Labor Statistics, *Employment and Wages: Annual Averages*:			
AWWAGE	583.18	147.50	Average weekly wages ($) by year and by state

Revolvers' Demand

The demand for credit card loans by revolvers (customers who carry a balance on their cards) should be more responsive to the rate of interest than the demand for loans by all cardholders. While convenience customers may be indifferent to the rate their issuer charges, at least some revolvers (those who behave consistently with their intentions) are likely to be sensitive to the rates charged. To find out how revolvers respond to APR, equation (1) is estimated using the balance of overdue accounts for bank i in period t ($OVERDUE_{it}$) as a dependent variable in place of the total amount of outstanding loans. Revolvers should be less likely than convenience users to care about other features, such as annual fee or enhancements, since most of their cost is driven by APR. Therefore, the elasticity of demand based on the coefficient on APR in the equation with overdue loans should be larger than the elasticity of demand based on the coefficient in the overall equation (1), while the coefficient on fee

[14] The vector of attributes may include grace period, minimum finance charge, late payment fee, and so on, as well as dummies for each plan's additional enhancements, such as automobile insurance, travel discounts, and extended warranty.

and on enhancements should be smaller than in equation (1).

To estimate how revolvers' demand responds to credit card features, data would be needed on the entire balance those customers carry from one month to the next. Unfortunately, banks report only the balances carried by customers who failed to pay the minimum balance on their credit card loans. The overdue loans reported by banks in the Call Report are therefore delinquent loans, not the total amount of revolving credit.[15] To the extent that customers with delinquent loans are especially risky to the banks, they may not constitute a representative sample of all the revolvers. The reported results may not therefore yield the information about the sensitivity of all the revolvers to interest rates.

IV. Estimation Results

Changing terms of credit card plans is costly to the issuer. Issuers announce the terms in newspaper ads and in mailings sent to their prospective customers, while current customers are typically informed about the terms of their plans once a year. The terms can therefore be assumed to be fixed over a three-month period.[16] Since the Call Report data on outstanding loans were collected two months *after* the TCCP data (see Section I), terms of credit card plans can be assumed to be exogenous when the loan amount is determined. Demand for credit card loans can therefore be estimated as a function of plan features using ordinary least squares (OLS).

Overall Demand for Credit Card Loans

Equation (1) was estimated using OLS. The restriction of no fixed effects was rejected at the 5 percent level.[17] Bank dummies were therefore included in the estimation. Replacing time dummies with a continuous time variable was not rejected, however. The linear specification was used, as it performed better than the double log specification. The estimation results are shown in Table 2.[18] The coefficient on APR is negative and significantly different from zero, showing that charging a higher APR leads to reduced credit card lending. The estimated elasticity of demand for loans with respect to APR is minus 1.47 (calculated at the mean). In other words, if an average bank dropped its APR by 1 percentage point (for example, from 17 to 16 percent, a 6 percent drop), its outstanding loans would rise by about 8 percent ($42 million for the average bank).

Table 2
Dependent Variable: $CRLOAN_{it}$
(outstanding credit card loans for bank i in period t)
Fixed effects estimation, 1990 to 1995

Variable	Coefficient (millions of dollars)	T-statistic
APR	−41.8	−3.21*
FEE	−5.9	−1.39
GRACE	16.4	1.71***
REBATE	92.0	.69
WARRANT	22.5	1.89***
PROTECT	74.3	.68
ACCID	−4.1	−.07
TRADISC	17.2	2.07**
AUTO	18.4	1.83***
BUYDISC	−11.9	−.10
REGIS	13.5	1.60
OTHER	−89.3	−1.23
I_PERSON	43.2	1.29
TIME	72.9	6.30*
INTERCEPT	−14.7	−.03

$R^2 = 0.13$ F = 37.32 N = 860
Note: See Table 1 for variable definitions and sources.
*Significant at the .01 level.
**Significant at the .05 level.
***Significant at the .10 level.

The coefficient on annual fee is negative, but not significantly different from zero. Credit card customers are, however, sensitive to the length of grace period, as shown by the positive and significant coefficient on GRACE. On average, extending the grace period by one day increases a bank's outstanding credit card loans by $16 million (a 3 percent increase). Among the enhancements added to credit card plans, consumers turned out to be most responsive to extension of manufacturer's warranty, travel accident insurance, and automobile rental insurance. Adding one of the three features raised an average bank's

[15] Approximately two-thirds of all credit card users are revolvers, while only about 3 percent of credit card loans are reported as overdue in the Call Report. Apparently most revolvers pay the minimum balance but do not pay the balance in full.

[16] In the TCCP sample, APR often remains constant over a period of one year or longer for a given credit card issuer.

[17] Fixed effects refers to estimation where intercepts are allowed to vary across individuals, here across individual banks.

[18] Since APR is a nominal interest rate, nominal dollars were used in the estimation.

outstanding credit card loans by $22.5 million (4.6 percent), $17.2 million (3.5 percent), and $18.4 million (3.8 percent), respectively.

Given the estimated effects of adding individual enhancements to a credit card plan, it is possible to calculate how much consumers spend on average on the added enhancements. By adding extension of manufacturer's warranty, a bank can expect to increase its outstanding credit card loans by 4.6 percent. Assuming that all accounts (those that pay interest and those that do not) raise their outstanding loans equally, an average interest-paying customer would increase his or her outstanding loan by 4.6 percent as well. An average credit card account carries a balance of $1,585.[19] A 4.6 percent increase would raise that average by $72.90. With a 17 percent average APR, adding extended warranty would increase an average cardholder's interest charges by $12.40 per year.[20] Similarly, adding travel accident insurance would raise an average cardholder's interest charges by about $9.40, while adding automobile rental insurance would raise his charges by $10.25. Although it is difficult to estimate whether enhancements added to credit cards are worth the money an average customer spends on them, evidence suggests that enhancements offer no savings to cardholders.[21] Cardholders who carry a balance on their cards are likely to minimize their spending by borrowing at the lowest possible APR and ignoring the added enhancements.

Credit card customers are sensitive to the rate of interest and also to the length of the grace period, extension of manufacturer's warranty, travel accident insurance, and automobile rental insurance.

In the final specification, minimum finance charge, cash advance fee, late payment fee, and over-the-limit fee were omitted. None of these variables obtained coefficients that were significantly different from zero and each had several missing values, limiting the number of observations used in the estimation. The finding that consumers' demand for loans is not sensitive to minimum financing charges or late penalties explains why banks have been raising those penalties—customers seem to be less responsive to them than to other features.

As explained above, interest rates on credit cards are fixed in the short run and can therefore be treated as exogenous in the demand estimation. However, to test for possible endogeneity of the interest rates due to banks' screening (see footnote 13), the demand for credit card loans was also estimated using three-stage least squares (3SLS). The demand equation (1) was estimated jointly with the following supply equation:

$$APR_{it} = \gamma_0 + \gamma_1 CRLOAN_{it} + \gamma_2 i_{1\ yr\ TB} + \gamma_3 w_t + \gamma_4 DEPOSITS_{it} + \gamma_5 t + \xi_{it} \qquad (2)$$

where $i_{1\ yr\ TB}$ is the cost of funds (measured as the interest rate on 1-year Treasury bill), w_t is the average weekly wage per employee in the finance sector for period t by state,[22] $DEPOSITS_{it}$ is the bank's deposits in domestic offices,[23] and t measures time in six-month intervals. The cost of funds, local wages in the finance sector, and the bank's deposits are exogenous instruments. In addition, an exogenous measure of income (GDP) was included as an instrument in the demand equation. The results were not significantly different from the OLS results: APR was statistically significant (although its coefficient was slightly lower in magnitude), the coefficient on GRACE remained statistically significant, and the coefficient on FEE was not significantly different from zero.

[19] Ausubel (1995) calculated the average outstanding credit card balance up to 1993. Extrapolating his numbers for 1994 and 1995 and then computing an average over the 1990–95 period yields $1,585. According to the Federal Reserve's *Functional Cost Analysis* (*FCA: National Average Report* 1994), the average size of an active credit card account is $1,028 for banks with deposits over $200 million. *FCA* samples only relatively small banks, however.

[20] That number is likely to underestimate the true interest and fee charges. For example, according to Ausubel's (1995) calculations, cardholders' monthly payment rate is only about 14 percent of their balances. The remaining balance is carried over to the next month. Such a low payment rate is likely to bring additional fees, such as late payment fees.

[21] Press articles have suggested that enhancements usually do not save customers any money, and they have cited anecdotal evidence that customers charged on their credit cards just to get the perks. For example, see *American Banker*, 9/22/95, p. 12 and *U.S. News & World Report*, 1/24/94, p. 68.

[22] The wages are taken from U.S. Bureau of Labor Statistics, *Employment and Wages: Annual Averages* for each year by state for finance, insurance, and real estate (4-digit SIC).

[23] To avoid endogeneity, deposits reported during the period preceding each TCCP survey were used in this estimation.

Table 3
Dependent Variable: $OVERDUE_{it}$
(overdue credit card loans for bank i in period t)
Fixed effects estimation, 1990 to 1995

Variable	Coefficient (millions of dollars)	T-statistic
APR	−3.48	−1.89***
FEE	.49	.64
GRACE	10.00	3.02*
MINFIN	−96.90	−1.44
CASH	−10.00	−1.68***
LATE	−7.23	−2.32**
OVER	6.62	1.60
REBATE	.42	.02
WARRANT	−44.30	−1.44
PROTECT	8.22	.30
ACCID	.25	.02
TRADISC	−10.10	−.63
AUTO	3.17	.13
BUYDISC	−46.00	−1.86***
OTHER	−14.80	−.87
I_PERSON	5.00	.70
TIME	−3.25	−1.16
INTERCEPT	−155.00	−1.19

$R^2 = 0.32$ F = 21.80 N = 157
Note: See Table 1 for variable definitions and sources.
*Significant at the .01 level.
**Significant at the .05 level.
***Significant at the .10 level.

Revolvers' Demand for Credit Card Loans

With the caveats noted in Section III, equation (1) was estimated using reported overdue credit card loan amounts as a dependent variable. The results are reported in Table 3. The results indicate, as expected, that the amount of overdue loans increases at a higher rate than loans in general as APR falls: The coefficient on APR is negative and significantly different from zero, yielding an elasticity of demand for overdue loans of 2.71 (calculated at the mean). The elasticity implies that if an average bank raised its APR by 1 percentage point (for example, from 17 to 18, a 6 percent increase), its overdue (delinquent) loans would decrease by 16 percent (about $3.5 million for an average bank), a much larger drop than for total loans. By the same token, lowering APR would increase the delinquent loans at a significantly higher rate than loans in general.

The amount of overdue loans is also more sensitive to the length of grace period (GRACE), the late payment fee (LATE), and the transaction fee for cash advances (CASH). The last two variables were insignificant in the overall demand equation. Higher charges for late payments, just like higher interest rates, seem to encourage customers to make payments on time or to switch to another plan.

Income from Credit Cards

Since demand for credit card loans is elastic with respect to APR, banks may be expected to lose income from credit cards as they raise their rates. On the other hand, since delinquent loans increase at a higher rate than credit card loans in general when APR falls, banks' income from credit card interest and fees could *increase* when they raise the interest rates charged. To examine which of the two effects dominates, the following regression was estimated:

$$INCOME = \lambda_0 + \lambda_1 APR_{it} + \lambda_2 FEE_{it} + \lambda_3 i_{1\,yr\,TB} + \lambda_4 w_t + \lambda_5 DEPOSITS_{it} + \lambda_6 t + \zeta_{it} \quad (3)$$

The results indicate that a bank's income increases both with the interest rate and with the annual fee the bank charges on its credit card loans, holding the bank's deposits and costs constant. As Table 4 shows, a 1 percentage point increase in APR (a 6 percent change) is associated with a $4.4 million average increase in the interest and fee income (a 10 percent rise), while a $1 increase in the annual fee translates into a $1 million average increase in the interest and fee income (a 2.3 percent rise) during a

Table 4
Dependent Variable: $CREDINC_{it}$
(income from card interest and fees for bank i in period t)
1990 to 1995

Variable	Coefficient (thousands of dollars)	T-statistic
APR	4391.2	3.86*
FEE	1044.8	2.88*
DEPOSITS	.000006	17.05*
I_1YRTB	1207.5	.73
AWWAGE	−54.8	−3.14*
TIME	1886.3	1.91***
INTERCEPT	56600	−2.02**

$R^2 = 0.17$ F = 57.40 N = 1650
Note: See Table 1 for variable definitions and sources.
*Significant at the .01 level
**Significant at the .05 level
***Significant at the .10 level

six-month period, holding the bank's deposits, cost of funds, and wages constant. Thus, banks benefit from raising both APR and annual fees.

The above results are significant. Although the overall outstanding credit card loans on a plan increase when the interest rate drops (the estimated demand elasticity is greater than one), banks collect less income when they lower their rates. Customers charge more as the rate of interest drops, but overdue loans increase even faster, indicating that cardholders repay at a lower rate. The reason for the discrepancy is the adverse selection problem faced by credit card banks: When lowering their rates, banks attract high-risk customers who are more likely to default (or their existing risky customers borrow more). Since banks lose income at lower interest rates, they maintain high APRs despite declines in the cost of funds.[24] The best strategy for banks to maximize their income is to charge high interest rates but entice their customers with additional perks, which cost banks little and are likely to attract "good" customers with a low probability of defaulting.

Customers charge more as the rate of interest drops, but overdue loans increase even faster, indicating that they repay at a lower rate. The reason is the adverse selection problem faced by credit card banks: When lowering their rates, banks attract high-risk customers.

Fixed Effects

As noted earlier, non-fixed effects specification was rejected by the data. The significance of fixed effects indicates that some individual banks' characteristics, besides the features of their credit card plans, affect demand for credit card loans. One hypothesis, examined in the next section, is that the size of a bank affects the demand for its credit card loans.

Table 5
Dependent Variable: APR_{it}
(annual percentage rate charged by bank i in period t)
1990 to 1995

Variable	Coefficient	T-statistic
DEPOSITS	3.25 e-11	4.59*
I_1YRTB	.061	1.82***
AWWAGE	−.001	−1.78***
TIME	−.227	−11.91*
INTERCEPT	18.673	56.68*

$R^2 = 0.13$ F = 68.07 N = 1884
Note: See Table 1 for variable definitions and sources.
*Significant at the .01 level.
**Significant at the .05 level.
***Significant at the .10 level.

V. Bank Size and the Credit Card Market

Although most credit card plans are offered nationally, some consumers may be more likely to turn to the bank that holds their deposits when applying for a credit card. Even if banks offer the same credit card terms to their clients as they do to others, consumers' preferences may give larger banks (those holding more accounts) an advantage in the credit card market. To test whether larger banks have a market power advantage that lets them charge higher interest rates on their credit card plans, APR was regressed on bank deposits.[25] The coefficient on deposits was positive and statistically significant (see Table 5), showing that banks with higher deposits charge higher interest rates on their credit card loans, holding the cost of funds and local wages constant. The effect was small, however; a $1 billion increase in deposits translated into an increase in the credit card interest rate of only 0.03 percent. The result indicates that large banks can take advantage of their market power and charge a somewhat higher rate of interest on credit card loans.[26]

[24] However, if banks raised their interest rates too high, they might lose even the "good" customers, who would forgo the convenience of borrowing on a credit card and borrow elsewhere.

[25] To avoid endogeneity, deposits reported during the period preceding each TCCP survey were used in this estimation.

[26] The size of deposits did not, however, affect banks' annual fees, other charges, or enhancements.

Table 6
Elasticity of Demand for Credit Card Loans at Banks by Size of Deposits

Deposits Category Banks	Average Credit Card Loans (millions of dollars)	Elasticity of Demand for Credit Card Loans
Less than $500 million	348.4	2.44
$500 million–$2 billion	359.3	.71
$2 billion–$5 billion	868.0	.69
Over $5 billion	674.5	.81

Since larger issuers, on average, charge higher interest rates, they should face more inelastic demand for credit card loans than do smaller banks (otherwise their strategy would not be profitable). Is demand for loans on credit cards issued by larger banks significantly less elastic than demand faced by smaller banks? When equation (1) was estimated by size-of-deposits category,[27] the smallest banks turned out to be the only group where demand for credit card loans was elastic with respect to the interest rates. See Table 6 for the results.[28] If an average bank from the first category (deposits below $500 million) raised its APR by 1 percentage point, its outstanding credit card loans would drop by 14.4 percent (a $50 million decrease). By comparison, a 1 percentage point increase in the APR charged by one of the largest banks would lead to only about a 4.6 percent decrease in the bank's credit card loans (a $31 million drop). Thus, small banks face much more elastic demand for credit card loans than large banks do. However, even when bank size was included in the regression, non-fixed effects estimation was rejected by the data. That indicates that other bank-specific factors in addition to size affect demand for credit card loans.

[27] Four deposit categories were used: below $500 million, between $500 million and $2 billion, between $2 billion and $5 billion, and over $5 billion. Each category contained approximately one-fourth of the sampled banks.

[28] When APR was interacted with the deposit size category in a pooled regression, the general result was confirmed: the larger the size category, the smaller the effect of APR on the amount of credit card loans. The pooled specification was rejected by the Chow test, however.

VI. Summary and Conclusions

Using data on the terms of credit card plans and on issuing banks' financial information, this study finds evidence that consumers' demand for credit card loans is elastic with respect to interest rates charged by issuing banks. The estimated demand elasticity for the overall market is 1.47, while the elasticity of demand for the delinquent credit card loans (loans that are at least 30 days overdue) is 2.71. Consumers' demand for loans was also found to be responsive to the length of the grace period and to some of the enhancements added by the issuing banks: extension of manufacturer's warranty, travel accident insurance, and automobile rental insurance.

Since demand for delinquent loans is significantly more elastic than demand for loans in general, lowering APR would attract disproportionately larger increases in delinquent loans than in loans borrowed by customers who pay back. Banks therefore face an adverse selection problem that induces them to maintain high interest rates on credit card loans in order to minimize their losses. The adverse selection hypothesis is further supported by the result that banks' income from credit card fees and interest increases with APR. Even though lowering APR would raise banks' outstanding credit card loans, the marginal customers attracted by lower APR would be risky and more likely to default, or the existing customers would borrow more than their incomes could service.

Significant fixed effects of individual banks exist in the credit card market. In particular, the largest banks (as measured by the size of deposits) charge slightly higher rates of interest and face more inelastic demand for credit card loans. Future research should examine whether the bank size effect is associated with market *power* or market *structure* characteristics. Even after controlling for bank size, fixed effects estimation cannot be rejected. That indicates that still another bank-specific factor may be affecting demand for credit card loans, in addition to size.

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The Growing Inequality of Family Incomes: Changing Families and Changing Wages

Katharine L. Bradbury

Vice President and Economist, Federal Reserve Bank of Boston. *Catherine Jew and Alicia C. Sasser provided valuable research assistance.*

It is widely known that the incomes of U.S. families became more unequal during the 1980s. The reasons for this rise, however, are not at all clear; numerous factors have been implicated. Economists observe that inequality typically declines during periods of economic expansion as the benefits of growth "trickle down" through the entire income distribution; yet most of the 1980s was characterized by economic growth as well as by growing income inequality.

One well-documented aspect of the rise in inequality is a growing gap between the wages of highly educated workers and those of high school dropouts or workers with only a high school degree, but no one factor satisfactorily explains this growing educational premium. Some of the rise in wage inequality translates directly into inequality of family incomes, since the wages of family members comprise most of family income. Furthermore, the United States experienced shifts in the mix of family types and changes in the work patterns of family members in the 1980s and 1990s that contributed to the increase.

This article examines inequality in the United States since the 1970s and investigates a number of hypotheses about why it is rising. Part I describes the 1973–94 increase in inequality of family incomes and related shifts in wage inequality, work trends, and family patterns. Several key facts emerge. Inequality has risen much more steeply among families than among individual workers, and much of the rise is due to increases in two categories of families concentrated at the top and bottom of the income distribution, respectively: two-earner married couples and families headed by one person with no spouse present. One-head families are concentrated closer to the bottom of the income distribution than two-earner married-couple families because they have fewer workers (by definition), and also because one-head family workers average fewer work hours and earn less per hour. Combining all family types, incomes have risen for the highest-income families and declined for the lowest as families near the top of the income distribution gained in number of

workers, hours per worker, and earnings per hour relative to those near the bottom.

Part II examines patterns of inequality among the nine Census regions in the United States as well as differences in their economic and demographic characteristics. Part III investigates the relationship between family income inequality and these factors, sorting out and quantifying their contributions to overall U.S. and regional changes in inequality. In brief, changes in both economic factors and family structure have been associated with rising family income inequality over the last two decades, with the increase in single parenthood and the growing wage premium to college education playing key roles. Among regions, part-time work, low labor force participation, and large minority populations are associated with greater inequality. A discussion of policy options for reducing income inequality concludes the article.

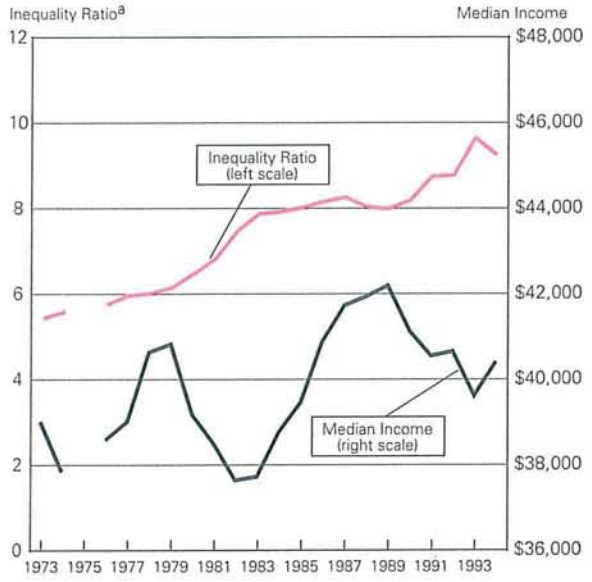
I. Patterns of Family Income Inequality in the United States, 1973 to 1994

Family income inequality increased fairly steadily during the past two decades. Figure 1 shows this upward trend in one measure of inequality. (Appendix A shows trends for alternative measures.) The measure of inequality used here is the ratio of the income of a high-income family, defined as the 90th percentile family (meaning 90 percent of families have lower incomes), to the income of a low-income family—the 10th percentile family (10 percent of families have lower incomes).[1] To facilitate comparisons between worker incomes and family incomes, this analysis focuses on nonelderly families, defined as those whose head is under age 65. Single individuals (living alone or with non-relatives) are also excluded from the analysis.[2]

As Figure 1 shows, the high-income family had about five and one-half times the income of the low-income family in 1973. In 1994, the high-income family had over nine times as much. The increase in inequality between 1973 and 1994 reflects both

Figure 1

Nonelderly Family Median Income and Income Inequality Ratio



Note: Data are missing for 1975. Median income in 1993 dollars, calculated using U.S. CPI-U-X1.

[a] Inequality measured as ratio of 90th percentile family income to 10th percentile.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey,* March 1974 through March 1995.

rising real incomes for those at the top of the income distribution and falling real incomes for those at the bottom. Figure 2 summarizes real income changes for families at the 10th through 90th percentiles of the income distribution, showing a significant drop in real income at the bottom, substantial real growth at the top, and remarkably regular corresponding changes in between.[3]

Inequality rises in recessions as the negative shocks to income caused by recession job cutbacks are borne more heavily by those at the bottom, and declines in expansions as jobs are regained and prosperity "trickles down" through the income distribution.[4] While Figure 1 shows increases in inequality in most years, it rose faster in the recession years 1980–82 and 1990–91 than in the expansions of the late 1970s and mid to late 1980s. Furthermore, this

[1] The income measure is the total money income of the family; no adjustments are made for family size or for in-kind, unreported, or unrealized income.

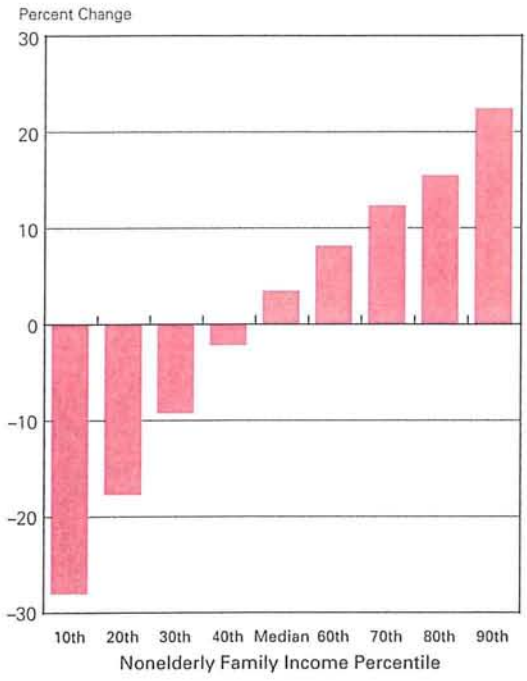
[2] The analysis is limited to "families" as defined by the U.S. Census Bureau, which consist of two or more related individuals living together. Families comprised just over 70 percent of households in the United States in 1994, down from just over 75 percent in 1973.

[3] Note that these data do not track individual families over time; rather, they measure changes from one year to another in the annual cross section of family incomes.

[4] See Appendix B for a description of year-to-year income changes at the top and bottom of the income distribution.

Figure 2

Change in Real Family Income, 1973 to 1994



Note: Constant-dollar income calculated using U.S. CPI-U-X1.
Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974 and March 1995.

measure of inequality actually declined toward the end of the expansion in 1988 and 1989, and again in 1994. Counter to the typical cyclical pattern, however, inequality climbed steeply in the recovery year of 1993 (median income declined that year, too).[5] The data thus reveal a strong upward trend to income inequality since the mid 1970s and also a cyclical pattern.

Three categories of long- and short-run influences affect the distribution of family incomes: (1) factors such as education and work hours that influence the distribution of earnings of individual workers; (2) the mix of family types and work patterns within families (which family members work and how many hours); and (3) the relationship between family earnings and other sources of family income.

Individual Earnings

Since the earnings of individual family members combine to form family earnings, and earnings comprise the bulk of family income, factors affecting the shape of the earnings distribution undoubtedly alter the shape of the family income distribution as well. Unlike the distribution of family incomes, however, individual earnings were not characterized by growing inequality until the 1980s. Presumably, changes in family structure and work patterns altered the shape of the family income distribution relative to that of individual earnings in the 1970s.

Parallel to the measure used for family incomes, earnings inequality is measured here as the ratio of annual earned income of a nonelderly high-earning (90th percentile) worker to that of a nonelderly low-earning (10th percentile) worker. For nonelderly men working full-time and year-round,[6] earnings inequality held steady between 1973 and 1979, and then the ratio rose from 4.0 in 1979 to 5.4 in 1994 (Table 1). For women, the ratio fell in the 1970s and then rose from 3.2 in 1979 to 4.5 in 1994.

For men and women combined, inequality rose in the 1980s and '90s, but less than for either men or women considered separately. The 90th percentile full-time, full-year worker had about 4.4 times the earnings of the 10th percentile worker in 1979, and 5.5 times as much in 1994. The overall earnings distribution spread out less than its male and female components because of two changes in the work roles of women. Women represented a growing share of the full-time, full-year work force, rising from 34 percent in 1979 to 40 percent in 1994. Furthermore, their labor market success was increasing: Women's earnings rose in real terms, on average, while men's real earnings declined (center panel of Table 1). In a sense, women filled out what had been the middle of the male earnings distribution: The median full-time, full-year woman's earnings rose from being about equal to the 20th percentile man's earnings in 1979 to above the 30th percentile male earner in 1994, while the 90th percentile woman moved up from the male median to above the 70th percentile man.

Many studies have sought to uncover economic and demographic causes for the growth in earnings inequality.[7] The most thoroughly studied aspect of the 1980s growth in earnings inequality is the increase in the educational wage premium: The amount by which the pay of college-educated workers exceeded that of

[5] Some of the 1993 jump and 1994 decline in inequality may reflect changes in Current Population Survey methodology; see Ryscavage (1995).

[6] Full-time is defined as working 35 or more hours per week in a majority of the weeks worked in the calendar year. Year-round (full-year) is working 50 or more weeks in the year.

[7] This literature was reviewed by Levy and Murnane in 1992 and, most recently, summarized by Kodrzycki (1996).

Table 1
Inequality of Individual Earnings
Annual Earned Income of Workers

	1973	1979	1989	1994	Change 1979–94
Earnings inequality ratio—full-time, full-year workers					Difference
Male	4.0	4.0	4.7	5.4	+1.5
Female	3.6	3.2	4.1	4.5	+1.3
Combined	4.5	4.4	5.0	5.5	+1.1
Women as percent of all full-time, full-year workers	30.8	34.5	38.8	40.0	Percentage Point +5.5
Median earnings of full-time, full-year workers (1993 $000)					Percent
Male	33.1	33.2	31.5	29.3	−11.9
Female	18.4	19.5	21.6	21.5	+9.9
Combined	27.6	27.3	26.8	25.4	−7.3
Earnings inequality ratio—workers on all work schedules					Difference
Male	18.0	16.8	16.7	17.1	+.4
Female	42.4	28.8	23.1	22.5	−6.3
Combined	33.3	26.8	21.0	21.4	−5.4
Percent of workers on full-time, full-year schedules					Percentage Point
Male	68.5	66.9	70.3	70.6	+3.7
Female	41.9	43.9	51.9	53.7	+9.8
Combined	57.3	56.7	61.8	62.7	+6.0

Note: Inequality measured as ratio of earnings of 90th percentile worker to earnings of 10th percentile worker. Constant-dollar earnings calculated using U.S. CPI-U-X1.
Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, 1980, 1990, and 1995.

high school graduates or high school dropouts grew considerably in the 1980s. (See the box.) Slower growth in the supply of college-educated workers than in the demand for them (and the opposite for workers with less education) is commonly cited as the cause of the growing gap in pay between college-educated and less educated workers. Supply changes include two demographic shifts—a slowdown in the rate of increase in the college-educated work force as the baby boom generation moved beyond the typical college-completing years, and a shift in the mix of immigrants toward less-skilled workers. Most reviews cite three key factors contributing to faster growth in demand for college graduates than for those less well-educated: shifting industry mix; increased international trade, especially with less developed nations; and technological changes that raise the productivity of more skilled workers relative to less skilled.

Overall wage inequality rose because disparities grew not only between educational groups, but also between experience (or age) categories, and even among workers with similar experience and education. Along with the causes of the growing educational wage premium, institutional factors, such as the changing role of unions and a declining real minimum wage, are thought to have contributed to greater earnings inequality.

Industry mix. Shifts in industry mix, notably the loss of manufacturing jobs and growth in services industries, have received considerable public attention although the economics research does not assign them great importance in explaining growing wage inequality.[8] Part of the reason for the public attention is that plant closings and layoffs in the manufacturing sector have been quite visible in specific local labor markets, and the associated disruptions to the work lives of individual workers have not been evenly spread among the population.

People often characterize the lost manufacturing jobs as "good" jobs, with relatively high pay and substantial fringe benefits, while services and retail trade jobs—the largest component of the nonmanufacturing additions—are considered "bad" jobs, with low pay and few fringe benefits. In fact, services and retail trade differ in their earnings profiles, with trade concentrated at the low end of the earnings distribution while services workers are more spread out at both the bottom and the top of the earnings distribution than manufacturing.[9] Thus, as workers shifted, on net, out

[8] Most studies find changes in industry mix to be a contributing factor, but a small one. The simplest evidence that the role of industry mix shifts is limited is the fact that wage inequality has expanded substantially *within* virtually all industry categories.

[9] For those working full-time and all year, earnings inequality is higher in services and retail trade than in manufacturing. The 90th/10th ratio was 3.5 for manufacturing workers in 1973 as compared with 5.2 in retail trade and 4.8 in services; by 1994, the ratios had risen to 4.7 in manufacturing, 5.5 in retail, and 5.6 in services. In both 1973 and 1994, the 10th percentile worker in retail trade had lower earnings than in services, and both earned less than the 10th percentile worker in manufacturing. The 90th percentile worker in services earned more in 1973 than the 90th percentile worker in manufacturing, but this was not the case in later years; the services worker earned slightly less.

The Educational Wage Premium

The table documents the size of the earnings premium garnered by college graduates compared with high school graduates and those not completing high school. For example, the median full-time, full-year male worker with a college degree earned 32 percent more than the median high school graduate working full-time all year in 1979, and 53 percent more than the median high school dropout holding a full-time year-round job. By 1994, those differences had risen to 71 percent and 171 percent, respectively. The educational wage premiums for women and for men and women combined were similar to those for men in each year.

The earnings premium for a college degree is higher when part-time and part-year workers are included (third panel in table), because less educated individuals are more likely to hold part-time or part-year jobs and thereby earn less annually on that score as well as on account of their education. This difference is particularly noticeable for women, since a greater fraction of female workers are on part-time or part-year work schedules. The premiums are even higher when nonworkers are included as well (data not shown), since less educated persons are also less likely to be working than college graduates.

The table also shows the nature of the increase in wage inequality between educational groups: Real wages declined precipitously for those without a high school degree, even those working full-time year-round, and rose or held steady for workers with a college degree or more. For men, the rise was very slight in real terms, but for women, it was sizable.

The Educational Wage Premium
Annual Earnings

Workers Age 25 to 64	Men	Women	Both
Full-Time, Full-Year (FTFY) Workers:			
Ratio: BA/no HS[a]			
1979	1.53	1.56	1.57
1989	1.74	1.72	1.75
1994	2.71	2.75	2.67
1979–94 Change	+1.17	+1.19	+1.10
Ratio: BA/HS[a]			
1979	1.32	1.41	1.48
1989	1.60	1.67	1.68
1994	1.71	1.74	1.73
1979–94 Change	+.39	+.32	+.25
All Workers:			
Ratio: BA/HS			
1979	1.35	1.60	1.64
1989	1.69	1.92	1.76
1994	1.68	2.00	1.84
1979–94 Change	+.33	+.40	+.20
Percent Change in FTFY Median Real Earnings, 1979–94:			
no HS	−43.2	−33.4	−41.1
HS	−22.6	−4.2	−14.6
BA	.3	17.7	−.1

[a]no HS = did not complete high school; HS = high school graduate with no college; BA = college degree or higher.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1980, March 1990, March 1995.

of manufacturing and into services jobs, the overall earnings distribution would be expected to become more unequal. A shift toward retail trade jobs would increase the weight at the bottom of the distribution.

However, the earnings distributions within each major industry also became more unequal during the 1980s. And while the manufacturing and services industries' earnings distributions both became more unequal, they were less different from one another in 1994 than they had been in 1973.[10] Both of these changes reduce the link between the tilt away from manufacturing and the rise in inequality.

Part-time and part-year work. Most analyses of earnings inequality focus on full-time, full-year workers because the inclusion of part-time or part-year workers, especially when earnings are measured on an annual basis, mixes together changes (or differences) in work schedules with changes in rates of pay. Yet in considering how individual earnings inequality may translate into family income inequality, these work schedule differences are also relevant.

[10] See Schweitzer and Dupuy (1995) for a careful analysis of this convergence.

Variations in work schedules make the earnings distribution for part-time or part-year workers considerably more unequal than the full-time, full-year distribution. Earnings inequality in 1973 and 1979 was over four times greater for men on all work schedules (working full-time or part-time, full-year or part-year) than for full-time, full-year men. Male inequality rose very slightly between 1979 and 1994. (See the lower panels of Table 1.) For women on all work schedules, by contrast, earnings inequality declined. And among all working men and women combined, inequality also declined. It was the marked shift of women into full-time, full-year work that brought inequality down.[11] The fraction of women working full-time and all year rose by 10 percentage points between 1979 and 1994 even as their numbers in the work force also increased. Men also shifted toward full-time, full-year work, but less markedly.

Family Structure and Family Work Patterns

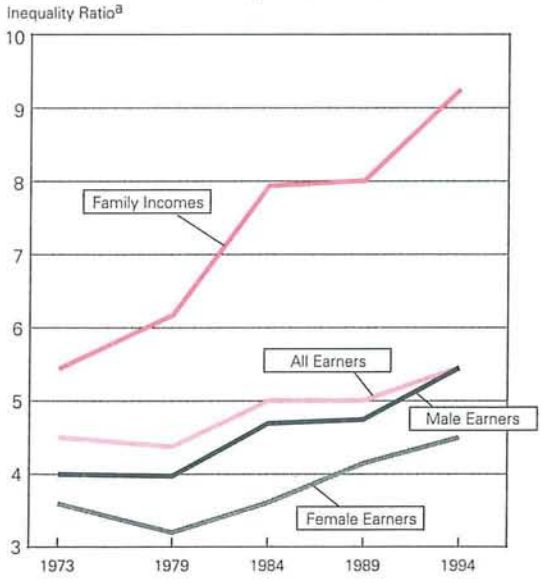
If each family had one worker and no nonearnings income, the family income distribution would match the overall distribution of earnings (for men and women on all work schedules combined). In fact, of course, some families have more than one earner and some have none. Furthermore, while earnings are the primary source, nonearnings account for about one-tenth of total family income, on average.

Figure 3 compares the inequality of family incomes with individual earnings inequality for selected years between 1973 and 1994. Two facts stand out: First, total income is distributed more unequally among families than are earnings among full-time, full-year workers. That is, the $90^{th}/10^{th}$ ratio is higher for family income than for individual earnings. Second, inequality increased considerably more among families than among individual full-time, full-year earners.[12] These facts imply that factors other than individual earnings inequality have made important contributions to family income inequality and its rise in the last two decades. The paragraphs that follow explore the characteristics of families and family earnings in terms of the individual earners.

Figure 3

Inequality of Individual Earnings and Family Incomes

Incomes of Nonelderly Families and Earnings of Full-Time, Full-Year Nonelderly Workers, Selected Years



[a] Inequality measured as ratio of 90th percentile earnings or income to 10th percentile.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, March 1980, March 1985, March 1990, and March 1995.

The most important distinction among family types from an income-generating point of view is between those headed by a married couple and those headed by an individual with no spouse present (either the head has no spouse or the spouse is absent).[13] The mix of such families has shifted consid-

[11] Looking at all work schedules, the ratio of 90^{th} percentile earnings to 10^{th} percentile earnings is quite high (as can be seen in Table 1), mostly because the 10^{th} percentile worker earns relatively little. Part-time or part-year workers typically have lower hourly earnings than full-time, full-year workers, and their lower hours and/or weeks mean that their annual earnings are even lower. Using the ratio of 80^{th} percentile earnings to 20^{th} percentile earnings reduces the sensitivity of the measure to very low part-time or part-year earnings at the bottom; but even using this measure, inequality rose for men and fell for women and for both genders combined.

[12] And much much more than among earners on all work schedules since, as just noted, earnings inequality did not increase at all for this group—it declined. Among the three full-time, full-year earnings inequality measures shown in Figure 3, the male measure most closely tracks the slopes, timing, and magnitude of changes in family income inequality. While full-time, full-year men comprise only three-eighths of all workers, over three-fifths of all families in 1994 included a full-time, full-year male worker.

[13] This section discusses family income as a function of work patterns within families, ignoring the nonearnings component of family income. Differences in family structure and the number and work schedules of family earners turn out to be much more important than variations in nonearnings income in understanding the rise in family income inequality as well as in explaining Figure 3's differences between inequality trends for family incomes and individual earnings. The nonearnings component of family income is discussed below.

erably in the last two decades. Married-couple families comprised 86 percent of all nonelderly families in 1973, but only 76 percent by 1994.[14]

Married-couple families typically have higher incomes than "one-head" families. The average income of married-couple families was $47,100 in 1973, while one-head families averaged roughly half that much—$24,500. Thus, the shift toward one-head families would cause a drop in average family income, other things equal. But other things were not equal: The average real income of married-couple families rose over the 21 years to $54,500, while the average income of one-head families declined slightly to $24,200. These changes in average incomes, combined with the shift toward one-head families, add weight to both ends of the income distribution; that is, they increase inequality.

The most important distinction among family types from an income-generating point of view is between those headed by a married couple and those headed by an individual with no spouse present.

The rise in married couples' incomes is partly attributable to widely recognized changes in work patterns in married-couple families. The fraction of married-couple families in which both the husband and wife worked rose from 55 percent in 1973 to 70 percent in 1994, while the fraction in which the husband or wife (but not both) was working declined from 42 to 26 percent.[15] Furthermore, this rise in two-earner couples has not occurred evenly across the income distribution; it has been more pronounced in families in which the husband's earnings are higher.[16]

Not surprisingly, the incomes of two-earner married-couple families generally exceed those of one-earner married-couple families.[17] While a sizable portion of the gain in married-couple family incomes was attributable to the shift toward families in which husband and wife were both working, the incomes of two-earner couples also increased noticeably relative to other married-couple families over the 1973–94 period.[18]

The data in Figure 4 illustrate the importance of work patterns in the determination of family income. In 1994, families in which both the husband and wife worked (shown in the left-most box; the key to data items is outside the top box) reaped the earnings from about 70 percent more work hours per year than families in which only one or the other worked (the adjacent box). Differences in the number of work hours are partially offset, however, by higher hourly earnings for workers in one-earner married-couple families.[19]

Hourly earnings, more than work hours, are the crucial determinant of income differences between one-head families in which the head works and married-couple families. The family work hours of one-head families with the head working are similar to those of one-earner married-couple families (Figure 4 again). But these one-head families average less than 65 percent of the family income (and family earnings, data not shown) of one-earner married-couple families because heads of families with no spouse command considerably lower hourly earnings than the working husband or wife in married-couple families.

[14] Throughout this research, data drawn from the U.S. Current Population Survey—reporting incomes for a calendar year and demographic characteristics as of the following March—are referred to as if the demographic data also referred to the calendar year. The text sentence to which this footnote is attached, for example, actually reports on the mix of family types in March 1974 and March 1995.

[15] The fraction of married-couple families in which neither head nor spouse worked rose from 3 percent in 1973 to 4 percent in 1994.

[16] See Cancian, Danziger, and Gottschalk (1993).

[17] The "two-earner" and "one-earner" labels are used henceforth as shorthand for husband-and-wife-working and husband-or-wife-working married-couple families, respectively. The characterization is actually inaccurate, however, since husband-and-wife-working families averaged 2.3 to 2.4 workers (depending on year) and husband-or-wife-working ("one-earner") families averaged 1.3 to 1.4 workers. These additional workers are typically the couple's working-age children.

[18] The combined effect of these two changes is indicated by the fact that in 1973, 60 percent of the families in the top quintile were married-couple families with husband and wife both working; in 1994, that figure was 78 percent. This rise occurred despite the declining share of married couples among nonelderly families overall. Note that each quintile, by definition, includes one-fifth of all nonelderly families ranked by income.

[19] Hourly earnings are defined as total family earned income divided by total family work hours (the sum of annual work hours of all workers in the family). The higher hourly earnings in one-earner families are not exogenous, of course: The decision to send the second spouse into the labor market reflects, in many cases, the earnings opportunities of the "primary" earner; and even high-earning primary earners may have the family's average earnings per hour pulled down by a lower-earning second earner.

Figure 4

Income Characteristics by Type of Family, 1994

All families with nonelderly head

Key	
• Number in thousands (share (%) of all)	58,746 (100.0)
• 1973-94 change in share	n.a.
• % in lowest income quintile~highest	20 ~ 20
• Average # of members, # of workers	3.3, 1.8
• Percent with children	62
• Average annual family work hours	3,200
• Average family income ($)	47,300
• 1973-94 income change (%)	+7.8

Married-couple families

	44,639 (76.0)
	-9.6
	11 ~ 25
	3.4, 1.9
	58
	3,600
	54,500
	+15.8

Husband and wife work

31,338 (70.2)[a]	
+15.3[a]	
5 ~ 29	
3.4, 2.3	
59	
4,200	
60,300	
+19.8	

Husband or wife works

11,419 (25.6)[a]	
-16.9[a]	
19 ~ 17	
3.6, 1.3	
60	
2,400	
44,100	
-.9	

Neither husband nor wife works

	Other adult works	No adult works
	318 (.7)[a]	1,563 (3.5)[a]
	+0.2[a]	+1.5[a]
	30 ~ 7	61 ~ 4
	4.1, 1.2	2.8, 0
	38	29
	1,800	0
	33,000	20,000
	+2.4	+23.4

One-head families

	14,107 (24.0)
	+9.6
	49 ~ 4
	2.9, 1.2
	75
	1,900
	24,200
	-1.0

Head works

10,855 (77.0)[b]	
+4.9[b]	
39 ~ 5	
2.8, 1.5	
74	
2,400	
28,200	
-1.1	

Head does not work

	Other adult works	No adult works
	777 (5.5)[b]	2,475 (17.5)[b]
	-3.5[b]	-1.4[b]
	52 ~ 3	94 ~ 0
	3.7, 1.2	3.2, 0
	55	89
	1,600	0
	22,100	7,200
	+2.4	-29.5

Note: hours and income rounded to nearest 100; demographic characteristics as of March 1995.

[a]Percent of married-couple families (or change in that percentage).
[b]Percent of one-head families (or change in that percentage).

Source: Author's calculations based on U.S. Census Bureau, *Current Population Survey*, March 1995.

Table 2
Educational Attainment, 1994
By income quintile and family type

Family Income Quintile	Percent Who Graduated from High School			Percent Who Graduated from College		
	Husbands	Wives	Female Heads[a]	Husbands	Wives	Female Heads[a]
Poorest	60	64	58	10	8	3
Second	78	80	68	11	9	8
Middle	87	89	67	17	15	17
Fourth	92	94	60	31	27	21
Richest	92	96	54	53	45	22
All	85	88	62	28	23	9

[a]Female head of family, no husband present; over four-fifths of one-head families are headed by a woman. Income in 1994; other characteristics as of March 1995.
Source: Author's calculations from U.S. Bureau of the Census, *Current Population Survey*, March 1995.

Married-couple families with one spouse working are a mixture of those whose primary worker's earnings are high enough that the spouse can "afford" not to work, those who choose to have one spouse at home taking care of children or other obligations, and those whose spouse has such poor labor market prospects that his or her earnings would not offset the costs of going to work. But one-head families, by definition, cannot choose the best earner among two spouses. Furthermore, most (over 80 percent) family heads with no spouse present are women. They typically have less education, fewer skills, and less work experience than the working spouse (male or female) in one-earner married-couple families. Table 2 shows markedly lower educational attainment for female heads of families with no spouse present than for husbands or wives in married-couple families.[20]

Most one-head families are in the lower income quintiles. Those who do make it into the upper quintiles have similar numbers of workers to married-couple families in those quintiles (Table 3). That is, additional workers contribute the earnings that put families higher on the income scale, but very few one-head families have those additional workers.

As noted earlier, the average income of one-head families declined in real terms between 1973 and 1994. This loss was not attributable to a decline in the fraction with the head working; that fraction actually rose. Work hours also rose.[21] The loss was, instead, attributable to declining hourly wages for working family heads with no spouse present. Since these heads are mostly women and real wages for most women, unlike most men, rose over the period, the decline is undoubtedly associated with the low education levels shown in Table 2, which put most single heads on the losing side of the rising educational wage premium.

The relative importance of differences in work patterns, work effort, and earning power is revealed in Table 4. When the data are summarized by quintile (combining all family types), they indicate that differences in all three factors contribute to the income differences that define the quintiles. The more than 14-to-1 ratio of top-quintile to bottom-quintile annual family earnings is the product of disparities in hourly earnings, number of workers, and hours per worker. The hourly earnings differences are most marked, with workers in top-quintile families averaging over four times the hourly earnings of those in the bottom quintile.[22]

Furthermore, the increase in disparity between earnings in the top and bottom quintiles is attributable to increasing disparities in all three factors.[23] That is, the number of workers per family, hours per worker, and real earnings per hour all fell for the poorest quintile relative to the richest quintile during the 1980s and 1990s. Hourly earnings showed the greatest difference in growth rates between rich and poor families over the period, but all three gaps expanded.

In sum, shifts in the mix of families combined with changes in the earnings and hours of workers in various types of families to raise family income inequality over the period 1973 to 1994. Two-earner married-couple families and one-head families both increased as a share of all families. The incomes of

[20] While it may seem puzzling that the average high school graduation rates of female heads are not higher in high-income quintiles, note that data in Table 3 indicate that workers other than the head, whose education levels are not shown, account for increasing shares of income in higher quintiles.

[21] This statement applies to the 1979–94 period; hours are not reported in a comparable way for 1973. Appendix C (Figure C) reports Figure 4's data for 1979.

[22] Differences in hourly earnings also make bigger contributions to disparities between top and middle, middle and bottom, and fourth and second quintiles than the other two factors.

[23] This statement applies to the 1979–94 period. The calculation cannot be done for 1973 because hours data are not reported in a comparable way. Appendix C (Table C) reports Table 4's data for 1979.

Table 3
Number of Workers, 1994
By income quintile and family type

Family Income Quintile	Percent of Families: Married-Couple Families	Percent of Families: One-Head Families	Average Number of Workers in Family: Married-Couple Families	Average Number of Workers in Family: One-Head Families
Poorest	8.1	11.9	1.2	.8
Second	14.1	5.9	1.7	1.4
Middle	16.6	3.5	1.9	1.7
Fourth	18.1	1.8	2.1	1.9
Richest	19.1	.9	2.3	2.1
All	76.0	24.0	1.9	1.2

Note: Family type as of March 1995; income and number of workers in calendar year 1994.
Source: Author's calculations from U.S. Bureau of the Census, *Current Population Survey*, March 1995.

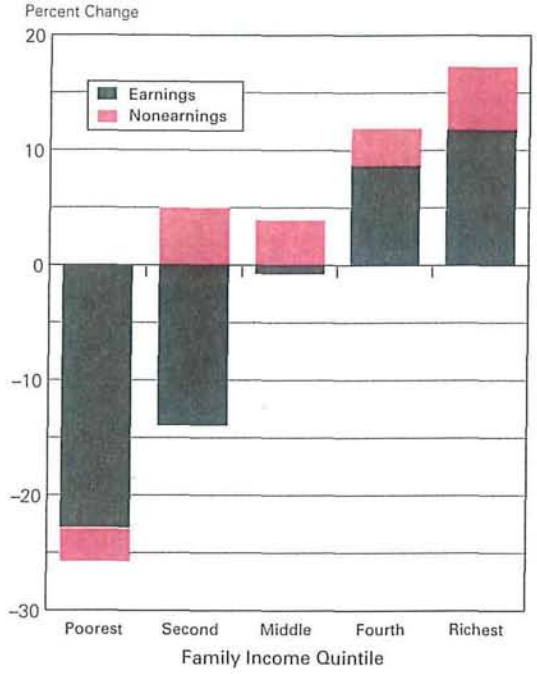
the former are well above average and the latter well below average because of differences in number of workers in the family (by definition), work hours per family worker, and earnings per family work hour. Hence the shift in mix raised inequality, other things equal. Compounding the shift is the fact that the incomes of two-earner married-couple families grew faster than those of one-head families, largely because of different growth rates in their hourly earnings. This

Table 4
Family Work Characteristics by Income Quintile, 1994

Family Income Quintile	Average Annual Family Earnings	Average Number of Family Workers	Average Annual Hours per Worker	Average Earnings per Hour
Poorest	$ 6,250	1.0	1,380	$ 4.77
Second	21,600	1.6	1,710	7.84
Middle	36,600	1.9	1,830	10.60
Fourth	54,150	2.1	1,880	13.72
Richest	90,750	2.3	1,930	20.73
All	$41,850	1.8	1,800	$13.21
Ratio: Highest to Lowest	14.5	2.4	1.4	4.3

Note: Average annual family earnings rounded to nearest $50; average hours rounded to nearest 10.
Source: Author's calculations from U.S. Bureau of the Census, *Current Population Survey*, March 1995.

Figure 5
Changes in Family Income, 1973 to 1994
Changes Attributable to Earnings and Nonearnings, by Quintile



Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974 and March 1995.

difference, in turn, presumably reflects differences in their initial earnings levels and the rising inequality of hourly earnings in the economy as a whole—declines for individual earners at the bottom and growth at the top.[24]

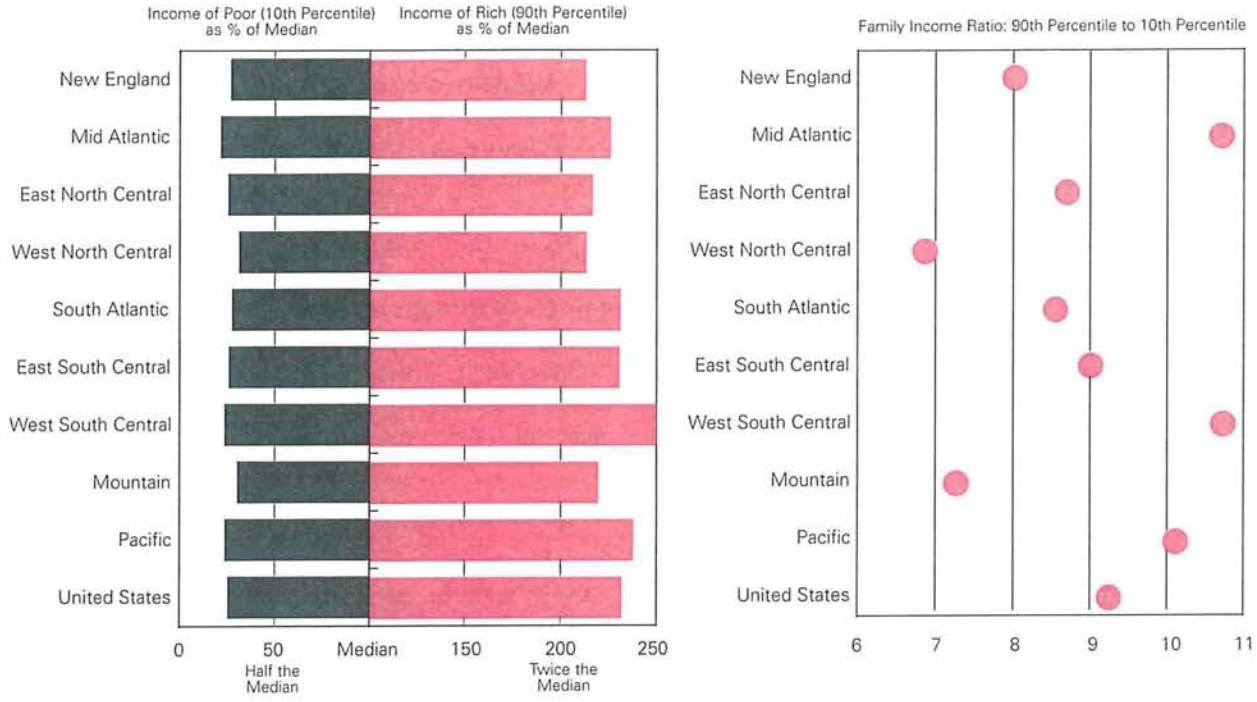
Nonearnings Income

The analysis to this point has focused on work patterns because earnings are by far the largest source of family income, comprising 88 to 92 percent of total family income (depending on the year). But as can be seen in Figure 5, changes in income other than earnings generally reinforced the growing earnings disparities between the richest and poorest families. Nonearnings income declined in real terms for the

[24] The rising educational wage premium that underlies a large part of the rise in wage inequality presumably explains some of these differences among family types in the growth rate of earnings. As Table 2 documents, heads with no spouse present have much lower educational attainment, on average, than either spouse in married-couple families.

Figure 6

Income Gap between Rich and Poor, 1994



See Table 5 for region definitions.
Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1995.

lowest income quintile and rose for the four higher quintiles.

The Census identifies two types of earnings (wage and salary income and self-employment income) and four broad types of nonearnings income—property income (including interest, dividends, and rental income), retirement income, transfer income, and "other" income. Wages and salaries accounted for 82 to 83 percent of income and self-employment averaged over 6 percent. The other sources each comprised 2 to 4 percent of total income. The importance of retirement and property income grew over the two decades (even though the families included here were headed by individuals under age 65) and transfers shrank.

Low-income families had a noticeably different mix of income sources than high-income families. The fraction of income from earnings is lowest for the poorest quintile (although still greater than 60 percent). Transfers (and retirement income) are more important in the bottom quintile than for families further up the income ladder. Transfers comprised a larger share of the poorest quintile's family income in 1994 than in 1973 (rising from 19 to 22 percent), even though the poorest quintile's average transfer income actually dropped 15 percent in real terms (that is, corrected for inflation).[25] Families in the poorest quintile experienced real declines in all sources of family income over the 1973–94 period.

II. Regional Differences in Family Income Inequality

While national business cycles and structural changes in the economy and in family organization affected all regions of the country, their magnitude

[25] Earnings, property income, and retirement income all declined faster, in real terms, for the poorest quintile than did transfers, so the transfer share rose.

Table 5
Bureau of the Census Definitions of U.S. Regions

New England (NE): Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, Vermont
Middle Atlantic (MA): New Jersey, New York, Pennsylvania
East North Central (ENC): Illinois, Indiana, Michigan, Ohio, Wisconsin
West North Central (WNC): Iowa, Kansas, Minnesota, Missouri, Nebraska, North Dakota, South Dakota
South Atlantic (SA): Delaware, District of Columbia, Florida, Georgia, Maryland, North Carolina, South Carolina, Virginia, West Virginia
East South Central (ESC): Alabama, Kentucky, Mississippi, Tennessee
West South Central (WSC): Arkansas, Louisiana, Oklahoma, Texas
Mountain (MTN): Arizona, Colorado, Idaho, Montana, Nevada, New Mexico, Utah, Wyoming
Pacific (PAC): Alaska, California, Hawaii, Oregon, Washington

and impact on the income distribution differed. The extent of family income inequality varied noticeably among the regions in 1994. The right side of Figure 6 indicates that the greatest degree of inequality was present in the West South Central, Middle Atlantic, and Pacific regions, where rich (90th percentile) families had over 10 times as much income as poor (10th percentile) families. (See Table 5 for a list of states in each region.) The distribution was least unequal in the West North Central and Mountain regions, in which the rich/poor ratio was around 7-to-1.

The left side of Figure 6 shows the poor family's income in each region relative to the region's median income, and the rich family's income relative to the median. The "leaders" in inequality, the West South Central, Middle Atlantic, and Pacific regions, had large spreads at both the top and bottom of their income distributions—the 10th percentile family had less than one-quarter of the income of the median family, while the 90th percentile family had 2¼ to 2½ times the income of the median family. The 90th/10th ratio measure of inequality is especially sensitive to differences at the bottom, and what distinguishes the low-inequality Mountain and West North Central regions is that the 10th percentile family had about 30 percent of the median family's income.

While inequality rose in all regions, their relative rankings were not entirely stable. Figure 7 summarizes the evolution of nonelderly family income inequality in the nine regions over roughly five-year intervals from 1973 to 1994. Inequality increased in most of the regions in most of the periods shown in the chart (the major exception being the mid to late 1980s), and the increases were due predominantly to a deterioration in the incomes of poor families relative to the median in their region.

While some of the regions' relative rankings moved over time, the picture in Figure 7 is not one of widespread reversals of inequality rankings. Thus, for example, the New England, West North Central, and Mountain regions retained relatively low inequality over the entire period, while the Pacific, West South Central, and East South Central regions were fairly consistently at the high end. The next section compares the inequality of family incomes with wage inequality in the regions, and the following section discusses some of the economic factors that might account for these regional patterns.

Wage Inequality among the Regions

Figure 8 compares levels of wage inequality and family income inequality among the nine regions and the United States in selected years. The scatterplot shows a fairly strong positive association—region-years with higher wage inequality typically show higher family income inequality as well. Some of the association simply reflects the fact that both types of inequality rose over time. Nonetheless, all the points do not fall on a line, so factors other than wage inequality contribute to the inequality of family incomes.

Figure 9 plots the time patterns of wage and family income inequality for four regions: New England, Pacific, Mid-Atlantic, and East South Central. Each region's level and pattern of year-to-year changes in wage inequality are roughly, but far from exactly, reflected in family income inequality.

The Regional Economies

In considering how regional economic performance might be associated with regional inequality, economists can draw on two bodies of research relating national economic performance to inequality: studies of inequality over the business cycle and studies of inequality differences among nations at different stages of economic development. This latter research asks whether economic growth and inequality are complementary or competitive. The debate often traces its roots to Kuznets (1955), who argued

Figure 7

Family Income Inequality by Region, 1973 to 1994, Selected Years

See Table 5 for region definitions.

[a] Inequality measured as ratio of 90th percentile family income to 10th percentile.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, March 1980, March 1985, March 1990, and March 1995.

that inequality rises at the early stages of economic development and industrialization, and falls only later as growth continues. More recently, a consensus seems to have developed on the facts (which are inconsistent with Kuznets' hypothesis): In a broad cross section of nations, greater economic growth is associated with lower inequality. The mechanisms that underlie this association, and even the direction of causation, are still hotly debated, however. (See, for example, articles by Albelda and Tilly 1995; Birdsall, Ross, and Sabot 1995; and Chang 1994.)

Most U.S. regions' economic fortunes are linked to national business cycles. Local manufacturers sell their products in national and international markets, and other local businesses and consumers are affected by nationwide factors such as interest rates and exchange rates. Nonetheless, regions do not move in lockstep with the nation. Occasionally a region will experience a recession or boom that does not mirror the national cycle. More frequently, the timing and amplitude of the cycle will differ among the regions.

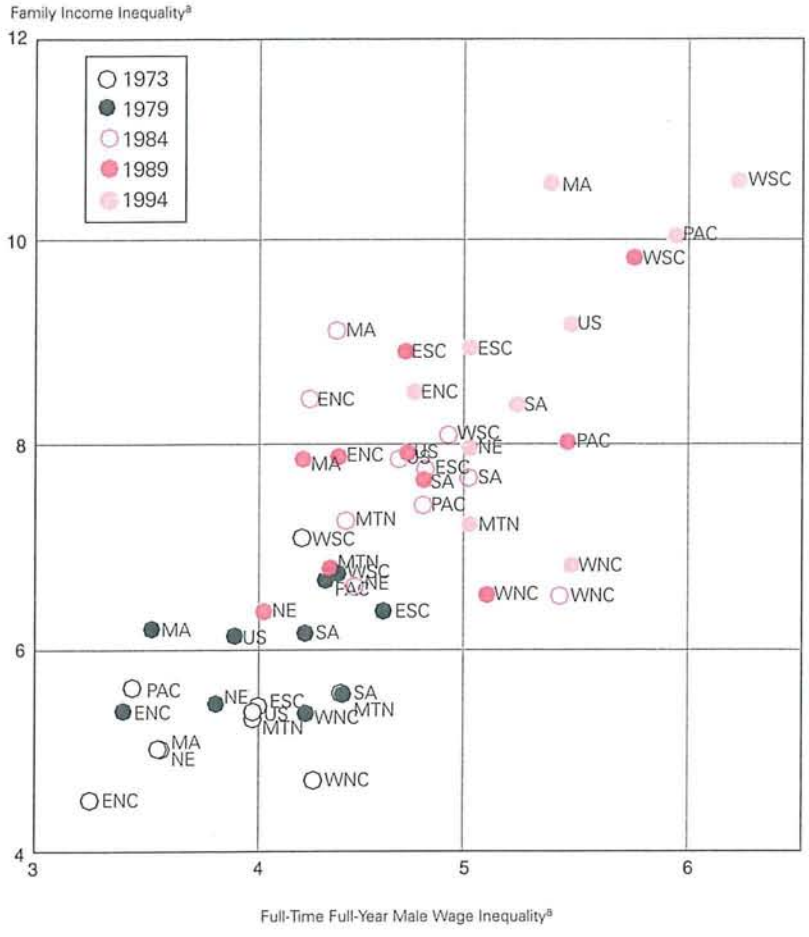
For example, while the country as a whole did not see falling inequality during the 1980s expansion, the experience of New England as compared with other regions is consistent with the "typical" cyclical pattern. New England experienced an economic boom of unusual proportions, enjoying unemployment rates in the 3 to 4 percent range in the late 1980s while the national rate was running at 5 to 6 percent. Although income inequality rose in all regions during the 1980s, it rose least in New England. Indeed, New England was the only one of the nine Census divisions in which the income of the 10th percentile family rose in real terms between 1979 and 1989; in the other eight regions, the income of poor families fell.

Similarly, Texas (in the West South Central region) and the other oil states suffered a severe economic decline while the rest of the nation expanded in the early to mid 1980s; Figure 7 shows a steep increase in inequality there between 1984 and 1989 (when inequality was declining in some other regions). The "rust belt" East North Central and Middle Atlantic states enjoyed their fastest employment growth in the 1984–89 period—initially reflecting their recovery from the severe recession of 1981–82—and simultaneously experienced declines in income inequality.

Figure 8

Inequality of Wages and Family Incomes

Scatter Plot by Region, Selected Years 1973 to 1994



[a] Inequality measured as ratio of 90th percentile income or earnings to 10th percentile.

See Table 5 for region definitions.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, March 1980, March 1985, March 1990, and March 1995.

New England's boom was followed by a downturn that began earlier and was much more severe than the national recession of 1990–91, in which the region lost one in 10 jobs and saw inequality rise noticeably. The Mountain states, by contrast, suffered virtually no employment loss in the early '90s recession and showed only a very small rise in inequality between 1989 and 1994.

These regional patterns are consistent with the hypothesis that economic growth reduces inequality while tough times are accompanied by rising inequality. When a region's pace or direction of economic growth and decline does not match the national cycle, the relative degree of inequality in the region often shifts accordingly.

Regions also vary considerably in their industry mix, another factor hypothesized to be associated with earnings inequality. All the regions saw a sizable downtrend in the fraction of employment in manufacturing over the past two decades. Before the recession of 1973–75, the East North Central, East South Central, and New England regions had the highest fraction of nonagricultural employment in manufacturing (30 percent or more), and the Mountain states were alone at the other extreme, with less than 15 percent. By 1994, only the East North Central and East South Central regions had more than 20 percent manufacturing jobs. The Mountain region's manufacturing share had fallen the least but remained at the low end with about 10 percent of nonfarm jobs in manufacturing.[26]

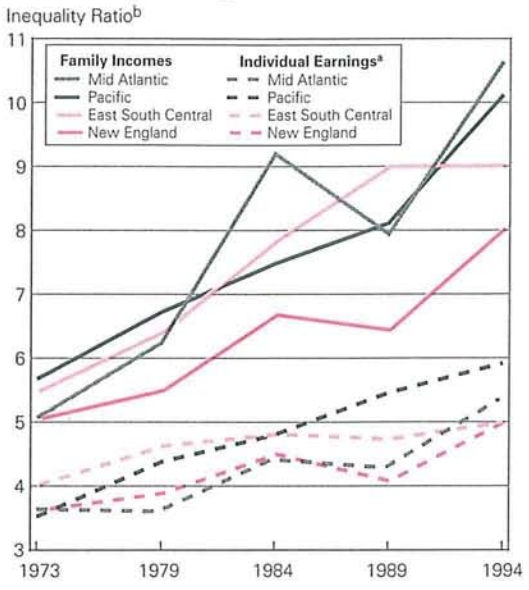
Patterns of part-time work also vary among the regions. In the nation as a whole, the fraction of workers on part-time schedules rose from 14.6 percent in 1973 to 15.3 percent in 1994. The part-time fractions in the West North Central and

[26] If industry mix were a crucial determinant of inequality, one would expect inequality to be higher in the Mountain states than in New England, the East North Central, and East South Central regions. If changes in industry mix were important, then inequality would be rising everywhere, but relatively high in New England and low in the Mountain states, East North Central, and East South Central. Figure 7, however, shows the New England and Mountain regions with consistently low inequality over the 1973–94 period, while East North Central rose from low to medium, and East South Central stayed near the middle.

Figure 9

Inequality of Family Incomes and Individual Earnings[a]

Selected Regions, Selected Years



[a] Earnings inequality measured for full-time, full-year earnings of men.

[b] Inequality measured as ratio of 90th percentile to 10th percentile.

Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, March 1980, March 1985, March 1990, and March 1995.

New England regions were consistently above the national average (over 18 percent and about 17 percent, respectively) but did not rise appreciably. The South Atlantic region's part-time fraction was below average, but also changed little, staying below 14 percent over the period.

Regional Differences in Family Structure, Work Patterns, and Income Sources

In addition to the differences in their economies outlined above, the regions have somewhat different demographic profiles, cultures, traditions, or histories that may cause their family structure and work patterns—and hence income inequality—to vary. For example, the fraction of nonelderly families headed by a single parent was highest (30 percent) in the relatively high-inequality Pacific and East South Central regions in 1994 and lowest (24 percent) in the relatively low-inequality Mountain states.

Among single-parent families, variations in family work patterns should also be associated with differences in inequality. The average number of workers in one-head families was highest in 1994 in the relatively low-inequality Mountain and West North Central regions and lowest in the relatively high-inequality Middle Atlantic states.

For some regional population characteristics, however, the relationship with family income inequality, even hypothetically, may be more complicated. The relationship between educational attainment and inequality is especially ambiguous. As discussed earlier, the premium to education reflects the interaction between the demand for educated workers and the supply, and its rise is generally attributed to faster growth in demand for college-educated workers than in their supply. Thus, if all regions faced a similar demand for educated workers, one would expect lower wage premiums and lower inequality in regions with more educated workers. However, if the demand for educated workers is higher in regions with more educated workers, as seems likely based on theories of how businesses decide where to expand or locate, the wage premium need not be lower in those regions. And to the degree that demand is increasing faster than supply, the premium to education might actually increase more rapidly in areas with more educated workers.[27] The data show the 1994 fraction of the adult population with a college degree (or more) ranging from 31 percent in relatively low-inequality New England to 18 percent in the more unequal East South Central region.

The average family's mix of income sources also varied among the regions, although not very markedly. For example, transfers as a percent of nonelderly family income varied by only 2 percentage points in 1994 from the highest region (Pacific) to the lowest (Mountain).

The next section uses multiple regression analysis to sort out and quantify the relationships between a variety of demographic and economic characteristics, on the one hand, and the degree of income inequality, on the other, both over time and among U.S. regions.

[27] A related hypothesis is more mechanical: If the national rise in the educational wage premium in the 1980s stretched out the top of the income distribution, one would expect that regions with a more educated population would have greater inequality. However, as noted in the next sentence, the college percentage is over 15 percent even in the lowest region, so increased returns to the college-educated might not be captured by the measure of inequality used here, since it reflects only the 90th and 10th income percentiles. Thus, increased returns to education might raise the incomes of the top 30 percent of families in New England and only the top 15 percent in the East South Central region; in both regions, thus, the 90th percentile's income would be higher compared to the 10th's than before the returns to education rose.

Since some of the demographic and cultural changes that are hypothesized to be causes of income inequality have moved only gradually over time and, in some cases, quite steadily in one direction, their contributions are very difficult to identify in time-series analyses of the upward trend in U.S. inequality. To the degree that they differ among regions, using regions as observations may allow some isolation of their effects.

III. Family Income Inequality: Differences over Time and among Regions

The preceding sections of this article have introduced (explicitly or implicitly) a number of hypotheses regarding causes of increased family income inequality in the United States. This section attempts to explore these hypotheses econometrically, using data on U.S. regions over the 1973–94 period. Immediately below, the data are described. The next subsection presents the econometric results and summarizes the findings regarding each group of hypotheses.

Measures of both economic prosperity and demographics and family structure seem to influence the degree of family income inequality, independent of the local degree of wage inequality.

A third subsection interprets the results, quantifying specific factors' contributions to interregional differences in inequality and to the 1973 to 1994 rise in U.S. inequality.

The Data

Measures of nonelderly family income inequality are regressed on a variety of explanatory variables using a pooled time-series, cross-section set of 90 observations: nine U.S. regions by ten years (1973, 1979, 1984, and 1988 to 1994). The regressions use data mostly drawn from the *U.S. Current Population Survey* (CPS), March Supplement, which reports demographic characteristics of families as of the survey date in March and income and work characteristics for the preceding calendar year. Table 6 reports summary statistics and sources.

Hypotheses related to the health of the local economy and economic growth are examined using measures of employment growth, unemployment rates, and median family income. Inequality-related characteristics of regional labor markets are summarized with measures of labor force participation and the prevalence of part-time work. To gauge the importance of the educational wage premium, the fraction of the adult population (age 25 and older) with a college degree is included. Industry mix is proxied by the share of employment in manufacturing.

Demographic variations among regions and over time are captured in variables tallying the fraction of (nonelderly) families headed by single parents and the fractions of the population that are black or Hispanic. Family work patterns are reflected in the average number of workers in married-couple and one-head families.

Regression Results

Table 7 reports the estimated coefficients from several equations "explaining" family income inequality, measured as the ratio of the 90th percentile nonelderly family's income to that of the 10th percentile family. Column (1) reports results for an inclusive set of economic, demographic, and family structure variables. Columns (2) through (4) try out three additional types of influences: industry mix, income composition, and wage inequality.

Table 8 examines disparities in the top and bottom of the income distribution separately, using logarithms so the results are additive. Column (1) reports an equation identical to column (1) of Table 7, except that the dependent variable is in logs. Columns (2) and (3) include the same explanatory variables but have as dependent variables the log of the 90th/50th ratio and the log of the 50th/10th ratio.

Table 9 also uses the same set of explanatory variables as column (1) of Table 7. Three alternative inequality measures are used as dependent variables: individual wage inequality as measured by the 90th/10th annual earnings ratio for full-time, full-year men; the ratio of 80th percentile family income to 20th percentile family income; and the ratio of average income in the richest quintile to average income in the poorest quintile.

Most of the variables in most of the equations shown in Tables 7 to 9 obtain coefficients that are

Table 6
Variable Means and Sources
Pooled regional cross-section time series, selected years 1973 to 1994 (N = 90)

Variable Name:	Mean	Standard Deviation	Minimum	Maximum	Source	Notes
Inequality ratio: 90th/10th	7.8	1.5	4.6	10.7	CPS	These ratios are all based on money income data for nonelderly families; selected percentiles.
Logarithm of ratio: 90th/10th	2.03	.20	1.52	2.37	CPS	
Logarithm of ratio: 90th/median	.77	.07	.61	.92	CPS	
Logarithm of ratio: median/10th	1.27	.15	.91	1.55	CPS	
Inequality ratio: 80th/20th	3.5	.5	2.5	4.5	CPS	
Quintile average income ratio: richest/poorest	9.0	1.6	5.4	12.2	CPS	Ratio of average income of richest fifth of families to poorest fifth
Wage inequality (male 90th/10th)	4.8	.6	3.3	6.2	CPS	Annual earnings, full-time, full-year men.
% change in employment from year earlier	2.6	2.2	−5.0	7.6	BLS	Nonagricultural employment
Unemployment rate previous year (%)	6.4	1.6	3.1	12.3	BLS	Unemployed/labor force, lagged
Labor force participation rate (%)	65.5	3.0	57.5	71.8	BLS	Labor force/civ. noninst. pop.
Percentage of workers part-time	15.5	1.8	12.1	19.2	CPS	
Median family income ($000)	40.3	4.8	30.3	54.8	CPS	See note on ratios above
% adults (25+) with college degree	19.9	4.6	8.5	30.5	CPS	
% families headed by single parents	25.0	4.5	12.7	32.7	CPS	
% population black	11.0	6.8	1.9	22.8	CPS	
% population Hispanic	7.7	6.9	.2	24.7	CPS	
Avg. no. of workers, married-couple families	2.0	.1	1.8	2.1	CPS	
Avg. no. of workers, one-head families	1.3	.1	1.1	1.5	CPS	
% employment in manufacturing	18.9	5.0	10.2	33.7	BLS	% of nonagricultural employment
% of income from earnings	89.0	1.2	86.9	93.6	CPS	Earnings = wages & salaries plus self-employment income

Note: CPS is U.S. Bureau of the Census, *Current Population Survey*, March, various years. BLS is U.S. Bureau of Labor Statistics, household and establishment series.

significantly different from zero and of the expected signs. The results are discussed and interpreted in the following subsections.

The economy. Most of the hypotheses related to the health and structure of the economy are supported by the data. In these equations, regions (years) in which employment expanded faster from the prior year have lower inequality, other things equal. In addition, where (or when) median family income is higher, income inequality tends to be lower, all else equal.[28] Furthermore, higher unemployment (lagged one year) is associated with greater inequality, reflecting the fact that unemployment is not spread evenly across the work force.

This uneven impact is revealed further in Table 8, where the estimated coefficients indicate that higher unemployment has no discernible effect on the spread at the top of the distribution, but is associated with greater disparities in income between middle-income families and poor families. Employment growth also has its impact on inequality through raising the bottom of the distribution relative to the median, not by raising the median relative to the top (or bringing down the top). By contrast, most of the inequality-lowering impact of a higher median income appears to work through compression of the top of the distribution rather than the bottom.[29]

When percent of employment in manufacturing

[28] Note that the income figures are adjusted for inflation using the U.S. consumer price index, but they are not corrected for regional differences in the cost of living. These differences can be substantial. For example, consumer prices rose about 10 percent more in Boston during the 1980s than in the nation as a whole. (Neither the U.S. adjustment nor regional adjustments would affect the $90^{th}/10^{th}$ percentile inequality measure because it is a ratio of two income numbers, both multiplicatively adjusted in the same way.)

[29] Since the median is the numerator of the dependent variable in column (3) of Table 8, however, it may be that the indistinguishable-from-zero impact reflects a truly inequality-lowering impact of a higher median offset by the arithmetic effect of a higher numerator in the measure. The arithmetic and inequality effects are reinforcing for the dependent variable in column (2), in which the median is the denominator.

Table 7

Regression Results – Family Income Inequality

Pooled regional cross-section time series, selected years 1973 to 1994
(Standard errors in parentheses below estimated coefficients)

Explanatory Variables:	Dependent Variable: Ratio: 90th/10th (1)	Ratio: 90th/10th (2)	Ratio: 90th/10th (3)	Ratio: 90th/10th (4)
Constant	3.7	3.6	−10	2.4
	(2.9)	(2.9)	(8)	(2.6)
% change in employment from year earlier	−.089***	−.087***	−.088***	−.090***
	(.033)	(.034)	(.033)	(.029)
Unemployment rate (%), previous year	.14***	.13***	.16***	.13***
	(.04)	(.04)	(.04)	(.03)
Labor force participation rate (%)	−.21***	−.21***	−.22***	−.19***
	(.04)	(.04)	(.04)	(.04)
Percentage of workers part-time	.26***	.27***	.25***	.21***
	(.08)	(.08)	(.08)	(.07)
Median family income (000)	−.073***	−.080***	−.083***	−.014
	(.025)	(.028)	(.025)	(.025)
% adults (25+) with college degree	.054	.065	.079*	.0050
	(.039)	(.045)	(.041)	(.036)
% adults (25+) with col. degree, post-1979	.044***	.045***	.039***	.030**
	(.014)	(.015)	(.014)	(.013)
% families headed by single parents	.12***	.11***	.15***	.12***
	(.03)	(.03)	(.03)	(.03)
% population black	.12***	.12***	.11***	.090***
	(.02)	(.02)	(.02)	(.021)
% population Hispanic	.12***	.12***	.12***	.089***
	(.02)	(.02)	(.02)	(.015)
Avg. no. of workers, married-couple families	6.4***	6.5***	6.2***	4.6***
	(1.9)	(1.9)	(1.8)	(1.7)
Avg. no. of workers, one-head families	−2.6**	−2.7***	−2.1**	−2.5***
	(1.0)	(1.0)	(1.0)	(.9)
% of employment in manufacturing		.0095		
		(.019)		
% of income from earnings			.15*	
			(.08)	
Wage inequality ratio: male earnings 90th/10th				.76***
				(.16)
No. of observations	90	90	90	90
Adjusted R-squared	.90	.90	.90	.92

*Significantly different from zero with 90 percent confidence.
**Significantly different from zero with 95 percent confidence.
***Significantly different from zero with 99 percent confidence.
Source: Author's calculations; see Table 6 for variable definitions.

industries is included (Column (2) of Table 7), it fails to obtain an estimated coefficient that is significantly different from zero.[30] This lack of impact may partly reflect the fact, noted earlier, that the earnings distributions of manufacturing and services became more alike over the period. Nonetheless, despite the strong association suggested by national trends (loss of manufacturing jobs accompanied by rising inequality), this time-series cross section of regions finds no independent effect of the composition of jobs among industries on family income inequality once the other included economic and demographic variables are controlled for.

Higher labor force participation rates are also associated with lower inequality. Apparently the broader the involvement of the working-age population in the labor market, the more evenly shared are incomes. And regions (years) with more workers on part-time schedules have a more unequal income distribution because part-time workers' lower annual earnings pull down the bottom of the distribution.

Thus, the typical cyclical pattern shows up in these pooled time-series, cross-section data, as does the pattern typically found in international data. That is, prosperity (high family income, low unemployment rate, faster employment growth, more people working, and working more hours) is associated with lower inequality, other things equal. However, the economy's structural shift away from manufacturing and toward services is not part of the story.

Educational attainment. The educational wage premium also shows clearly in these data. Regions (years) with a higher fraction of the population college-educated show more inequality, other things equal. Because the earnings inequality literature finds that the educational wage premium rose in the 1980s but not the 1970s, a separate college coefficient is estimated for the years after 1979.[31] While the college coefficient is

[30] Similarly, variables measuring percent of employment in services industries or the change in either manufacturing or services' share of employment obtain coefficients that are statistically indistinguishable from zero (results not shown).

always significant different from zero in the 1980s and '90s, in most of the equations reported here, educational attainment has no significant effect in the 1970s. Reflecting the fact that worker(s) in the 10th percentile family typically are not college-educated,[32] the decomposition of inequality in Table 8 suggests that regions (years) with more college-educated residents have a higher median income and an even higher (relative to the median) 90th percentile income.

As noted earlier, the growing college wage premium reflects faster growth in the *demand* for college-educated workers than in the *supply*. As the emphasis on education rises, employers would be increasingly attracted to locate or expand in regions with a more educated work force. In this case, the coefficient would reflect proportionally greater increases in demand for educated workers (hence higher incomes for them) in regions like New England that have relatively high levels of educational attainment than in the East South Central and West South Central regions with lower attainment.[33]

Population and family characteristics. Demographic factors are strongly associated with income inequality. Inequality is higher where (when) there are more single-parent families. Inequality is also higher where the fraction of the population that is black or Hispanic is higher. All of these groups (single-parent families, black families, and Hispanic families) are typically underrepresented at the top of the income distribution. Interestingly, the decomposition in Table 8 suggests that, of these demographic variables, only the single-parent fraction has its dominant impact on inequality through pulling down the bottom of the income distribution. The other variables are associated with the spread at the top of the distribution as well as at the bottom, although the estimated coefficients for the bottom—the median/10th equation—are larger than for the top.

Table 8
Regression Results — Family Income Inequality Decomposition
Pooled regional cross-section time series, selected years 1973 to 1994
(Standard errors in parentheses below estimated coefficients)

	Dependent Variable		
Explanatory Variables:	Logarithm of Ratio: 90th/10th (1)	Logarithm of Ratio: 90th/median (2)	Logarithm of Ratio: Median/10th (3)
Constant	1.4***	.59***	.76**
	(.3)	(.15)	(.30)
% change in employment from year earlier	−.010**	.0011	−.011***
	(.004)	(.0017)	(.003)
Unemployment rate (%), previous year	.018***	.0015	.017***
	(.005)	(.0021)	(.004)
Labor force participation rate (%)	−.025***	−.0046**	−.020***
	(.005)	(.0023)	(.004)
Percentage of workers part-time	.031***	−.00085	.032***
	(.010)	(.0044)	(.009)
Median family income (000)	−.0096***	−.0079***	−.0017
	(.0029)	(.0013)	(.0026)
% adults (25+) with college degree	.0073	.0061***	.0013
	(.0046)	(.0021)	(.0041)
% adults (25+) with col. degree, post-1979	.0064***	.0019**	.0045***
	(.0017)	(.0008)	(.0015)
% families headed by single parents	.016***	.0032**	.013***
	(.003)	(.0016)	(.003)
% population black	.015***	.0036***	.012***
	(.003)	(.0012)	(.002)
% population Hispanic	.015***	.0051***	.010***
	(.002)	(.0008)	(.002)
Avg. no. of workers, married-couple families	.80***	.24**	.56***
	(.22)	(.10)	(.19)
Avg. no. of workers, one-head families	−.33***	.0094	−.34***
	(.12)	(.053)	(.11)
No. of observations	90	90	90
Adjusted R-squared	.93	.87	.89

*Significantly different from zero with 90 percent confidence.
**Significantly different from zero with 95 percent confidence.
***Significantly different from zero with 99 percent confidence.
Source: Author's calculations; see Table 6 for variable definitions.

[31] The post-1979 coefficient is additive; that is, the estimated effect of college in the 1980s to '90s is indicated by the sum of the two coefficients. When the post-1979 coefficient is significantly different from zero, it indicates that the post-1979 effect of college is significantly higher than the 1970s effect.

[32] Table 2 shows five to seven times as many college graduates in the richest quintile of families as in the poorest quintile.

[33] The more mechanical explanation is that families with college-educated workers, in any region, pull ahead of those without a

Table 9
Regression Results — Alternative Dependent Variables
Pooled regional cross-section time series, selected years 1973 to 1994
(Standard errors in parentheses below estimated coefficients)

Explanatory Variables:	Dependent Variable: Wage Inequality Ratio: Male Earnings 90th/10th (1)	80th/20th Percentile Family Income Ratio (2)	Quintile Average Family Income Ratio: Richest/Poorest (3)
Constant	1.7	2.7***	7.3**
	(1.9)	(.9)	(2.8)
% change in employment from year earlier	.0015	−.018*	−.049
	(.021)	(.010)	(.032)
Unemployment rate (%), previous year	.0087	.033***	.12***
	(.025)	(.012)	(.04)
Labor force participation rate (%)	−.025	−.040***	−.18***
	(.028)	(.013)	(.04)
Percentage of workers part-time	.072	.063**	.21**
	(.054)	(.026)	(.08)
Median family income (000)	−.078***	−.033***	−.12***
	(.016)	(.007)	(.02)
% adults (25+) with college degree	.065**	.018	.036
	(.025)	(.012)	(.038)
% adults (25+) with col. degree, post-1979	.018**	.011**	.063***
	(.009)	(.004)	(.014)
% families headed by single parents	−.0024	.039***	.12***
	(.019)	(.009)	(.03)
% population black	.041***	.032***	.11***
	(.015)	(.007)	(.02)
% population Hispanic	.043***	.035***	.11***
	(.010)	(.005)	(.02)
Avg. no. of workers, married-couple families	2.3*	.93	4.6**
	(1.2)	(.57)	(1.8)
Avg. no. of workers, one-head families	−.24	−.27	−1.4
	(.65)	(.31)	(1.0)
No. of observations	90	90	90
Adjusted R-squared	.72	.90	.92

*Significantly different from zero with 90 percent confidence.
**Significantly different from zero with 95 percent confidence.
***Significantly different from zero with 99 percent confidence.
Source: Author's calculations; see Table 6 for variable definitions.

Inequality is higher where (when) the average number of workers in married-couple families is higher and where the average number of workers in one-head families is lower. Given the data examined earlier showing that most one-head families are near the bottom of the income distribution and two-earner married-couple families are more prevalent in the middle and top, a simple interpretation would be that more work among one-head families brings up the bottom while more workers in married-couple families stretch the top of the income distribution. Table 8 tells a slightly more complicated story for married-couple families, however: More work effort by members of married-couple families stretches the bottom of the distribution, presumably by raising the median relative to the 10th percentile income, and stretches the top by raising the 90th percentile income even more.

The mix of income sources barely influences the degree of family income inequality. Column (3) of Table 7 reports a version of the equation that includes the fraction of income from earnings. The coefficient is positive (regions or years in which earnings comprise a greater share of family income tend to have higher inequality, other things equal), but significantly different from zero at only the 10 percent level. Other measures of income mix show no relationship with inequality.

The final column of Table 7 reports an equation that adds wage inequality to the list of variables. While this equation is an interesting descriptive exercise, the usual caveats regarding causative interpretations for estimated coefficients apply with particular force here. The most striking result, however, is how little the other coefficient estimates are affected by the inclusion of wage inequality (measured among full-time, full-year male workers), given that many of the explanatory variables would be expected to have their effects on family income inequality by affecting wage inequality. While the coefficients on a number of variables are somewhat smaller in column (4) than in column (1), the only coefficient that falls to zero is the one on median family income. Thus, measures of both economic prosperity and demographics and family structure seem to influence the degree of family in-

college degree and stretch out the distribution in proportion to their fraction of the population.

come inequality, independent of the local degree of wage inequality.

An equation with wage inequality as its dependent variable might be expected to shed further light on which factors' effects on family income inequality occur mostly through their effects on wage inequality, but the results (shown in column (1) of Table 9) are somewhat surprising. The estimates indicate that most of the economic factors (employment growth, lagged unemployment, labor force participation, and part-time work) are not associated with wage inequality, while one of the family variables is (workers in married-couple families).[34] In addition to these unexpected findings, the equation does confirm a rising college wage premium (the college coefficient is significantly higher after 1979 than before), as well as greater wage inequality in regions with a higher percentage black or Hispanic population, and higher median income, other things equal.

The two most important factors associated with the observed rise in U.S. income inequality from 1973 to 1994 were the increasing fraction of families headed by single parents and the rising payoff to college education in the 1980s.

The alternative measures of family income inequality used as dependent variables in Table 9 produce results that are very similar to those for the basic family inequality measure used in Table 7. Column (2) of Table 9 uses the ratio of the 80th percentile family income to the 20th percentile family; column (3) has the ratio of average income in the richest quintile to average income in the poorest quintile. The similarity of coefficient patterns is not surprising since all three ratios moved similarly over the 1973–94 period (see Appendix A for a comparison of the measures). The employment growth coefficient is noticeably weaker in the 80th/20th equation than in the 90th/10th, and becomes indistinguishable from zero in column (3). Among the other variables, a number of the estimated coefficients are slightly weaker with the alternative inequality measures, but only the coefficients on number of workers in one-head families fall to zero.

Interpretation of Results

One way to understand what the regressions imply about inequality is to use the coefficient estimates to "explain" inequality differences among regions or the observed rise in U.S. inequality from 1973 to 1994. Multiplying actual variable values in a specific year or region by the estimated coefficients shown in column (1) of Table 7 provides a "prediction" of the inequality ratio for that year or region. Comparing those predictions over time or across regions allows the predicted change (or difference) in inequality to be associated with changes (or differences) in specific factors.

Nationally, the ratio of 90th percentile income to 10th percentile income rose by 3.8 (from 5.4 to 9.2) between 1973 and 1994. The predicted value (based on U.S. variable values) rose by almost as much (3.5). The two most important factors associated with the rise were the increasing fraction of U.S. families headed by single parents and the rising payoff to college education in the 1980s (see the first panel of Table 10).[35] An expansion in the Hispanic fraction of the population was also a factor, as was the rising number of workers in the average married-couple family. Changes in the economy (employment growth was slower and the unemployment rate and part-time fractions were higher in 1994 than in 1973) also contributed to the rise in inequality, but to a lesser extent, and so did a slight

[34] Since the wage inequality measure is for full-time, full-year workers, the zero effect of labor force participation and part-time work is not surprising—those variables determine whether or not a worker is considered full-time, full-year.

[35] Wage inequality also rose over the 1973–94 period, but separating out the direct effect of rising wage inequality does not change the nature of these results. If the figures in Table 10 were computed using the coefficients from the equation including wage inequality (column (4) of Table 7 instead of column (1)), the "college" contribution would be 0.7 instead of 1.6; the direct contribution of changes in the wage inequality variable is 1.1; the other estimates are essentially unchanged. Thus, the effect of the rise in single parenting on family income inequality is virtually the same whether or not wage inequality is included in the equation, consistent with the expectation that family structure would affect family income inequality but not the distribution of wages. College education, by contrast, known to have contributed substantially to the rise in wage inequality, makes a smaller independent addition to family income inequality once wage inequality's effect is removed. And when included, the increase in wage inequality itself (presumably including whatever part of that increase is attributable to the growing college premium) is the second most important factor associated with rising family inequality.

Table 10
Major Factors in Family Income Inequality

	Difference or Change in Inequality Ratio
I. Trend: Increase in U.S. family income inequality, 1973–94	
Actual	+3.8
Predicted (fitted value of regression)	+3.5
Estimated contribution of 1973–94 actual U.S. increase in:	
% families headed by single parents	+1.5
% adults (25+) with college degree[a]	+1.6
% population Hispanic	+.6
Avg. no. of workers, married-couple families	+.3
Labor force participation rate	−1.3
II. Cycle: Increase in U.S. family income inequality, 1990–92	
Actual	+.6
Predicted	+.6
Estimated contribution of 1990–92 actual U.S. change in:	
Employment growth	+.1
Unemployment rate (lagged one year)	+.2
III. Regional level: Difference between New England average and all-regions average	
Actual	−1.0
Predicted	−1.0
Estimated contribution of difference between New England and all regions in:	
% population black	−.9
Median family income	−.6
Labor force participation rate	−.5
% population Hispanic	−.5
Avg. no. of workers, married-couple families	+.8
% adults (25+) with college degree[a]	+.5
% of workers part-time	+.4
IV. Regional boom: Decrease in New England inequality, 1984–89	
Actual	−.3
Predicted	−1.5
Estimated contribution of 1984–89 actual New England change in:	
Median family income	−.7
Unemployment rate (lagged one year)	−.5
Labor force participation rate	−.4
Avg. no. workers, one-head families	−.3
% families headed by single parents	−.3
V. Regional bust: Increase in New England inequality, 1989–94	
Actual	+1.6
Predicted	+2.7
Estimated contribution of 1989–94 actual New England change in:	
Median family income	+.4
Unemployment rate (lagged one year)	+.5
Labor force participation rate	+.2
Avg. no. workers, one-head families	+.8
% families headed by single parents	+.7

[a]College "contribution" reflects both rise in fraction with college and rising return to college (higher estimated college coefficient after 1979).

Note: Inequality measured as ratio of 90th percentile income to 10th. "Contributions" are based on coefficients reported in Table 7, column (1).

Source: Author's calculations; see Table 6 for variable definitions.

decline in the average number of workers in one-head families and a small rise in the black fraction of the population.[36] Over this period, the overall labor force participation rate rose; in the absence of this improvement, family income inequality in the United States would have risen even more than it did.

The factors the equations identify as most important in explaining the two-decade rise in inequality are, not surprisingly, trend variables—increases in single-parenthood and college graduation rates have cumulated over a relatively long time period. But other variables may be more important in explaining interregional or year-to-year variations in inequality. For example, both employment growth and the unemployment rate move considerably over the business cycle, but were only marginally less favorable in 1994 (an expansion year) than in 1973 (a pre-recession peak). The rise in unemployment and decline in employment growth that occurred during the 1990–91 recession, however, accounted for about one-half of the increase in inequality that occurred between 1990 and 1992 (see second panel in Table 10).[37]

Across all the regions, the greatest variations in inequality are attributable to interregional differences in racial and ethnic mix, the prevalence of part-time work, and labor force participation in general and especially among married-couple families.[38] The estimated coefficients indicate, for example, that New England's comparatively low family in-

[36] These factors making smaller estimated contributions—about 0.2 each—are not shown in Table 10.

[37] U.S. employment grew 1.4 percent in 1990 and only 0.3 percent in 1992; the nation's unemployment rate rose from 5.3 percent in 1989 (recall that unemployment is lagged one year in the regression) to 6.7 percent in 1991.

[38] Differences between the highest and lowest region in either 1973 or 1994 on each of the listed variables "explains" a difference in the inequality ratio of more than 1.0. In the 1980s and '90s (but not the '70s), interregional differences in median family income and college degrees also made contributions this large.

come inequality over the entire 1973–94 period was associated with a number of factors, including its above-average median income and labor force participation, along with below-average fractions of black and Hispanic population. (See the third panel of Table 10.) Partially offsetting these factors were more workers in married-couple families, a higher percentage of college graduates than the average region, and a higher fraction of part-time workers.

The estimated coefficients can also be used to explore why New England's relative inequality fell during its economic boom and rose during the recent recession (see the final two panels of Table 10). The drop during the region's mid to late-1980s boom was associated with improvements in the economy (income rose and unemployment fell between 1984 and 1989), a rise in the region's overall labor force participation rate with an associated increase in the number of workers in one-head families, and a decline in the fraction of single-parent families. Most of these improvements unraveled in the ensuing bust: The region experienced higher unemployment, lower real family incomes, an increase in the fraction of families headed by single parents, and a decline in the average number of workers in one-head families, all of which were associated with the rise in inequality between 1989 and 1994.

IV. Conclusions

Both economic and demographic factors have contributed to increased family income inequality in the 1970s, '80s, and '90s. According to the estimated equations, the most important factors associated with the two-decade rise in inequality are a rising fraction of single-parent families and an increase in the payoff to college education, offset by swelling labor force participation. Also contributing was a rise in the number of workers in married-couple families. Year-to-year variations were attributable, in addition, to year-to-year changes in rates of employment growth and unemployment. Inequality differences among regions are explained by those factors plus sizable differences in racial and ethnic composition and labor market characteristics such as the prevalence of part-time work and labor force participation rates.[39]

[39] This choice of "most important" factors is based on the "predicted value" exercise for the United States reported in the previous section and on the size of estimated beta coefficients associated with the equations shown in Tables 7 to 9.

Empirical analyses such as those reported here capture relationships that may or may not reflect causation. Lacking a causal connection, if policies were to be directed at moderating or reversing this rise in inequality, those aimed at improving economic growth and facilitating and encouraging wide participation in the labor market (if such participation can be successful, not adding mostly to unemployment) would seem most promising. That is, boosting employment growth and labor force participation and reducing unemployment and (involuntary) part-time work are likely to raise the incomes of the poorest families, incomes that declined markedly in real terms over the last two decades. Furthermore, even if such efforts were not successful in reducing inequality, enhancing and steadying economic growth and tapping available human resources more effectively through greater access and education would have other salutary effects on the economy.

The payoff in terms of reduced inequality is likely to be large to any efforts that increase the employability and productivity of those currently at the bottom of the income distribution.

The demographic factors identified as important, by contrast, are not appropriately subject to policy pushes. No one would suggest reducing the number of workers in married-couple families as a method of reducing inequality. Rather, any policies undertaken might be aimed at changing the relationship between these population composition characteristics and income inequality. For example, the association between single parenthood and racial and ethnic composition, on the one hand, and income inequality, on the other, might be altered through improvements in the functioning of the economy and through policies targeted on encouraging work among families currently attached only marginally to the labor force. That is, with greater prosperity and better returns to work—stronger economic growth and broader (successful) participation in the labor market—more single-parent families and blacks and Hispanics might be able to move up the income ladder, reducing the spread, or the

weight, at the bottom of the distribution. Birdsall, Ross, and Sabot (1995) describe several mechanisms, "virtuous circles," by which policies that advance development and growth also reduce inequality by improving the economic prospects of a nation's poorest families.[40] Such arguments give a central role to education.

Education is probably the most important lever for enhancing growth and reducing inequality. While the coefficients estimated here, taken at face value, indicate that additions to a region's pool of college-educated workers are associated with increased inequality, they must be interpreted as reflecting faster growth in the demand for college-educated workers than in the supply in regions (years) with a more educated population. The policy goal, if any, must thus be to alter the relationship,[41] not to stanch the spread of higher education. The educational wage premium would decline as a result of changes in either the supply or demand for labor at various educational levels: If the supply of college-educated labor grew faster than the demand, the premium would fall. Increasing access to, and raising the quality of, elementary and secondary education in the United States would give more people the preparation needed for success in college and would also increase the quality of the non-college-educated work force. By the same token, training or schooling to provide job skills to the least prepared individuals and higher-quality high school educations could slow the rise in demand for college graduates by offering employers productive substitutes without the college credential. The payoff in terms of reduced inequality is likely to be large to any efforts that increase the employability and productivity of those currently at the bottom of the income distribution.

[40] The "circle" is completed by a second stage of the process, not examined here, through which the improvements in inequality in turn provide the means, both economic and political, for further investments (notably in increasing access to and the quality of education) that enhance growth. Albelda and Tilly (1995) point out a positive relationship between low inequality and gains in labor productivity.

[42] That is, to change the coefficient, as indeed it changed in the opposite, inequality-augmenting direction in the 1980s.

Appendix A

Measures of Family Income Inequality

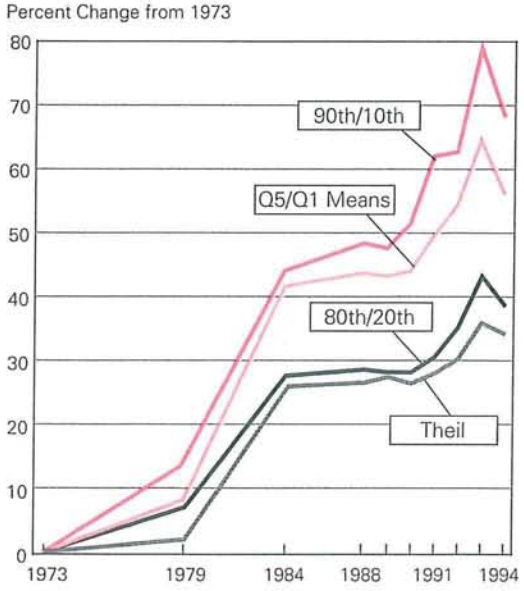
The analysis reported in this paper uses the ratio of income of the 90th percentile family to the income of the 10th percentile family to summarize the degree of family income inequality in a particular year or region. The major advantage of this measure is its simplicity; it is extremely easy to understand what differences in this ratio mean. In addition, the ratio is invariant to multiplicative transformations, an important characteristic when comparing inequality over time or across regions, when (where) levels of income may differ considerably.

Figure A compares the 90th/10th percentile income ratio for the United States with several other measures used in the literature—the ratio of average income in the richest quintile to average income in the poorest quintile, the 80th/20th percentile income ratio, and Theil's measure of income inequality. The chart shows the measures, indexed to their 1973 values, for 1973, 1979, 1984, and annually 1988 through 1994. All four measures move in tandem over the period—rising more steeply from 1979 to 1984 than from 1973 to 1979, leveling out in the 1984–89 period, and falling in 1994. The exact timing and amplitude of their movements, however, does not coincide, and the measure used in this study—the 90th/10th ratio—rose the most in percentage terms.

Figure A

Measures of Family Income Inequality

Selected Years



Source: Author's calculations based on U.S. Bureau of the Census, *Current Population Survey*, March 1974, March 1980, March 1985, March 1989 through March 1995.

Theil's measure, based on information theory, is preferred by many analysts on conceptual grounds. It reflects the shape of the entire income distribution (unlike the 90th/10th ratio or any measure based on only two points in the distribution) although it is more sensitive to income differences at the low end of the distribution. Theil's index also has useful decomposition properties. But it is much more complicated to calculate and to understand than the 90th/10th ratio.

The interpretation of the ratio of average incomes in the richest and poorest quintiles is just as uncomplicated as the 90th/10th ratio. Indeed, the 90th and 10th percentiles are the median incomes within the richest and poorest quintiles. Thus, the discussion that follows points out the relative disadvantages of using quintile means rather than quintile medians to reflect the typical well-being of families in each quintile. A major drawback of the ratio of richest to poorest quintile income averages is that averages are skewed by outliers, which the 90th and 10th percentiles (quintile medians) are not. In the poorest quintile, the average income can be pulled down substantially by negative incomes, but the 10th percentile income is positive in all cases. At the top, a few very high-income families can skew the highest quintile's average income.

Furthermore, income data in the U.S. Current Population Survey are top-coded: Income values above a specific amount (the top code) are reported as equal to the top code. Top-coding obviously biases downward the average income in the richest quintile.[42] But since top-coding affects only the richest 1 to 5 percent of families, the 90th percentile income is not affected.

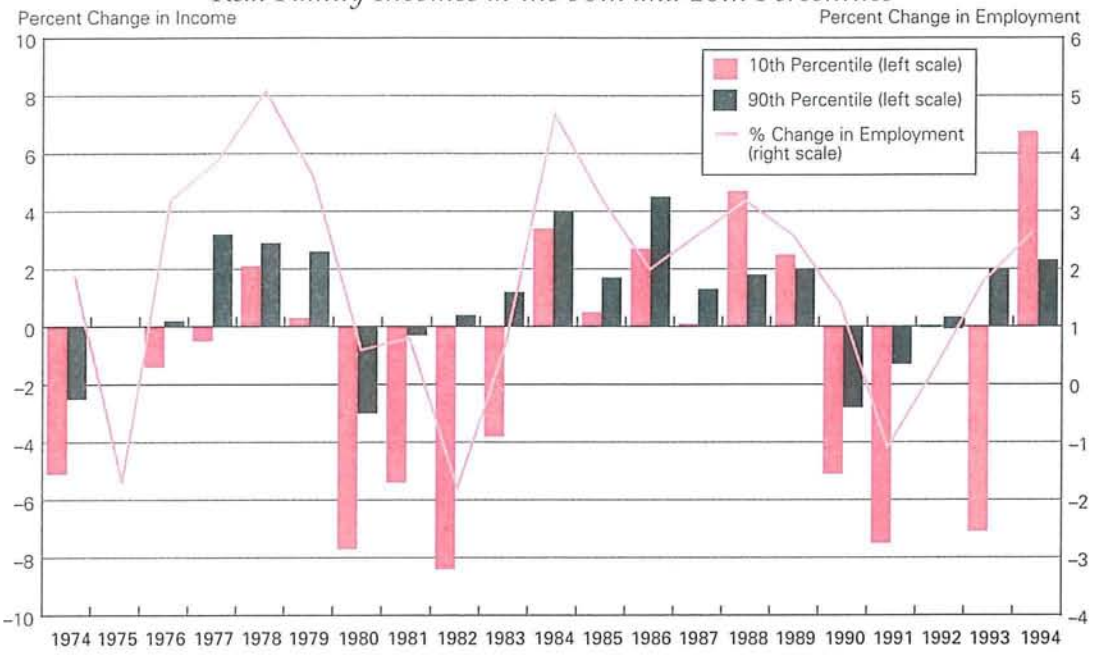
The ratio of income of the 80th percentile family to income of the 20th percentile family is, by definition, less sensitive to changes in the lowest and highest one-fifth of the distribution than the 90th/10th ratio. The choice between the two therefore comes down to choosing a balance between over-sensitivity to outliers (which one would probably see in the 95th/5th ratio) and capturing changes in the relative positions of "poor" and "rich" families.

The regressions reported in Tables 7 and 9 provide another means of comparison between the 90th/10th ratio and two of the alternatives—the 80th/20th and the richest/poorest quintile averages. Columns (2) and (3) of Table 9 report equations specified to be identical to that in Column (1) of Table 7, except the dependent variables in Table 9 are the alternative measures of inequality. The general coefficient patterns are very similar in the equations. The alternative measures of inequality, however, are less strongly related to employment growth, part-time work, and work patterns in one-head families than is the 90th/10th ratio. This finding may reflect these factors' strong effects on the bottom of the income distribution and the greater sensitivity of the 90th/10th ratio to differences in income at the bottom.

[42] Despite this downward bias, the ratio of richest to poorest quintile income averages is greater than the ratio of richest to poorest quintile medians (90th/10th). This occurs because outliers raise the ratio of averages (pulling the average income in the top quintile up and the bottom quintile's average down) more than top-coding reduces it (pulling the top-quintile average down).

Figure B

Year-to-Year Changes in Total Employment and in Real Family Incomes at the 90th and 10th Percentiles



Note: Income data not available for 1975; 1976 bars show 1974-76 change at annual rate. Constant-dollar income calculated using U.S. CPI-U-X1.
Source: Author's calculations based on U.S. Bureau of Labor Statistics, *Employment and Earnings*, and U.S. Bureau of the Census, *Current Population Survey*, March 1974 through March 1995.

Appendix B

Year-by-Year Changes in the Incomes of Rich and Poor Families

Figure B shows year-to-year changes in total employment along with income changes for families at the 10th and 90th percentiles. In all the recession years (1974–76, 1980–82, and 1990–91), real incomes declined at the bottom of the distribution, but in some of those years incomes rose at the top. In non-recession years, real incomes at the top always increased, while those at the bottom sometimes fell. Out of two decades of annual income changes, in only three years—1988, 1989, and 1994—did the incomes of poor Americans grow faster than the incomes of the rich.

Appendix C

1979 Data for Income Characteristics by Type of Family and Family Work Characteristics by Income Quintile

Figure C reports 1979 characteristics by family type corresponding to the 1994 data shown in text Figure 4. Table C decomposes the 1979 disparities in average family earnings among the quintiles into differences in average numbers of workers, annual work hours per worker, and earnings per hour, as did text Table 4 for 1994.

Table C

Family Work Characteristics by Income Quintile, 1979

Family Income Quintile	Average Annual Family Earnings	Average Number of Family Workers	Average Annual Hours per Worker	Average Earnings per Hour
Poorest	$ 8,450	1.1	1,380	$ 5.56
Second	24,500	1.6	1,650	9.03
Middle	37,650	1.9	1,720	11.78
Fourth	50,950	2.1	1,750	13.96
Richest	80,000	2.4	1,760	18.61
All	$40,300	1.8	1,680	$13.10
Ratio: Highest to Lowest	9.5	2.2	1.3	3.3

Note: Average annual family earnings rounded to nearest $50; average hours rounded to nearest 10.
Source: Author's calculations from U.S. Bureau of the Census, *Current Population Survey*, March 1980.

Figure C

Income Characteristics by Type of Family, 1979

All families with nonelderly head

Key

- Number in thousands (share (%) of all) — 50,907 (100.0)
- % in lowest income quintile~highest — 20 ~ 20
- Average # of members, # of workers — 3.4, 1.8
- Percent with children — 63
- Average annual family work hours — 3,100
- Average family income ($) — 45,100

Married-couple families

41,707 (81.9)
13 ~ 23
3.5, 1.9
61
3,300
49,500

Husband and wife work

25,675 (61.6)[a]
8 ~ 27
3.5, 2.4
61
4,000
53,200

Husband or wife works

14,612 (35.0)[a]
18 ~19
3.7, 1.4
64
2,500
45,500

Neither husband nor wife works

Other adult works	No adult works
274 (.7)[a]	1,146 (2.7)[a]
27 ~ 8	68 ~ 4
4.3, 1.3	2.8, 0
52	27
1,600	0
35,500	21,100

One-head families

9,200 . 8.0)
53 ~ 5
3.1, 1.3
74
1,900
25,100

Head works

6,857 (74.5)[b]
43 ~ 6
2.9, 1.6
73
2,400
29,000

Head does not work

Other adult works	No adult works
726 (7.9)[b]	1,617 (17.6)[b]
58 ~ 4	94 ~ 0
3.8, 1.4	3.3, 0
61	87
1,600	0
24,600	8,800

Note: hours and income rounded to nearest 100; demographic characteristics as of March 1980.

[a]Percent of married-couple families.
[b]Percent of one-head families.

Source: Author's calculations based on U.S. Census Bureau, *Current Population Survey*, March 1960.

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Correction

Correction, May/June 1996 *New England Economic Review*

In the article "Technology and Skill Requirements: Implications for Establishment Wage Structures," by Peter Cappelli, incorrect summary statistics were given for Tables 2, 3, and 4. The correct figures are as follows:

Table 2, on page 149: $R^2 = .47$, $\overline{R^2} = .45$, $F = 20.849$

Table 3, on page 149: $R^2 = .40$, $\overline{R^2} = .38$, $F = 16.001$

Table 4, on page 150: $R^2 = .12$, $\overline{R^2} = .08$, $F = 3.12$

Please enter these corrections on your copy of the May/June 1996 issue.

The Boston Fed Has a Home on the World Wide Web

The Federal Reserve Bank of Boston has established a site on the World Wide Web (WWW). The new "home page" is a gateway to Federal Reserve economic and statistical information and offers on-line access to most Boston Fed publications.

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