TRENDS IN U.S. WAGE INEQUALITY: REVISING THE REVISIONISTS

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Abstract—A recent "revisionist" literature characterizes the pronounced rise in U.S. wage inequality since 1980 as an "episodic" event of the first half of the 1980s driven by nonmarket factors (particularly a falling real minimum wage) and concludes that continued increases in wage inequality since the late 1980s substantially reflect the mechanical confounding effects of changes in labor force composition. Analyzing data from the Current Population Survey for 1963 to 2005, we find limited support for these claims. The slowing of the growth of overall wage inequality in the 1990s hides a divergence in the paths of upper-tail (90/50) inequality—which has increased steadily since 1980, even adjusting for changes in labor force composition—and lower-tail (50/10) inequality, which rose sharply in the first half of the 1980s and plateaued or contracted thereafter. Fluctuations in the real minimum wage are not a plausible explanation for these trends since the bulk of inequality growth occurs above the median of the wage distribution. Models emphasizing rapid secular growth in the relative demand for skills—attributable to skill-biased technical change—and a sharp deceleration in the relative supply of college workers in the 1980s do an excellent job of capturing the evolution of the college/high school wage premium over four decades. But these models also imply a puzzling deceleration in relative demand growth for college workers in the early 1990s, also visible in a recent "polarization" of skill demands in which employment has expanded in high-wage and low-wage work at the expense of middle-wage jobs. These patterns are potentially reconciled by a modified version of the skill-biased technical change hypothesis that emphasizes the role of information technology in complementing abstract (high-education) tasks and substituting for routine (middle-education) tasks.

I. Introduction

A large literature documents a substantial widening of the U.S. wage structure during the 1980s (Bound & Johnson, 1992; Katz & Murphy, 1992; Murphy & Welch, 1992; Juhn, Murphy, & Pierce, 1993). Wage differentials by education, by occupation, and by age and experience group all rose substantially.[1] Residual wage inequality—that is, wage dispersion within demographic and skill groups—increased simultaneously. The growth of wage inequality was reinforced by changes in *nonwage* compensation leading to a large increase in total compensation inequality (Hamermesh, 1999; Pierce, 2001). These wage structure changes translated into a pronounced rise in both household income inequality and consumption inequality, implying a marked increase in the disparities of economic well-being for U.S. families (Cutler & Katz, 1992; Karoly & Burtless, 1995).

This literature reaches two broad conclusions. First, much of the rise in U.S. earnings inequality during the 1980s appears to be explained by shifts in the supply of and demand for skills combined with the erosion of labor market institutions—including labor unions and the minimum wage—that protected the earnings of low- and middle-wage workers.[2] Second, a number of influential studies argue that the surge of inequality evident in the 1980s reflected an ongoing, secular rise in the demand for skill that commenced decades earlier and perhaps accelerated during the 1980s with the onset of the computer revolution. When this secular demand shift met with an abrupt slowdown in the growth of the relative supply of college-equivalent workers during the 1980s—itself a consequence of slowing educational attainment for cohorts born after 1949 and of smaller entering labor force cohorts—wage differentials expanded rapidly (Katz & Murphy, 1992; Autor, Katz, & Krueger, 1998; Goldin & Katz, 2001; Card & Lemieux, 2001; Acemoglu, 2002).

Drawing on more recent data, however, some recent studies challenge these conclusions. Most notably, Card and DiNardo (2002) stake two dissenting claims. First, they argue that the rise of inequality during the 1980s is largely explained by nonmarket factors, most prominently, the declining real value of the minimum wage, a view that was earlier articulated by Lee (1999).[3] Second, Card and DiNardo conclude that the growth of U.S. earnings inequality was primarily a one-time ("episodic") event of the early 1980s, which plateaued by the mid-1980s and did not recur. Building on this line of argument, Lemieux (2006b) concludes that the rise of residual inequality in the 1980s was also an episodic event accounted for by the declining value of the minimum wage and that apparent increased residual inequality since the mid-1980s reflects the mechanical effects of the changing labor force composition (rising education and experience).

This "revisionist" literature has the potential to amend the description and interpretation of U.S. earnings inequality trends. If the rise of U.S. earnings inequality was a brief, nonrecurring episode of the early 1980s, the probable causes are likely to be one-time precipitating events such as the 1980s decline in the real minimum wage. Alternatively, if the growth of earnings inequality reflects a long-term movement toward greater dispersion of earnings and higher skill differentials, then it is more likely to be explained by

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[1] A narrowing of gender wage differentials is the primary exception to the widening U.S. wage structure since 1980.

[2] See Katz and Autor (1999), Goldin and Katz (2001), and Acemoglu (2002) for overviews of this literature. See Berman, Bound, and Machin (1998) and Machin and Van Reenen (1998) for international comparisons.

[3] DiNardo, Fortin, and Lemieux (1996) also conclude that labor market institutions are the most important factor explaining rising wage inequality in the 1980s, but they do not attribute the majority of the increase to this factor.

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fundamental, secular factors, affecting the supply of and demand for skills.[4]

In this paper, we reevaluate the traditional and revisionist explanations for changes in the U.S. wage inequality over the last four decades, paying particular attention to two main claims of the revisionists: (i) that the growth of inequality was an episodic rather than secular phenomenon; and (ii) that it is explained largely by nonmarket forces and the mechanical effects of labor force composition changes. We explore these issues using wage and employment data from the March Current Population Surveys (CPS) covering 1963 to 2005, the May CPS samples for 1973 to 1978 combined with the CPS Outgoing Rotation Group (ORG) files for 1979 to 2005, and decennial population Census samples for 1980, 1990, and 2000.

In partial support of the revisionist literature, we find that past is not prologue: overall wage inequality continued growing from 1990 to 2005 but at a slower pace than in the 1980s, and the secular demand increases favoring more educated workers were, by our estimates, less rapid in the 1990s and early 2000s than from the 1960s to the 1980s—though we document a rapid ongoing rise of the relative earnings of workers with postcollege education (those with graduate and professional degrees). We concur that the falling minimum wage was a contributor to rising lower-tail (50/10 wage gap) wage inequality in the 1980s.

By contrast, we find little support for strong forms of the major revisionist claims. The growth of wage inequality is not accurately described as an episodic event. Inequality in the upper half of the male wage distribution (the 90/50 wage gap) grew rapidly and nearly continuously from 1980 to 2005 at the rate of about 1 log point per year—a marked, secular phenomenon.[5] The rapid secular growth of upper-tail wage inequality is apparent even after adjusting for labor force compositional changes. By contrast, inequality in the lower half of the distribution expanded rapidly in first half of the 1980s and then reversed course thereafter. The persistent rise in upper-tail inequality belies the claim that minimum wages can provide a coherent explanation for the bulk of the rise in earnings inequality.

We find some support for the revisionists' conclusions concerning residual inequality trends. Consistent with Lemieux (2006b), we confirm that changes in labor force composition exerted an upward force on residual wage dispersion for 1989 to 2005. But this compositional effect was concentrated in the lower tail of the earnings distribution and, moreover, served to offset a rapid compression of lower-tail prices. We find that "price" changes, changes in earnings dispersion within narrowly defined demographic groups, remain a key force in the evolution of both upper- and lower-tail U.S. residual wage inequality.

An organizing theme that emerges from our review of the key facts is that, following a monotone surge of inequality during 1979 through 1987 in which upper incomes rose and lower incomes fell, changes in the U.S. earnings distribution subsequently "polarized," with a strong, persistent rise in inequality in the upper half of the distribution and a slowing (or reversal) of inequality trends in the lower half of the distribution. This polarization is seen in overall inequality, in residual inequality, and in earnings trends among workers at different education levels. The earnings of workers with a postcollege (graduate) degree relative to noncollege workers have increased rapidly and continuously since 1979. By contrast, the earnings of college-only workers (those with a four-year college degree but *without* a graduate degree) relative to high school graduates rose rapidly from 1979 to 1987 and then plateaued.[6]

If these inequality trends are *not* primarily explained by episodic institutional shocks, can they plausibly be explained by market-driven changes in the supply and demand for skills? In the final section, we provide a simple summary test of this hypothesis using Census data to analyze the evolution of employment and wage changes by skill over the 1980 to 2000 period. These data yield clear evidence that wage changes by earnings level and employment changes by skill level track each other closely in both decades. In the 1980s, during which wage growth was essentially monotone in skill, employment shares in the highest-educated and highest-paid occupations expanded substantially while employment shares in the lowest-skill occupations contracted. During the subsequent decade of the 1990s—in which earnings growth polarized—employment shares in very low- and very high-skill occupations increased while employment shares in moderately skilled occupations contracted. The roughly parallel movement of earnings and employment growth in each decade suggests that demand forces have played a key role in shaping wage structure changes during the inequality surge of the 1980s and the polarization that followed. Following Autor, Levy, and Murnane (2003) and Goos and Manning (2007), we find that these patterns may in part be explained by a richer version of the skill-biased technical change (SBTC) hypothesis in which information technology complements highly educated workers engaged in abstract tasks, substitutes for moderately educated workers performing routine tasks, and has less impact on low-skilled workers performing manual tasks.

The paper is organized as follows. Section II documents the evolution of the U.S. wage structure from 1963 to 2005.

[4] These explanations are not intrinsically at odds, and numerous studies focused on the experience of the 1980s support the view that institutions and market forces both contributed to rising inequality (Katz & Autor, 1999).

[5] Approximately 80% of the rise in 90/10 earnings inequality from 1980 to 2005 is accounted for by the rise in the 90/50 wage gap using hourly wages for all male wage and salary workers in both the CPS ORG and March data.

[6] Murphy and Welch (2001) and Lemieux (2006a) also find a "convexification" in the returns to schooling. Lemieux concludes that the majority of the rise in wage inequality from 1973 to 2005 is accounted for by increased returns to postsecondary schooling.

Section III presents time series models to assess the role of demand, supply, and institutional factors for changes in educational wage differentials and overall wage inequality. Section IV uses the kernel reweighting methods of DiNardo, Fortin, and Lemieux (1996) and Lemieux (2006b) to analyze the role of prices and labor force composition in changes in overall and residual inequality—focusing on the divergent trends in the bottom and top halves of the distribution. Section V provides summary tests of the relevance of demand shifts to wage structure changes. Section VI concludes.

II. U.S. Wage Structure Changes over the Past Four Decades: Key Facts

To summarize the basic changes in the U.S. wage structure over the last four decades, we draw on two large and representative household data sources: the March CPS and the combined May CPS and Outgoing Rotation Group samples. We describe these sources briefly here and provide the details on the construction of our analysis samples in the data appendix. The March CPS data provide reasonably comparable data on prior year's annual earnings, weeks worked, and hours worked per week for four decades. We use the March files from 1964 to 2006 (covering earnings from 1963 to 2005) to form a sample of real weekly earnings for workers ages 16 to 64 who participate in the labor force on a full-time, full-year (FTFY) basis, defined as working 35-plus hours per week and forty-plus weeks per year.[7]

We complement the March FTFY series data with data on hourly wages of all current labor force participants using May CPS samples for 1973 through 1978 and CPS Outgoing Rotation Group samples for 1979 through 2003 May (CPS/ORG). From these sources, we construct hourly wage data for all wage and salary workers employed during the CPS sample survey reference week. Unlike the retrospective annual earnings data in the March CPS, the May/ORG data provide point-in-time measures of usual hourly or weekly earnings. We weight May/ORG hourly earnings data by hours worked and the appropriate CPS sampling weight to provide a measure of the entire distribution of hours paid.

As detailed in Autor, Katz, and Kearney (2005) and Lemieux (2006b), both the March and May/ORG CPS surveys have limitations that reduce their consistency over the forty-year period studied. The March CPS data are not ideal for analyzing the hourly wage distribution since they lack a point-in-time wage measure and thereby hourly wages must be computed by dividing annual earnings by the product of weeks worked last year and usual weekly hours last year. Estimates of hours worked last year from the March CPS appear to be noisy and data on usual weekly hours last year are not available prior to the 1976 March CPS. The May/ORG samples provide more accurate measures of the hourly wage distribution (particularly for hourly workers) but cover a shorter time period than the March CPS. Both the March and May/ORG CPS samples have undergone various changes in processing procedures over several decades, especially involving the top-coding of high earnings, the flagging of earning imputations, and algorithms used for allocating earnings to those individuals who do not answer earnings questions in the survey. These create challenges in producing consistent data series over time, which we have tried to account for to the extent possible to make the wage series time consistent. The major redesign of the earnings questions in the CPS ORG in 1994 is likely to have created comparability problems that we are unable to fully redress.[8]

A. *Trends in Overall Inequality*

We begin laying out basic wage structure facts in figure 1, which uses data on FTFY workers from the March CPS to illustrate the widening of U.S. wage inequality for both men and women over the past four decades. This figure plots the change in log real weekly wages by percentile for men and for women from 1963 to 2005.[9] The figure displays a sizable expansion of wage inequality with the 90th percentile earners rising by approximately 45 log points (more than 55%) relative to 10th percentile earners for both men and women. The figure also indicates a monotone (and almost linear) spreading out of the entire wage distribution for women and for the wage distribution above around the 30th percentile for men. Notably, women have substantially gained on men throughout the wage distribution over the last four decades.

We focus on four inequality concepts: changes in overall wage inequality, summarized by the 90/10 log wage differential; changes in inequality in the upper and lower halves of the wage distribution, summarized by 90/50 and 50/10 log wage gaps (which we refer to as upper-tail and lower-

[7] We also drop from the sample (full-time) workers with weekly earnings below one-half the value of the real minimum wage in 1982 ($67 a week in 1982 dollars or $112 a week in 2000 dollars). Starting in 1976 (earnings year 1975), the March survey began collecting information on hours worked in the prior year, and this allows us to create a second March sample of hourly wage data for all wage and salary workers for earnings years 1975 to 2005. Supplemental tables using the hourly wage sample for the March CPS are available online at http://www.mitpressjournals.org/doi/suppl/10.1162/rest.90.2.300 (see online reference tables 1a, 1b, and 2).

[8] Autor, Katz, and Kearney (2005) and Lemieux (2006b) find large discrepancies in trends in residual inequality in the May/ORG versus March samples beginning in 1994. These series closely parallel each other from 1979 to 1994, and then diverge sharply, with the March data showing a continued rise in residual inequality from 1994 to 2005 for hourly workers while the ORG data show a flattening. Lemieux attributes the bulk of this divergence to a differential rise in measurement error in the March sample. Autor, Katz, and Kearney call attention to another source of the discrepancy: the redesign of the CPS ORG survey in 1994. This redesign changed the format (and increased the complexity) of the earnings component of the survey, and was followed by a striking increase in earnings nonresponse: from 15.3% in 1993 (immediately prior to the redesign) to 23.3% in the last quarter of 1995 (the first quarter in which allocation flags are available in the redesigned survey), reaching 31% by 2001 (Hirsch & Schumacher, 2004). The contemporaneous rise in the earnings imputation rate in the March survey was comparatively small.

[9] The top-coding of CPS wage data makes it not very useful for measuring changes in the very top part of the distribution. Thus, we symmetrically trim the top and bottom parts of the distribution in figure 1 and focus on wage changes from the 3rd to 97th percentile.

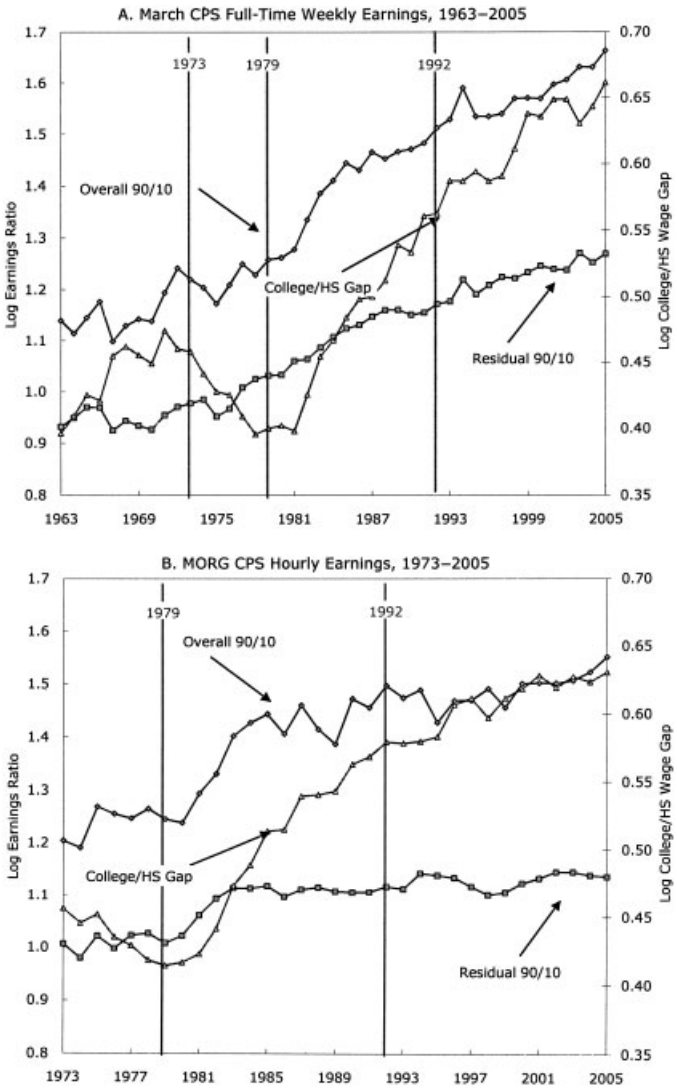
FIGURE 1.—CHANGE IN LOG REAL WEEKLY WAGE BY PERCENTILE, FULL-TIME WORKERS, 1963–2005

Source: March CPS data for earnings years 1963–2005, full-time, full-year workers ages 16 to 64 with 0 to 39 years of potential experience whose class of work in their longest job was private or government wage/salary employment. Full-time, full-year workers are those who usually worked 35-plus hours per week and worked forty plus weeks in the previous year. Weekly earnings are calculated as the logarithm of annual earnings divided by weeks worked. Calculations are weighted by CPS sampling weights and are deflated using the personal consumption expenditure (PCE) deflator. Earnings of below $67/week in 1982 dollars ($112/week in 2000 dollars) are dropped. Allocated earnings observations are excluded in earnings years 1967 forward using either family earnings allocation flags (1967–1974) or individual earnings allocation flags (1975 earnings year forward).

tail inequality); between-group wage differentials, illustrated using the college/high school wage premium; and within-group (residual) wage inequality, summarized by the 90/10, 90/50, and 50/10 residual wage gaps conditioning on measures of education, age/experience, and gender.[10]

Figures 2A and 2B display the evolution of the 90/10 overall and residual wage gaps for males and the college/high school log wage premium for our two core samples; March FTFY 1963 to 2005 and CPS May/ORG hourly 1973 to 2005. The estimated college/high school log wage premium represents a fixed weighted average of the college-plus/high school wage gaps separately estimated for males and for females in four different experience groups. The figure underscores a key, and oft-neglected, fact about the evolution of U.S. wage inequality, which is that the rise of inequality is not a unitary phenomenon. While all three inequality measures expand in tandem during the 1980s then flatten somewhat in the 1990s, the series diverged in both the 1970s and the 1960s. Specifically, while overall and residual inequality were either modestly rising (March) or flat (May/ORG) during the 1970s, the college wage premium declined sharply in this decade and then rebounded even more rapidly during the 1980s. The college wage premium expanded considerably during the 1960s, even while aggregate inequality was quiescent. These divergent patterns suggest that the growth of inequality is unlikely to be adequately explained by any single factor. Underlying the rapid growth of overall wage inequality during the 1980s followed by a deceleration in the 1990s is a divergence in inequality trends at the top and bottom of the wage distribution. This divergence is shown in figure 3,

FIGURE 2.—THREE MEASURES OF WAGE INEQUALITY: COLLEGE/HIGH SCHOOL PREMIUM, MALE 90/10 OVERALL INEQUALITY, AND MALE 90/10 RESIDUAL INEQUALITY



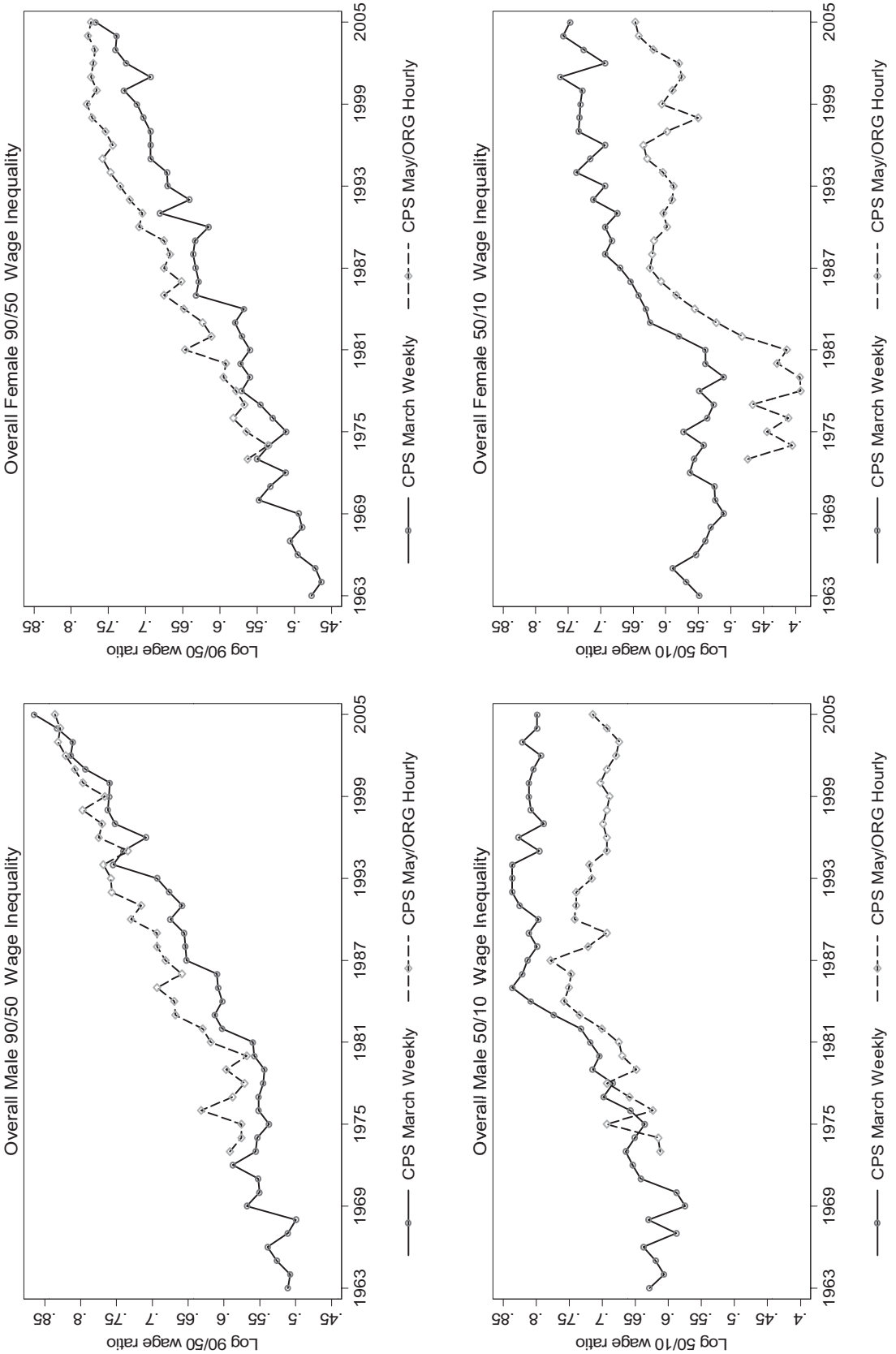
Sample for panel A is full-time, full-year workers from March CPS for earnings years 1963–2005. Sample for panel B is CPS May/ORG, all hourly workers for earnings years 1973–2005. Processing of March CPS data A is detailed in table 1 and figure 1 notes. For panel B, samples are drawn from May CPS for 1973 to 1978 and CPS Merged Outgoing Rotation Group for years 1979 to 2005. Sample is limited to wage/salary workers ages 16 to 64 with 0 to 39 years of potential experience in current employment. Calculations are weighted by CPS sample weight times hours worked in the prior week. Hourly wages are equal to the logarithm of reported hourly earnings for those paid by the hour and the logarithm of usual weekly earnings divided by hours worked last week for nonhourly workers. Top-coded earnings observations are multiplied by 1.5. Hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000 dollars) are dropped, as are hourly wages exceeding 1/35th the top-coded value of weekly earnings. All earnings are deflated by the chain-weighted (implicit) price deflator for personal consumption expenditures (PCE). Allocated earnings observations are excluded in all years, except where allocation flags are unavailable (January 1994 to August 1995). Where possible, we identify and drop nonflagged allocated observations by using the unedited earnings values provided in the source data.

The college/high school wage premium series depicts a fix-weighted ratio of college to high/school wages for a composition-constant set of sex-education-experience groups (two sexes, five education categories, and four potential experience categories). See table 1 notes and data appendix for further details.

The overall 90/10 inequality series depicts the difference between the 90th and 10th percentile of log weekly (March) or log hourly (May/ORG) male earnings. The residual 90/10 series depicts the 90/10 difference in wage residuals from a regression of the log wage measure on a full set of age dummies, dummies for nine discrete/schooling categories, and a full set of interactions among the schooling dummies and a quartic in age.

[10] The robustness of conclusions concerning the timing of changes in overall and residual wage inequality changes to the choice of wage concept and sample are illustrated in an online reference, tables 1a and 1b, which presents changes over consistent subperiods from 1975 to 2005 of different measures of inequality for males, females, and both combined using weekly earnings for full-time workers and hourly wages for all workers for the March CPS and May/ORG CPS.

FIGURE 3.—90/50 AND 50/10 WEEKLY WAGE INEQUALITY IN MARCH (FULL-TIME WORKERS) AND HOURLY WAGE INEQUALITY IN MAY/ORG (ALL WORKERS) CPS SERIES, 1963–2005



See notes to figure 2 for details on samples and data processing.

TABLE 1.—CHANGES IN REAL, COMPOSITION-ADJUSTED LOG WEEKLY WAGES FOR FULL-TIME, FULL-YEAR WORKERS, 1963–2005. (100 × CHANGE IN MEAN LOG REAL WEEKLY WAGES)

	1963–1971	1971–1979	1979–1987	1987–1995	1995–2005	1963–2005
All	19.5	0.6	−0.8	−4.8	7.6	22.2
Sex						
Men	21.1	0.1	−4.9	−7.8	6.7	15.3
Women	17.3	1.4	4.9	−0.7	9.0	31.8
Education (years of schooling)						
0–11	17.0	1.8	−8.4	−10.3	2.5	2.6
12	17.6	3.2	−3.2	−6.6	5.8	16.8
13–15	18.6	0.6	1.2	−5.3	9.5	24.6
16+	25.4	−4.2	6.8	2.8	12.5	43.3
16–17	22.9	−4.9	5.6	1.0	11.9	36.5
18+	31.3	−2.6	9.5	6.8	14.0	59.0
Experience (males)						
5 years	20.0	−3.6	−8.5	−7.6	9.0	9.3
25–35 years	21.6	3.4	−1.6	−8.1	3.8	19.2
Education and experience (males)						
Education 12						
Experience 5	19.4	0.7	−16.1	−10.3	7.1	0.7
Experience 25–35	17.0	6.3	−2.5	−7.6	0.3	13.6
Education 16+						
Experience 5	23.1	−11.0	9.3	−1.9	10.0	29.5
Experience 25–35	35.0	1.7	2.6	−2.2	13.8	50.9

Tabulated numbers are changes in the (composition-adjusted) mean log wage for each group, using data on full-time, full-year workers ages 16 to 64 from the March CPS covering earnings in calendar years 1963 to 2005. The data are sorted into sex-education-experience groups of two sexes, five education categories (high school dropout, high school graduate, some college, college graduate, and postcollege), and four potential experience categories (0–9, 10–19, 20–29, and 30–39 years). Log weekly wages of full-time, full-year workers are regressed in each year separately by sex on dummy variables for four education categories, a quartic in experience, three region dummies, black and other race dummies, and interactions of the experience quartic with three broad education categories (high school graduate, some college, and college plus). The (composition-adjusted) mean log wage for each of the forty groups in a given year is the predicted log wage from these regressions evaluated for whites, living in the mean geographic region, at the relevant experience level (5, 15, 25, or 35 years depending on the experience group). Mean log wages for broader groups in each year represent weighted averages of the relevant (composition-adjusted) cell means using a fixed set of weights, equal to the mean share of total hours worked by each group over 1963–2005. All earnings numbers are deflated by the chain-weighted (implicit) price deflator for personal consumption expenditures. Earnings of less than $67/week in 1982 dollars ($112/week in 2000 dollars) are dropped. Allocated earnings observations are excluded in earnings years 1967 forward using either family earnings allocation flags (1967–1974) or individual earnings allocation flags (1975 earnings year forward).

which compares the evolution of the 90/50 and 50/10 log hourly and full-time weekly wage gaps for males and females. Upper-tail and lower-tail wage inequality expanded rapidly in the first half of the 1980s for both men and women. But the 50/10 wage gap for the most part stopped growing after 1987—and the male hourly wage series from the CPS May/ORG shows an actual decline in the 50/10 since the late 1980s. By contrast, the 90/50 wage gap continues to grow smoothly from 1979 to 2005. Thus, the deceleration of overall inequality growth since 1987 actually reflects an abrupt halt or reversal in lower-tail inequality expansion paired with a secular rise in upper-tail inequality.[11]

The divergent growth of upper- and lower-tail wage inequality in the 1980s and 1990s is corroborated by microdata on wages and total compensation from the establishment-based Employment Cost Index (Pierce, 2001). And the steady growth of upper-tier earnings inequality is seen in rising shares of wages paid to the top 10% and top 1% of U.S. earners since the late 1970s in tax data (Piketty & Saez, 2003).

To summarize, the sharp growth in wage dispersion in the lower half of the wage distribution during the early to mid-1980s seems to have been an episodic event that has not reoccurred over the past fifteen years. By contrast, the steady growth of wage dispersion in the upper half of the wage distribution appears to represent a secular trend that has been ongoing for 25 years.

B. *Trends in Wage Levels and Between-Group Inequality*

Table 1 summarizes between-group wage structure changes by subperiod from 1963 to 2005 for groups defined by sex, education, and potential experience. Mean (predicted) log real weekly wages were computed in each year for forty sex-education-experience groups, and mean wages for broader groups are fixed-weighted averages of the relevant subgroup means, using the average share of total hours worked for each group over 1963 to 2005 as weights to adjust for compositional changes.[12]

The first row indicates that composition-adjusted mean real wages increased by 22.2 log points over the full period. Wage growth was rapid in the 1960s, stagnant or declining from 1971 to 1995, and rapid from 1995 to 2005. The next two rows show that women gained substantially on males—by 16.5 log points over the full sample—with women's relative earnings gains concentrated in the 1979 to 1995 period.

[11] The divergent growth of the 90/50 and 50/10 wage differentials has been previously emphasized by Murphy and Welch (2001) and Mishel, Bernstein, and Boushey (2002) and is noted by Lemieux (2006b). Using decennial Census earnings data, Angrist, Chernozhukov, and Fernándedez-Val (2006) document a sharp rise in residual inequality from 1980 to 1990, with a continuing increase from 1990 to 2000 concentrated in the upper half of the wage distribution.

[12] The March and May/ORG samples appear equally valid for measuring between-group wage trends and show almost identical patterns in between-group wage differentials since 1973. The March data cover an additional decade.

The following six rows highlight the expansion of educational wage differentials, with particularly large increases in the relative earnings of college graduates. The sharp differences across decades seen in figure 2 are evident in these detailed figures, with educational wage differentials rising in the 1960s, narrowing in the 1970s, increasing sharply in the 1980s, and growing at a slightly less torrid pace since 1995. The bottom part of the table contrasts changes in real wages for younger and older males. Experience differentials expanded for college and high school graduates, with the rise for college graduates concentrated in the 1960s and 1970s and the rise for high school graduates concentrated in the 1980s.

The expansion of between-group wage differentials has been less continuous—and undergone more reversals—than the trend toward increasing overall wage inequality over the last four decades.

III. The Sources of Rising Inequality: Proximate Causes

We now present an analysis of the leading proximate causes of overall and between-group wage inequality, focusing on supply and demand factors, unemployment, and the minimum wage. We start with simple time series models of the U.S. college wage premium covering 1963 to 2005 and augment the specification to allow for an impact of a key labor market institutional factor, the federal minimum wage.[13]

A. *Sources of the Rising College/High School Wage Premium*

Our illustrative conceptual framework starts with a CES production function for aggregate output Q with two factors, college equivalents (c) and high school equivalents (h):

$$Q_t = [\alpha_t(a_t N_{ct})^\rho + (1 - \alpha_t)(b_t N_{ht})^\rho]^{1/\rho}, \tag{1}$$

where N_{ct} and N_{ht} are the quantities employed of college equivalents (skilled labor) and high school equivalents (unskilled labor) in period t, a_t and b_t represent skilled and unskilled labor augmenting technological change, α, is a time-varying technology parameter that can be interpreted as indexing the share of work activities allocated to skilled labor, and ρ is a time-invariant production parameter. Skill-neutral technological improvements raise a_t and b_t by the same proportion. Skill-biased technical changes involve increases in a_t/b_t or α_t. The aggregate elasticity of substitution between college and high school equivalents is given by $\sigma = 1/(1 - \rho)$.

Under the assumption that college and high school equivalents are paid their marginal products, we can use equation (1) to solve for the ratio of marginal products of the two labor types yielding a relationship between relative wages in year t, w_{ct}/w_{ht}, and relative supplies in year t, N_{ct}/N_{ht} given by

$$\ln(w_{ct}/w_{ht}) = \ln[\alpha_t/(1 - \alpha_t)] + \rho \ln(a_t/b_t)] - (1/\sigma)\ln(N_{ct}/N_{ht}), \tag{2}$$

which can be rewritten as

$$\ln(w_{ct}/w_{ht}) = (1/\sigma)[D_t - \ln(N_{ct}/N_{ht})], \tag{3}$$

where D_t indexes relative demand shifts favoring college equivalents and is measured in log quantity units. The impact of changes in relative skill supplies on relative wages depends inversely on the magnitude of aggregate elasticity of substitution between the two skill groups. The greater is σ, the smaller the impact of shifts in relative supplies on relative wages and the greater must be fluctuations in demand shifts (D_t) to explain any given time series of relative wages for a given time series of relative quantities. Changes in D_t can arise from (disembodied) SBTC, nonneutral changes in the relative prices or quantities of nonlabor inputs, and shifts in product demand.

Following the approach of Katz and Murphy (1992), we directly estimate a version of equation (3) to explain the evolution from 1963 to 2005 of the overall log college/high school wage differential series for FTFY workers from the March CPS shown in panel A of figure 2. We substitute for the unobserved demand shifts D, with simple time trends and a measure of labor market cyclical conditions, the unemployment rate of males aged 25–54 years. We also include an index of the log relative supply of college/high school equivalents.[14] Our full model includes the log real minimum wage as a control variable:

$$\ln(w_{ct}/w_{ht}) = \gamma_0 + \gamma_1 t + \gamma_2 \ln(N_{ct}/N_{ht}) + \gamma_3(\text{RealMinWage}_t) + \gamma_4\text{Unemp}_t + \varepsilon_t, \tag{4}$$

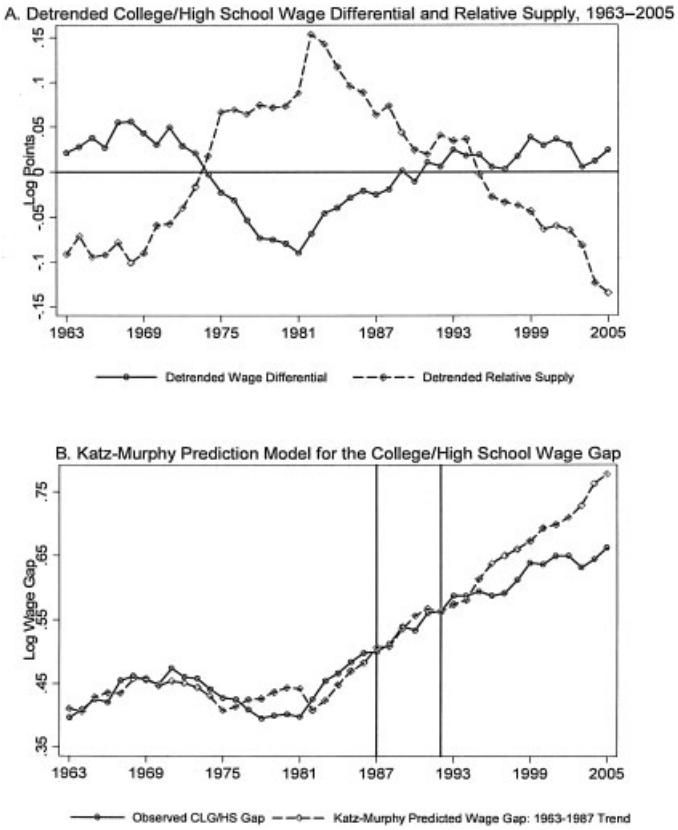
where γ_2 provides an estimate of $1/\sigma$.

The large increase in the college wage premium over the last forty years coincided with a substantial secular rise in the relative supply of college workers. The college graduate share of the full-time equivalent workforce increased from about 10.6% in 1960 to over 30% in 2005. Given this rapid growth in college graduate supply, a market-clearing model requires (even more) rapid growth in relative demand for college workers to reconcile increasing college supply with a rising college wage premium.

[13] The present analysis of the college wage premium extends earlier work in Katz and Murphy (1992) and Katz and Autor (1999), drawing on additional years of data.

[14] We use a standard measure of college/noncollege relative supply calculated in "efficiency units" to adjust for changes in labor force composition by gender and experience groups. Full details are provided in the data appendix.

FIGURE 4.—COLLEGE/HIGH SCHOOL RELATIVE SUPPLY AND WAGE DIFFERENTIAL, 1963–2005 (MARCH CPS)



Composition-adjusted college/high school relative wages are calculated using March FTFY earners data, sorted into sex-education-experience groups of two sexes, five education categories, and four potential experience categories. Mean log wages for broader groups in each year represent weighted averages of the relevant (composition-adjusted) cell means using a fixed set of weights that are equal to the mean share of total hours worked by each group over 1963 to 2005 from the March CPS. See table 1 notes for additional details.

The college/high school log relative supply index is the natural logarithm of the ratio of college-equivalent to noncollege-equivalent labor supply in efficiency units in each year. See the data appendix for details.

The detrended supply and wage series in panel A are the residuals from seperate OLS regressions of the relative supply and relative wage measures on a constant and a linear time trend. The Katz-Murphy predicted wage gap series in panel B is the fitted values from an OLS regression of the college/high school wage gap for years 1963 through 1987 on a constant and the college/high school relative supply measure. Plotted 1988 to 2005 values are out-of-sample predictions.

The upper panel of figure 4 plots the college relative supply and wage premium series over 1963 to 2005 deviated from a linear time trend. This figure reveals an acceleration of the growth in the relative supply of college workers in the 1970s relative to the 1960s, followed by a dramatic slowdown starting in 1982. These fluctuations in the growth rate of relative supply, paired with a constant trend growth in relative college demand, do an effective job of explaining the evolution of the college wage premium from 1963 to 2005. The figure illustrates that deviations in relative supply growth from a linear trend roughly fit the broad changes in the detrended college wage premium.

Table 2 presents representative regression models for the overall college/high school log wage gap following this approach. The first column uses the specification of Katz and Murphy (1992) for the 1963 to 1987 period (the period analyzed by Katz-Murphy) with only a linear time trend and the relative supply measure included as explanatory variables. Although our data processing methods differ somewhat from those of Katz and Murphy, we uncover quite similar results with an estimate of $\gamma_2 = 0.64$ (implying $\sigma = 1.57$) and with estimated trend growth in the college wage premium of 2.6% per annum. The lower panel of figure 4 uses this replication of the basic Katz-Murphy model from column 1 of table 2 to predict the evolution of the college wage premium for the full sample period of 1963 to 2005 and compares the predicted and actual college wage gap measures.

The Katz-Murphy model does an excellent job of forecasting the growth of the college wage premium through 1992 (with the exception of the late 1970s), but the continued slow growth of relative supply after 1992 leads it to overpredict the growth in the college wage premium over the last decade. This pattern implies there has been a slowdown in relative demand growth for college workers since 1992, as illustrated by a comparison of the models in

TABLE 2.—REGRESSION MODELS FOR THE COLLEGE/HIGH SCHOOL LOG WAGE GAP, 1963–2005

	(1) 1963–1987	(2)	(3)	(4)	(5) 1963–2005	(6)	(7)	(8)
CLG/HS relative supply	−0.636	−0.411	−0.619	−0.599	−0.609	−0.728	−0.403	
	(0.130)	(0.046)	(0.066)	(0.112)	(0.102)	(0.155)	(0.067)	
Log real minimum wage						−0.049	−0.117	−0.144
						(0.051)	(0.047)	(0.065)
Male prime-age unemp. rate						0.004	−0.001	−0.018
						(0.004)	(0.004)	(0.003)
Time	0.026	0.018	0.026	0.028	0.021	0.028	0.017	0.006
	(0.005)	(0.001)	(0.002)	(0.006)	(0.006)	(0.007)	(0.002)	(0.001)
Time^2/100				−0.011	0.030	0.017		
				(0.006)	(0.015)	(0.017)		
Time^3/1000					−0.006	−0.005		
					(0.002)	(0.002)		
Time × post-1992			−0.008					
			(0.002)					
Constant	−0.159	0.043	−0.146	−0.143	−0.124	−0.160	0.266	0.689
	(0.119)	(0.037)	(0.057)	(0.108)	(0.098)	(0.191)	(0.112)	(0.120)
Observations	25	43	43	43	43	43	43	43
R-squared	0.563	0.934	0.953	0.940	0.952	0.955	0.944	0.891

Standard errors in parentheses. Each column presents an OLS regression of the fixed-weighted college/high school wage differential on the indicated variables. The U.S. federal minimum wage is deflated by the personal consumption expenditure deflator. Source for labor supply and earnings measures is the March CPS, earnings years 1963–2005.

columns 2 and 3 of table 2 without and with allowing for a trend break in 1992.[15] The model in column 3 covering the full 1963–2005 period indicates a significant slowdown of demand growth after 1992 but still indicates a large impact of relative supply growth with an estimated aggregate elasticity of substitution of 1.62 (1/0.619).[16] Subsequent models in columns 4 through 6 that allow for a more flexible time trend—either a quadratic or cubic function—imply that trend demand growth for college relative to noncollege workers slowed in the early 1990s.

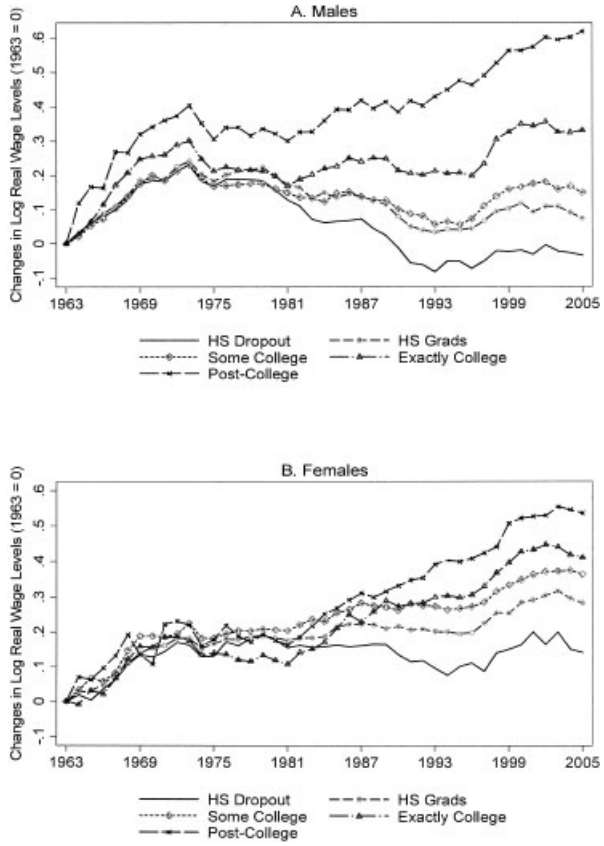
The implied slowdown in trend demand growth in the 1990s is potentially inconsistent with a naive SBTC story looking at the growth of computer investments since these continued rapidly in the 1990s. One potential explanation for this implied slowdown is the strong cyclical labor market of the expansion of the 1990s, leading to a low unemployment rate. The impacts of labor market institutions such as the erosion of the real value of the minimum wage since the early 1980s might also play a role.

The roles of cyclical conditions and the minimum wage are examined in the augmented models illustrated in columns 6 and 7 of table 2. The real minimum wage and prime-age male unemployment rates have modest additional explanatory power in the expected directions and reduce the extent of slowdown in trend demand growth over the last decade. But the inclusion of these variables does not much alter the central role for relative supply growth fluctuations and trend demand growth in explaining the evolution of the college wage premium. A model without the relative supply variable in column 7 leads to larger impacts of the real minimum wage but it also has less explanatory power and generates a puzzling negative impact of prime-age male unemployment. These cyclical and institutional factors are insufficient to resolve the puzzle posed by slowing trend relative demand for college workers in the 1990s.

A closer look at the data suggests why the simple CES model with two factors—college and high school equivalents—fails to provide an adequate explanation of the evolution of between-group wage inequality starting in the early 1990s. As shown in figure 5, the real, composition-adjusted earnings of full-time, full-year workers at different levels of educational attainment "polarized" after 1987 in a manner consistent with the divergent trends in 90/50 and 50/10 inequality documented in figure 3. In particular, the wage gap between males with a postcollege education and those with a high school education rose rapidly and monotonically from 1979 through 2005, increasing by 43.1 log points overall and 15.4, 15.7, and 12.0 points respectively between 1979–1988, 1988–1997, and 1997–2005.[17] By contrast, after increasing by 13.3 log points between 1979 and 1987, the wage gap between males with exactly a college degree and those with a high school education rose comparatively slowly thereafter, by 4.5 and 9.0 log points respectively between 1988–1997 and 1997–2005. By implication, between 1988 and 2005, the earnings of postcollege males rose by 14.2 log points more than the earnings of college-only males. Conversely, at the bottom of the wage distribution, the wage gap between high school graduates and high school dropouts increased steadily from 1979 through 1997, then flattened or reversed.

This pattern, in which wage gaps *within* college-educated and non-college-educated worker groups diverge, is inconsistent with the basic, two-factor CES model. In this model, the labor input of all college-educated worker subgroups is assumed to be perfectly substitutable up to a scalar multiple, and similarly for noncollege worker subgroups. Accordingly, the wage *ratio* of college-educated to postcollege-

FIGURE 5.—TRENDS IN COMPOSITION-ADJUSTED REAL LOG WEEKLY FULL-TIME WAGES BY GENDER AND EDUCATION, 1963–2005 (MARCH CPS)



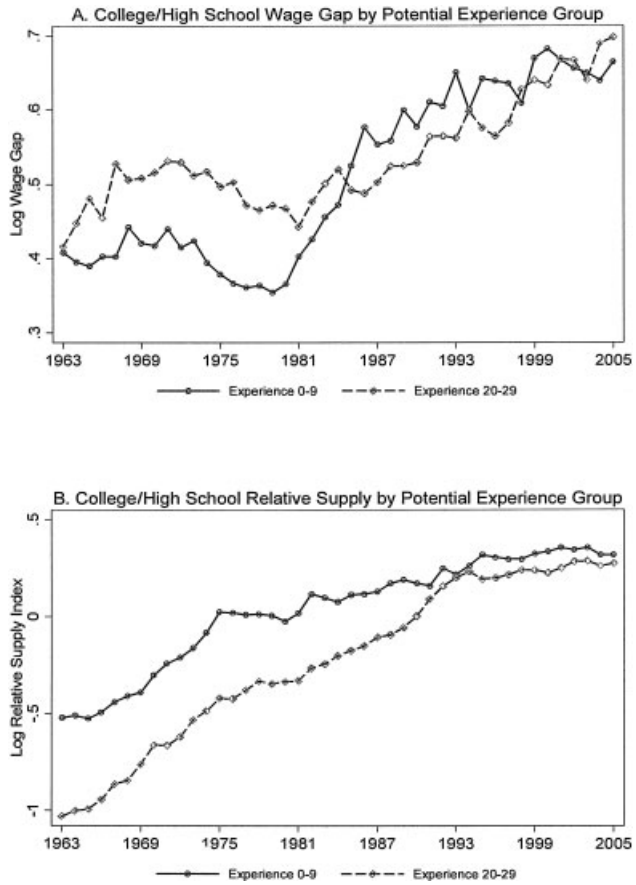
See notes to table 1 for details on samples and data processing.

[15] In fact, Goldin and Katz (2007) find that rather steady relative demand growth for college workers combined with relative supply growth fluctuations does a nice job of explaining the longer-run evolution of the U.S. college wage premium from 1915 to 2005. Their analysis implies a modest acceleration in demand growth after 1950 as well as a slowdown in the trend demand after 1990. The slowing of relative demand growth for college workers after 1990 also has been noted by Autor, Katz, and Krueger (1998), Katz and Autor (1999), and Card and DiNardo (2002).

[16] Similar conclusions of a significant slowdown in trend relative-demand growth for college workers arise in models allowing trend breaks in any year from 1989 to 1994.

[17] For females, earnings growth between 1988 and 2005 among postcollege-educated workers was substantially greater than for college-only workers, but the pattern was reversed for 1979 to 1988.

FIGURE 6.—COMPOSITION-ADJUSTED LOG RELATIVE COLLEGE/HIGH SCHOOL WAGE AND SUPPLY BY POTENTIAL EXPERIENCE AND AGE GROUPS, 1963–2003 (MARCH CPS)



See notes to figure 4 for details on construction of supply and wage series.

educated workers should be roughly constant, as should the wage ratio of high school dropouts to high school graduates. This two-factor assumption fits the data rather well from 1963 to 1987. However, the drastic rise in earnings of postsecondary relative to college-only workers after 1987 and the slightly increasing earnings of dropouts relative to high school graduates after 1997 represent significant departures from the assumptions of the model. Fundamentally, the two-factor model does not accommodate a setting in which the wages of very high- and very low-skilled workers rise relative to those of middle-educated workers—that is, a setting where wage growth *polarizes.* We consider the sources of this polarization in section V.

B. The College/High School Gap by Experience Group

As shown in table 1, changes in the college/high school wage gap differed substantially by age/experience groups over recent decades, with the rise in the college/high school gap concentrated among less experienced workers in the 1980s. We illustrate this pattern in figure 6 through a comparison of the evolution of the college premium (panel A) and college relative supply (panel B) for younger workers (those with 0–9 years of potential experience) and older workers (those with 20–29 years of potential experience). The return to college for younger workers has increased much more substantially since 1980 than for older workers. To the extent that workers with similar education but different ages or experience levels are imperfect substitutes in production, one should expect age-group or cohort-specific relative skill supplies—as well as aggregate relative skill supplies—to affect the evolution of the college/high school wage premium by age or experience as emphasized in a careful analysis by Card and Lemieux (2001). Consistent with this view, the lower panel of figure 6 shows a much more rapid deceleration in relative college supply among younger than older workers in the mid- to late 1970s.

In table 3, we take fuller account of these differing trends by estimating regression models for the college wage by experience group that extend the basic specification in equation (4) to include own experience group relative skill supplies. The first two columns of table 3 present regressions pooled across four potential experience groups (those with 0–9, 10–19, 20–29, and 30–39 years of experience) allowing for group-specific intercepts but constraining the other coefficients to be the same for all experience groups. These models estimate:

$$\begin{aligned}\ln(w_{cjt}/w_{hjt}) = \beta_0 + \beta_1[\ln(N_{cjt}/N_{hjt}) \\ - \ln(N_{ct}/N_{ht})] + \beta_2 \ln(N_{ct}/N_{ht}) \\ + X_t\beta_3 + \delta_j + \eta_{jt},\end{aligned} \tag{5}$$

where j indexes experience group, the δ_j are experience group main effects, and X_t, includes measures of time trends and other demand shifters. This specification arises from an aggregate CES production function in college and high school equivalents of the form of equation (1) where these aggregate inputs are themselves CES subaggregates of college and high school labor by experience group (Card & Lemieux, 2001). Under these assumptions, $-1/\beta_2$ provides an estimate of σ, the aggregate elasticity of substitution, and $-1/\beta_1$ provides an estimate of σ_E, the partial elasticity of substitution between different experience groups within the same education group.

The estimates in the first two columns of table 3 indicate substantial effects of both own-group and aggregate supplies on the evolution of the college wage premium by experience group. While the implied estimates of the aggregate elasticity of substitution in the table 3 models are very similar to the aggregate models in table 2, the implied value of the partial elasticity of substitution between experience groups is around 3.55 (somewhat lower than the estimates in Card & Lemieux, 2001). These estimates indicate that differences in own-group relative college supply growth go a substantial distance toward explaining variation across experience groups in the evolution of the college wage premium in recent decades. For example, as seen in figure 6, from 1980 to 2005 the college wage premium increased by

TABLE 3.—REGRESSION MODELS FOR THE COLLEGE/HIGH SCHOOL LOG WAGE GAP BY POTENTIAL EXPERIENCE GROUP, 1963–2005, MALES AND FEMALES POOLED

	Potential Experience Groups					
	All Experience Groups		0–9 yrs	10–19 yrs	20–29 yrs	30–39 yrs
Own supply minus aggregate supply	−0.282	−0.281	−0.169	−0.325	0.101	0.002
	(0.027)	(0.027)	(0.130)	(0.084)	(0.084)	(0.119)
Aggregate supply	−0.600	−0.705	−0.855	−0.474	−0.398	−0.544
	(0.087)	(0.131)	(0.262)	(0.182)	(0.224)	(0.239)
Log real minimum wage		−0.074	−0.340	−0.145	0.098	0.028
		(0.037)	(0.076)	(0.049)	(0.054)	(0.067)
Prime-age male unemployment		0.004	0.005	0.002	0.003	0.000
		(0.003)	(0.007)	(0.004)	(0.005)	(0.006)
Time	0.027	0.031	0.040	0.015	0.016	0.027
	(0.004)	(0.006)	(0.012)	(0.009)	(0.011)	(0.011)
$\text{Time}^2/100$	−0.009	−0.012	−0.025	0.010	0.000	−0.021
	(0.005)	(0.006)	(0.012)	(0.009)	(0.010)	(0.012)
Constant	−0.046	−0.013	0.268	0.359	−0.032	−0.065
	(0.079)	(0.151)	(0.300)	(0.223)	(0.256)	(0.275)
N	172	172	43	43	43	43
R-squared	0.863	0.868	0.926	0.969	0.898	0.663

Standard errors in parentheses. Each column presents an OLS regression of the fixed-weighted weighted college/high school wage differential on the indicated variables. The college/high school wage premium is calculated at the midpoint of each potential experience group. Real minimum wage is deflated by the personal consumption expenditure deflator. Columns 1 and 2 also include dummy variables for the four potential experience groups used in the table.

29.9 log points for the 0–9-year experience group and by 23.0 log points for the 20–29-year experience group. Over the same period the own-group relative college supply for the 0–9-year experience group grew by 26.7 log points less rapidly than for the 20–29-year experience group. Thus, using the implied own-group relative inverse substitution elasticity of −0.282 in column 1 of table 3, we find that the slower relative supply growth for the younger (0–9-year) experience group explains the entirety (7.53 log points of a 6.90 log point gap) of the larger increase in the college premium for the younger than for the older (20–29-year) experience group.

The final four columns of table 3 present regression models of the college wage premium separately estimated by experience group. Trend demand changes and relative skill supplies play a large role in changes in educational differentials for younger and prime-age workers. The college wage premium for younger workers appears more sensitive to own-group and aggregate relative skill supplies than the premium for older workers. The real minimum wage is a significant determinant of changes in the college wage premium for younger workers, but, plausibly, does not appear important for older workers.

In summary, a simple supply-demand framework emphasizing a secular increase in the relative demand for college workers combined with fluctuations in relative skill supplies can account for *some* of the key patterns in the recent evolution of between-group inequality, including the contraction and expansion of the college/high school gap during the 1970s and 1980s and the differential rise in the college/high school gap by experience group in the 1980s and 1990s.[18]

[18] However, the divergence of postcollege and college-only wages is inconsistent with this simple two skill group CES framework and demands its own explanation, to which we return below.

What drives these secular demand shifts? A large literature reviewed in Katz and Autor (1999) and Katz (2000) yields two consistent findings suggesting that SBTC has played an integral role.[19] The first is that the relative employment of more-educated workers and nonproduction workers has increased rapidly within detailed industries and within establishments in the United States during the 1980s and 1990s, despite the sharp rise in the relative wages of these groups (Dunne, Haltiwanger, & Troske, 1997; Autor, Katz, & Krueger, 1998). Similar patterns of within-industry increases in the proportion of skilled workers are apparent in other advanced nations (Berman, Bound, & Machin, 1998; Machin & Van Reenen, 1998). These findings suggest strong within-industry demand shifts favoring the more skilled.[20] Second, a wealth of quantitative and case-study evidence documents a striking correlation between the adoption of computer-based technologies (and associated organizational innovations) and the increased use of college-educated labor within detailed industries, within firms, and across plants within industries (Doms, Dunne, & Troske, 1997; Autor, Levy, & Murnane, 2002; Levy & Murnane, 2004; Bartel, Ichniowski, & Shaw, 2007).

C. *The Role of the Minimum Wage*

Several studies, including Lee (1999), Card and DiNardo (2002), and Lemieux (2006b), conclude the minimum wage plays a primary role in the rise of wage inequality since

[19] Skill-biased technological change refers to any introduction of a new technology, change in production methods, or change in the organization of work that increases the demand for more-skilled labor relative to less-skilled labor at fixed relative wages.

[20] Foreign outsourcing of less-skilled jobs is another possible explanation for this pattern (Feenstra & Hanson, 1999). But large within-industry shifts toward more-skilled workers are pervasive even in sectors with little or no observed foreign outsourcing activity. Foreign outsourcing appears likely to become increasingly important, however.

1980. Yet, our simple models above do not find the minimum wage to be important in the evolution of educational wage differentials, except possibly for young workers. Why do our conclusions differ? The discrepancy partially arises from a disjuncture between trends in between-group inequality (the college/high school gap) and trends in overall and residual inequality. As seen in figure 2, overall inequality was flat during the 1970s while between-group inequality fell; conversely, as between-group inequality continued to rise in the 1990s, residual inequality stabilized. Between-group and residual inequality move closely together only during 1979–1987.

Following our simple models for the college/high school earnings gap above, we provide a time series analysis for the proximate sources of the growth of overall, upper-tail, and lower-tail hourly wage inequality. As emphasized by Card and DiNardo (2002), there is a striking time series relationship between the real value of the federal minimum wage and hourly wage inequality, as measured by the 90/10 log earnings ratio. This relationship is depicted in figure 7. A simple regression of the 90/10 log hourly wage gap from the CPS May/ORG for the years 1973 to 2005 on the real minimum wage yields a coefficient of -0.74 and an R-squared of 0.71. Based in part on this tight correspondence, Card and DiNardo (2002) and Lemieux (2006b) argue that much of the rise in overall and residual inequality over the last two decades may be attributed to the minimum wage.[21] In a cross-state analysis of the minimum wage and wage inequality for the period 1979 to 1991, Lee (1999) reaches a similar conclusion.

A potential problem for this argument is that the majority of the rise in earnings inequality over the last two decades occurred in the upper half of the earnings distribution. Since it is not plausible that a declining minimum wage could cause large increases in upper-tail earnings inequality, this observation suggests that the minimum wage is unlikely to provide a satisfying explanation for the bulk of inequality growth. Not surprisingly, as shown in the upper panel of figure 7, the real minimum wage is highly correlated with lower-tail earnings inequality between 1973 and 2005: a log point rise in the minimum is associated with 0.26 log point compression in lower-tail inequality. Somewhat surprisingly, the minimum wage is also highly correlated with upper-tail inequality: a 1 log point rise in the minimum is associated with a 0.48 log point compression in upper-tail inequality (figure 7, lower panel).

These bivariate relationships may potentially mask other confounds. To explore these relationships in slightly greater detail, we estimated a set of descriptive regressions for 90/10, 90/50, and 50/10 hourly earnings inequality over 1973 to 2005. In addition to the minimum-wage measure used in figure 7, we augmented these models with a linear time trend, a measure of college/high school relative supply (calculated from the CPS May/ORG), the male prime-age unemployment rate (as a measure of labor market tightness), and in some specifications a post-1992 time trend, reflecting the estimated trend reduction in skill demand in the 1990s. The main finding from these models is that the strong relationship between the minimum wage and both upper- and lower-tail inequality is highly robust. In a specification that includes a linear time trend, the college/high school supply measure, and the prime-age unemployment rate variable, the minimum-wage measure has a coefficient of -0.23 for lower-tail inequality and a coefficient of -0.10 for upper-tail inequality (both significant).

These patterns suggest that the time series correlation between minimum wages and inequality is unlikely to provide causal estimates of minimum-wage impacts. Indeed, the relationship between the minimum wage and upper-tail inequality is potential evidence of spurious causation. Although the decline in the real minimum wage during the 1980s likely contributed to the expansion of lower-tail inequality—particularly for women—the robust correlation of the minimum wage with upper-tail inequality suggests other factors are at work.[22] One possibility is that federal minimum-wage changes (or inaction) during these decades were partially a response to political pressures associated with changing labor market conditions and costs to employers of a minimum-wage increase. This political economy story could help explain the coincidence of falling minimum wages and rising upper-tail inequality.[23]

IV. Rising Residual Inequality: The Role of Composition and Prices

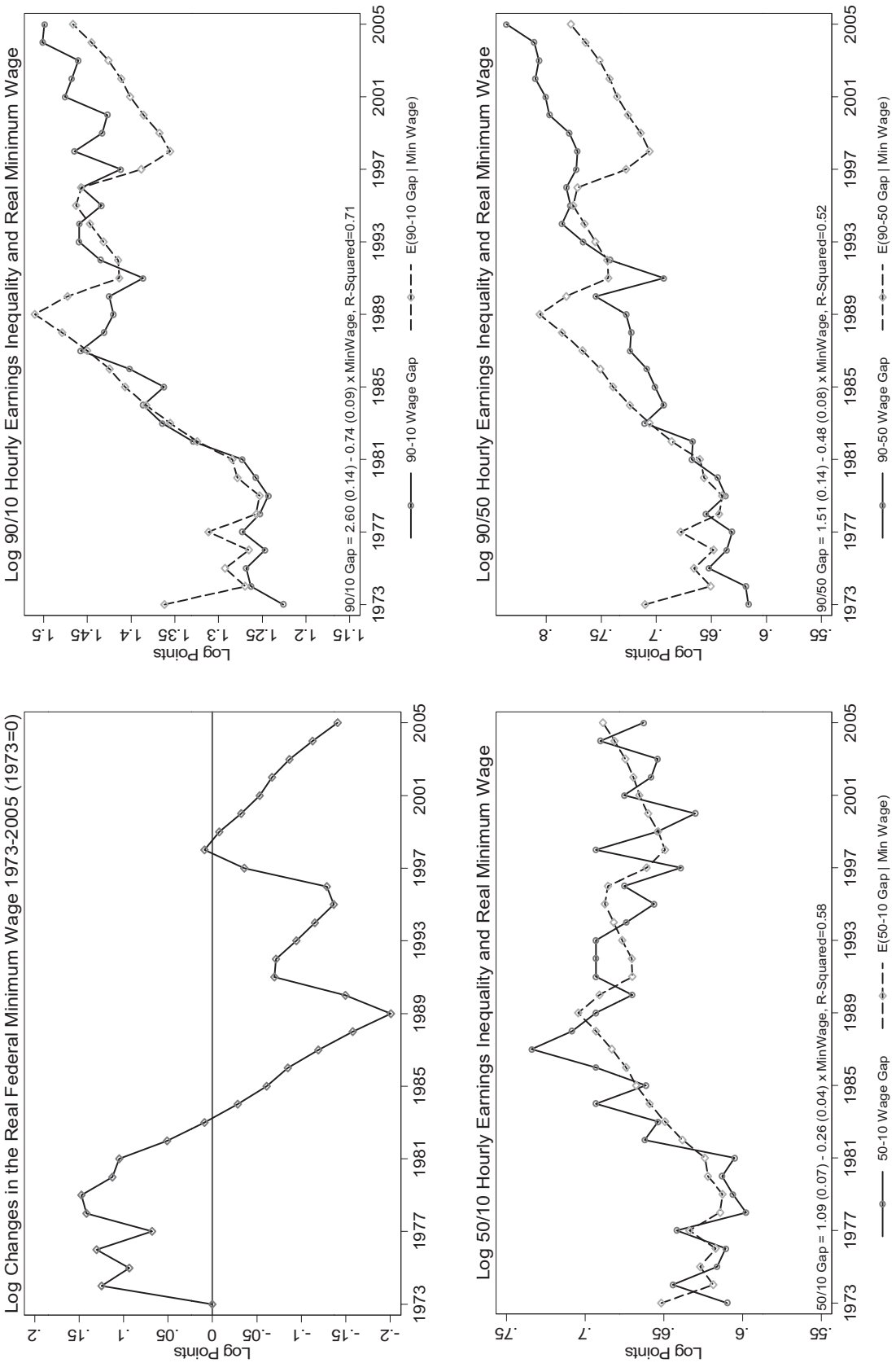
The educational attainment and labor market experience of the U.S. labor force rose substantially over the last thirty years as the large 1970s entering ("baby boom") college cohorts reached midcareer during the 1990s. The full-time

[21] Lemieux (2006b) focuses on the tight fit between the real minimum wage and residual wage variance for men and women from 1973 to 2003. We find greater time series explanatory power of the real minimum wage for residual wage inequality measures than for actual wage inequality measures. This is puzzling for the minimum-wage hypothesis since the minimum wage should "bite" more for actual low-wage workers than for residual low-wage workers.

[22] Lee (1999) also noted a puzzling relationship between the "effective" state minimum wage (the log difference between the state median and the state minimum) and upper-tail inequality. Opposite to the simple time series regressions above, Lee finds in a cross-state analysis that increases in the effective state minimum wage appear to reduce upper-tail inequality, both for males and for the pooled-gender distribution, leading him to advise caution in causally attributing trends in male and pooled-gender earnings inequality to the minimum wage.

[23] In a similar vein, Acemoglu, Aghion, and Violante (2001) argue that the decline in union penetration in the United States and the United Kingdom is partly explained by changing skill demands that reduced the viability of rent-sharing bargains between high- and low-skill workers. Furthermore, the direct effects of union decline on U.S. wage inequality growth appear to be modest. Card, Lemieux, and Riddell (2003) find that falling unionization explains about 14% of the growth of male wage variance from 1973 to 2001 (in models allowing for skill group differences in the impact of unions) with an even smaller union effect for the growth of female wage variance.

FIGURE 7.—LOG REAL FEDERAL MINIMUM WAGE AND LOG 90/10, 90/50, AND 50/10 HOURLY WAGE DIFFERENTIALS, 1973–2005 (CPS MAY/ORG).



Nominal minimum wages are deflated to real log values using the PCE deflator. In the first panel, the real log minimum-wage measure is normalized to zero in 1973. Subsequent panels depict the observed wage gap (90/10, 90/50, and 50/10) for all hourly workers from the CPS May/ORG samples in each year plotted alongside the predicted values from separate OLS regressions of the relevant wage gap on a constant and the contemporaneous real log minimum wage.

equivalent employment share of male workers with a college degree rose from less than one-fifth to approximately one-third of the U.S. male labor force from 1973 to 2005. The mean potential experience of male workers with high school or greater education also increased substantially (by two to five years) from 1973 to 2005, with the largest gains for college workers.

As discussed by Lemieux (2006b), these shifts in labor force composition may have played a role in changes in measured wage inequality. The canonical Mincer (1974) earnings model implies that earnings trajectories fan out as workers gain labor market experience. Hourly wage dispersion is also typically higher for college graduates than for less-educated workers. Thus, changes in the distribution of education or experience of the labor force can lead to changes in wage dispersion. These *compositional* effects are distinct from the standard *price* effects arising from shifts in supply-demand and institutional factors. Holding market prices constant, changes in labor force composition can *mechanically* raise or lower residual earnings dispersion simply by altering the employment share of worker groups that have more or less dispersed earnings. Similarly, changes in workforce composition can also raise or lower overall earnings dispersion by increasing or reducing heterogeneity in observed skills (Juhn, Murphy, & Pierce, 1993). These observations suggest that measured earnings dispersion may change because of the mechanical impact of composition without any underlying change in market prices.

Following such an approach, Lemieux (2006b) finds that most of the growth in residual wage dispersion from 1973 to 2003—and *all* of the growth after 1988—is explained by mechanical effects of changes in workforce composition rather than shifts in residual inequality within defined skill groups (what we call "price" effects). Lemieux concludes that the rise in residual earnings inequality is mainly attributable to institutional factors during the 1980s—especially the falling real minimum wage—and to mechanical labor force composition effects since the late 1980s.

We reassess these conclusions, adhering closely to the methods and data sources used by Lemieux (2006b). Our analysis differs from Lemieux in one key respect: whereas Lemieux focuses primarily on the contribution of prices and composition to the *variance* of wage residuals—thus aggregating over changes in the upper and lower tails of the distribution—we focus on the contribution of prices and composition to changes in upper-tail and lower-tail earnings inequality (both overall and residual). We conclude that changes in labor force composition do *not* substantially contribute to an explanation for the diverging path of upper- and lower-tail inequality, either overall or residual, over the past three decades.

A. Implementation

We employ a variant of the kernel reweighting approach introduced by DiNardo, Fortin, and Lemieux (1996, DFL hereafter).[24] We write the observed density of wages at times t and t' as

$$f(w|T=t) = \int g(w|x,T=t)h(x|T=t)dx$$
$$\text{and } f(w|T=t') = \int g(w|x,T=t')h(x|T=t')dx. \tag{6}$$

In this expression, w is the logarithm of the hourly wage, T is a variable denoting the year from which an observation is drawn, $g(w|x,T=t)$ is the density of wages for observable attributes x at survey year t, and $h(x|T=t)$ is the density of attributes x at survey year t. Equation (6) decomposes the density of wages into two functions: a "price" function, $g(\cdot)$, that provides the conditional distribution of wages for given attributes and time, and a "composition" function, $h(\cdot)$, that provides the density of attributes in that time period.

Using this decomposition, we can develop counterfactual wage densities by combining the price function $g_t(\cdot)$ from some period t with the composition function $h_t(\cdot)$ from an alternative period t'. As shown by DFL, calculating such a counterfactual simply requires reweighting the price function $g_t(\cdot)$ in year t by the ratio of the density of attributes x in year t' to the density of attributes in year t, $h_{t'}(\cdot)/h_t(\cdot)$. Applying Bayes's rule, this reweighting function can be written as

$$\frac{h(x|T=t')}{h(x|T=t)} = \frac{\Pr(T=t'|x)}{1-\Pr(T=t'|x)} \times \frac{1-\Pr(T=t')}{\Pr(T=t')}. \tag{7}$$

The reweighting function can be estimated using a logit model applied to the pooled data sources, $h(x)$, from years t and t'.

The validity of this counterfactual exercise rests on the partial equilibrium assumption that prices and quantities can be viewed as independent—that is, changes in labor market quantities, $h(x)$, do not affect labor market prices, $g(x)$. Although analytically convenient, this assumption is economically unappealing, and, moreover, is precisely opposite in spirit to our supply-demand analysis in section III. Given the dramatic changes in the education and experience of the labor market over the three decades and their attendant affects on labor market prices documented above, the partial equilibrium assumptions underlying this exercise are certain to be violated. Nevertheless, we view this analysis as worthwhile because it permits a direct assessment of the

[24] In Autor, Katz, & Kearney (2005), we provide a more complete reanalysis of Lemieux (2006b) using a quantile decomposition approach proposed by Machado and Mata (2005) and comparing the findings for the CPS March and May/ORG samples. Here we adopt the kernel reweighting approach of Lemieux to facilitate a direct comparison.

substantive conclusions of Lemieux (2006b), taking the methodology as given.[25]

To evaluate the importance of shifts in composition and prices to observed changes in overall and residual wage inequality, we draw on our core May/ORG hourly wage samples from 1973 to 2005 to construct counterfactual wage distributions. In each sample year *t,* we apply the labor force composition data, $h_t(x)$, to the price function, $g(w|x,T = t')$, from the years 1973, 1989, and 2005. This procedure simulates (via reweighting) a hypothetical set of cases where labor force composition is allowed to evolve as actually occurred over 1973 to 2005 while labor market prices are held at their start-of-period (1973), midperiod (1989), or end-of-period levels (2005). In calculating the reweighting function (equation 7), we employ the same covariates in the x vector as used by Lemieux (2006b). These include a full set of age dummies, dummies for nine discrete schooling categories, and a full set of interactions among the schooling dummies and a quartic in age. All models are estimated separately by gender.

The procedure outlined above is suitable for obtaining counterfactuals for *overall* inequality. To calculate analogous counterfactuals for *residual* inequality, we replace $g(w|x,T = t)$ in equation (6) with a "residual pricing" function, $\hat{g}(\varepsilon|x,T = t)$, which is obtained by regressing the logarithm of hourly wages in each year on the full set of covariates in x, then replacing the wage observations in $g(w|x, T = t)$ with their corresponding residuals from the OLS regression. This residual price function provides the conditional distribution of wage *residuals* in year t and can be used analogously to $g(w|x,T = t)$ for calculating counterfactual residual densities.

B. Results

Trends in observed and counterfactual overall and residual inequality are summarized in table 4, and plotted in figures 8 and 9. In these figures, differences in the vertical height of each series within a given year reflect the effect of prices on overall earnings inequality, holding labor force composition at the appointed year's level. The over-time change in the level of each series moving along the x axis reflects the effect of changes in labor force composition, holding prices at their 1973, 1989, or 2005 levels.

For comparison with Lemieux (2006b), we begin by discussing residual inequality. Panel A of table 4 shows that male 90/50 (upper-tail) residual wage inequality rose during both halves of the sample: by 4.4 log points from 1973 to 1989 and by 4.0 log points from 1989 to 2005. Holding labor force composition constant at its 1973, 1989, or 2005 levels does not change the basic message. In all cases, the composition-constant rise in residual 90/50 inequality is at least 65% as large as the unadjusted change. This finding is readily seen in the upper-left panel of figure 8, which plots actual and counterfactual male 90/50 residual inequality over 1973 to 2005. A comparison of the heights of the 1973, 1989, and 2005 series (that is, pairing the 1973, 1989, and 2005 prices with the observed labor force composition in each year) demonstrates that composition-constant male 90/50 residual wage inequality rose substantially, both between 1973 and 1989 and between 1989 and 2005. As is visible from the shallow upward slopes of each counterfactual series (moving along the x axis), compositional shifts also contributed to rising residual inequality, particularly after 1988. But these compositional shifts are modest relative to the price effects.

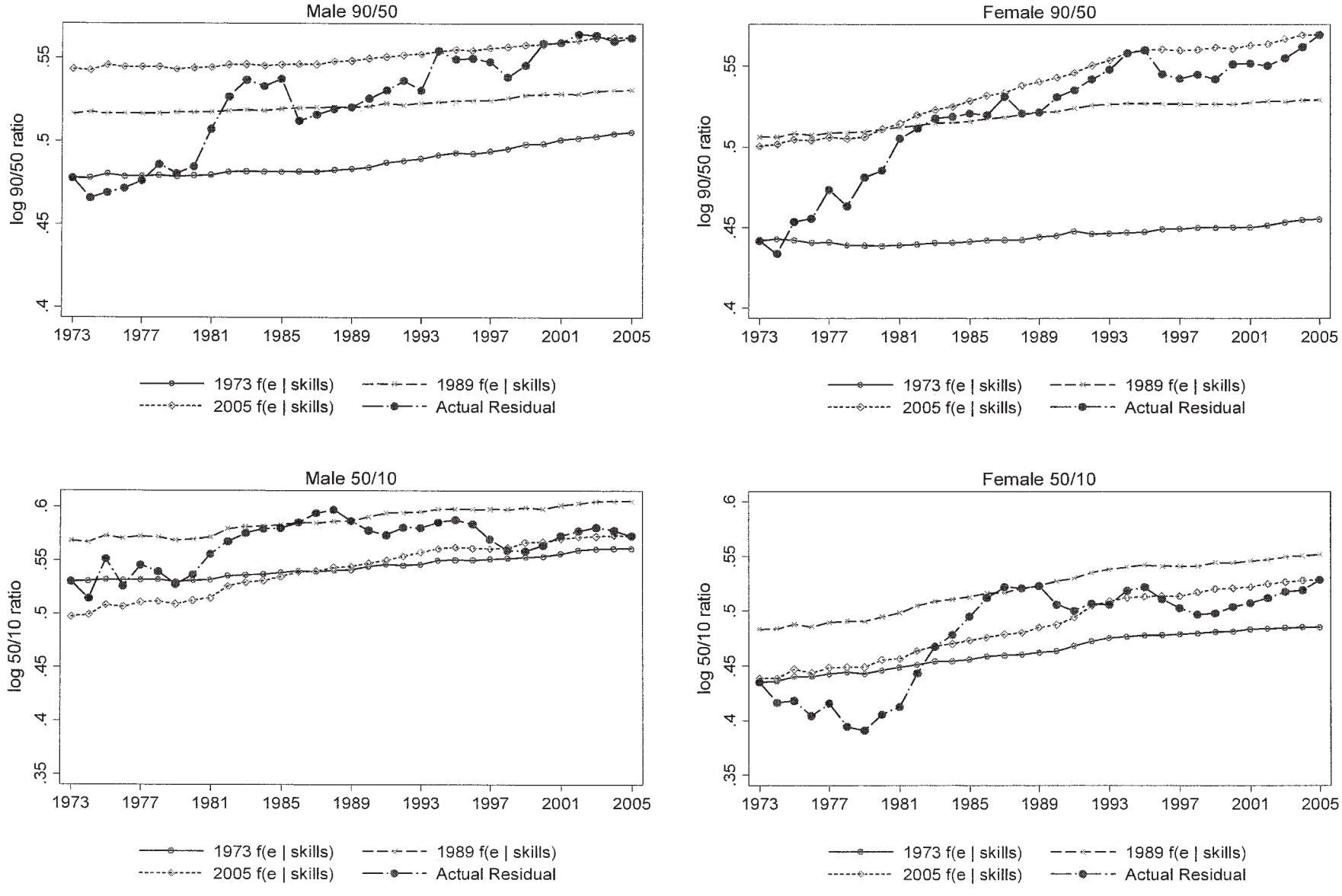
Next consider the evolution of *lower-tail* residual inequality during 1973 to 2005. Male 50/10 residual inequality rose by 5.7 log points between 1973 and 1989 and then fell by 1.3 log points between 1989 and 2005 (panel B of table 4). What are the roles of composition and prices in these shifts? Figure 8 shows that both the expansion and compression of lower-tail inequality are largely explained by *price* changes. In particular, during the first half of the sample, the composition-constant growth of residual 50/10 inequality was at least 65% as large as the unadjusted growth. In the latter half of the sample, price changes were also paramount: composition-constant lower-tail residual inequality *fell* by somewhere between 3.5 and 7.1 log points during 1989 to 2005, with the precise magnitude depending upon the choice of the base year. In short, the compression of residual prices during 1989 to 2005 was opposite in sign but comparable in magnitude to the expansion of residual prices during 1973 to 1989.

There is, however, an important difference between the earlier and latter halves of the sample. Consistent with Lemieux (2006b), figure 8 shows that labor force composition exerted a countervailing, upward effect on lower-tail residual inequality after 1989. In particular, the shallow upward slopes of each counterfactual inequality series starting in the mid- to late 1980s indicates that, holding prices constant, shifts in labor force composition served to increase measured residual inequality. Were it not for the countervailing effect of these compositional shifts, the observed decline in residual lower-tail inequality would have been even larger. The fact that we find a net decline in (non-composition-adjusted) lower-tail residual inequality between 1989 and 2005, however, indicates that the price effects reducing residual inequality dominated the composition effects raising it.

We have so far discussed residual inequality. Notably, our conclusions about the relative importance of price and

[25] In their analysis of the contributions of price and quantities to rising inequality during the 1980s, DiNardo, Fortin, and Lemieux (1996) perform supply and demand adjustments for observed wages by education by experience cells (in addition to applying a kernel reweighting procedure). DFL find these supply-demand adjustments to be important, explaining 21% to 33% of the growth in male 90/10 log hourly earnings inequality between 1979 and 1988. This indicates that the assumption that aggregate skill supplies do not affect skill prices is likely to be far from innocuous.

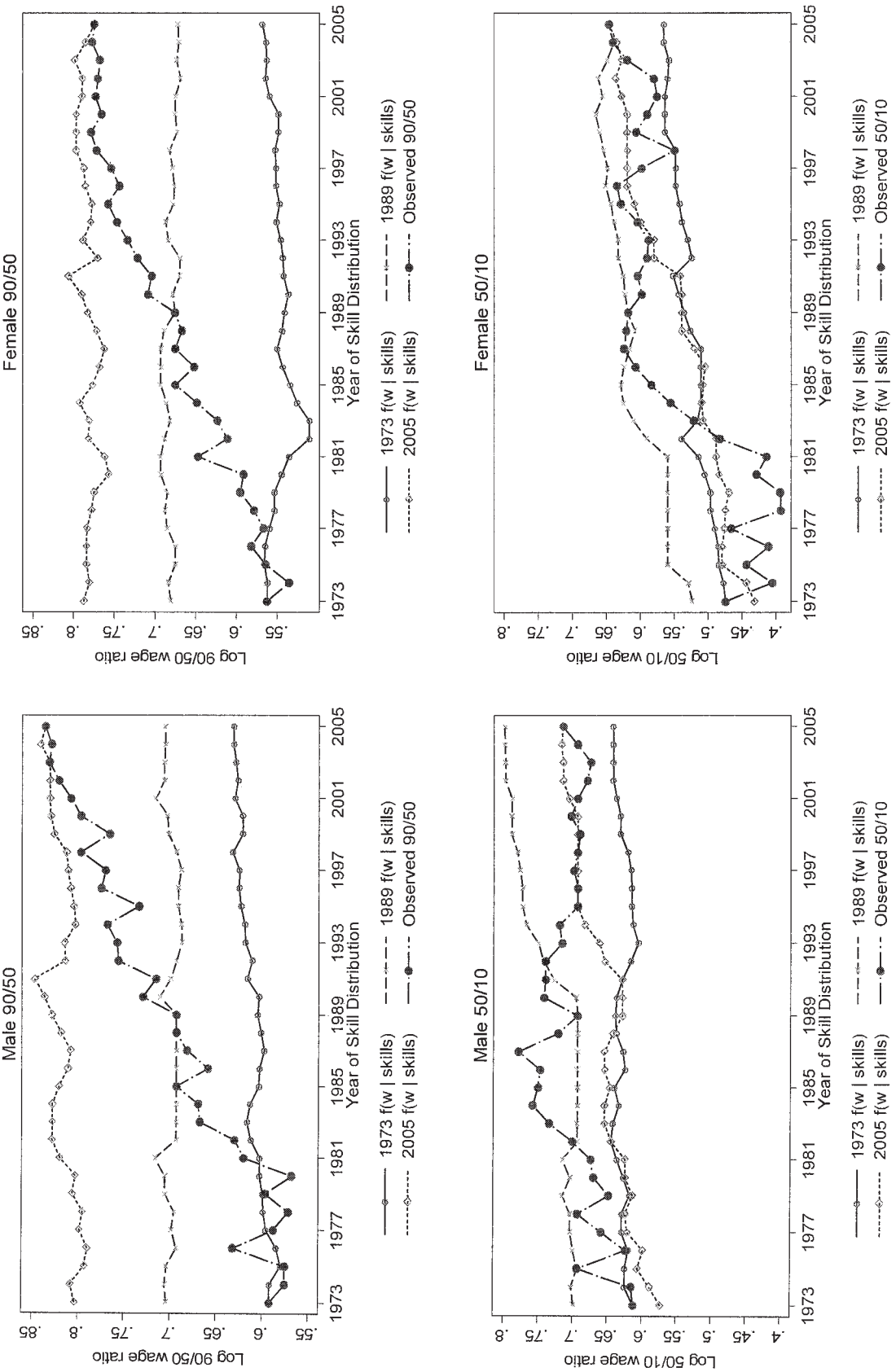
FIGURE 8.—ACTUAL AND COUNTERFACTUAL 90/50 AND 50/10 RESIDUAL HOURLY WAGE INEQUALITY, 1973–2005 (CPS MAY/ORG)



Data source: CPS May/ORG samples for all hourly workers are detailed in notes to figure 2. Series labeled "observed residual" presents the 90/50 or 50/10 difference in wage residuals from an OLS regression (weighted by CPS sampling weight times hours worked in the prior week) of log hourly earnings on a full set of age dummies, dummies for nine discrete schooling categories, and a full set of interactions among the schooling dummies and a quartic in age. All models are estimated separately by gender.

Series labeled "YEAR f(e | skills)" present 90/50 or 50/10 difference in wage residuals for a reweighted residual wage distribution from year "YEAR," where weights are proportional to ratio of the density of skills (defined by age, schooling, and their interactions) in the year depicted on the x axis to the corresponding density of skills in year "YEAR." Thus, the series labeled "1973 f(e | skills)" plots the 90/50 or 50/10 difference in wage residuals using the 1973 density of wage residuals reweighted using the skills distributions of years 1973 through 2005. See text for additional details of reweighting procedure.

FIGURE 9.—ACTUAL AND COUNTERFACTUAL 90/50 AND 50/10 OVERALL HOURLY WAGE INEQUALITY, 1973–2005 (CPS MAY/ORG)



Samples and data processing are identical to figure 8 except that observed hourly wages are used in place of hourly wage residuals.

TABLE 4.—100 × OBSERVED AND COMPOSITION-CONSTANT CHANGES IN OVERALL AND RESIDUAL HOURLY INEQUALITY MEASURES (MAY/ORG CPS 1973, 1989, AND 2005)

	Residual Inequality			Overall Inequality		
	1973–1989	1989–2005	1973–2005	1973–1989	1989–2005	1973–2005
A. Δ 90/50						
Males						
Observed	4.4	4.0	8.4	10.2	14.2	24.4
1973 X's	4.0	2.8	6.8	11.2	9.2	20.4
1989 X's	3.6	2.6	6.2	8.8	13.5	22.3
2005 X's	2.8	2.9	5.7	7.9	13.0	20.9
Females						
Observed	8.2	4.8	12.9	11.3	9.8	21.1
1973 X's	6.8	−0.9	5.9	11.8	9.9	21.7
1989 X's	7.6	1.9	9.5	13.2	10.5	23.7
2005 X's	8.0	4.0	12.0	10.7	10.1	20.8
B. Δ 50/10						
Males						
Observed	5.7	−1.3	4.4	8.1	2.1	10.2
1973 X's	3.9	−7.1	−3.2	8.7	−12.6	−3.9
1989 X's	6.3	−4.0	2.3	5.2	−6.4	−1.2
2005 X's	4.9	−3.5	1.4	15.7	−8.4	7.3
Females						
Observed	8.8	0.4	9.2	14.4	2.8	17.3
1973 X's	4.4	−4.5	−0.1	4.9	−8.4	−3.5
1989 X's	5.7	−3.9	1.8	8.5	−7.9	0.6
2005 X's	6.2	−2.5	3.7	8.3	−0.1	8.2
C. Δ 90/10						
Males						
Observed	10.1	2.8	12.8	18.3	16.4	34.6
1973 X's	7.9	−4.3	3.6	19.9	−3.4	16.5
1989 X's	8.0	−1.5	6.5	13.7	7.0	20.7
2005 X's	7.7	−0.7	7.0	23.4	4.6	28.0
Females						
Observed	16.9	5.2	22.1	25.7	12.7	38.4
1973 X's	11.2	−5.4	5.8	16.7	1.5	18.2
1989 X's	13.3	−2.0	11.3	21.7	2.6	24.3
2005 X's	14.2	1.4	15.6	18.9	9.9	28.8

Tabulated statistics present observed and composition-constant residual and overall wage inequality from CPS May/ORG hourly earnings series for 1973–2005. Residual inequality statistics are obtained from a regression of log hourly wages on a full set of age dummies, dummies for nine discrete schooling categories, and a full set of interactions among the schooling dummies and a quartic in age. All models are estimated separately by gender. Compositional adjustments are made using the DiNardo-Fortin-Lemieux (1996) kernel reweighting approach. See text for further details.

composition effects for the evolution of residual inequality apply with even greater force to overall inequality. As shown in figure 9, the dramatic rises in (overall) upper-tail inequality during both halves of the sample are almost completely accounted for by price changes; the contribution of compositional shifts is so small as to be almost undetectable. For lower-tail inequality, price and compositional changes both play significant, albeit countervailing, roles, with price changes causing lower-tail compression and compositional shifts leading to lower-tail expansion.

C. *Reconciling with Lemieux (2006b)*

How can our finding that composition plays only a secondary role in explaining the time patterns of residual inequality for males in the CPS May/ORG be reconciled with Lemieux's (2006b) conclusion that the mechanical effects of changes in labor force composition explain the full rise in residual inequality after 1988?[26] The answer is seen by studying the net rise in 90/10 residual inequality (thus summing over 90/50 and 50/10 residual inequality) between 1989 and 2005. Panel C of table 4 shows that the observed rise in 90/10 residual inequality of 2.8 log points over 1989 to 2005 is small relative to the substantial rise in upper-tail inequality and compression of lower-tail inequality. Consequently, the relatively modest contribution of compositional shifts to residual inequality over this period can be said to explain all—in fact, more than all—of the observed rise in 90/10 residual inequality. But this summary conclusion aggregates over two important countervailing forces. The first is the contraction in lower-tail inequality after 1988, which is due to compressing residual "prices" partially offset by changing composition. The second is the rise in upper-tail inequality, which appears to be almost

[26] A similar analysis performed using the March CPS samples also supports this conclusion (see online reference table 2).

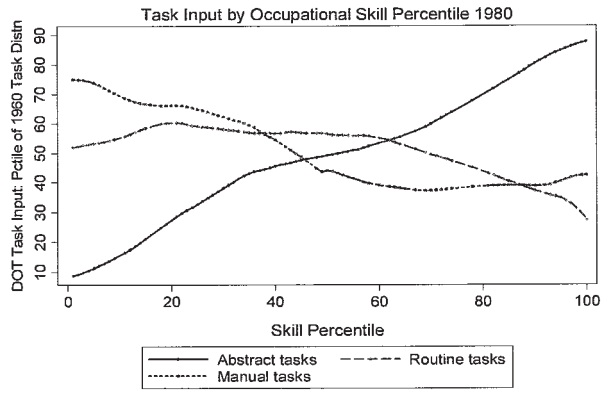
entirely explained by price changes. Because composition overexplains the former phenomenon and underexplains the latter, it is accurate to say that composition can "fully explain" the aggregate trend in residual inequality after 1988. But when upper- and lower-tail inequality are considered separately—as seems appropriate given their substantial divergence—composition does not appear a satisfying explanation for either.

V. What Explains the Polarization of Earnings Growth?

Following the monotonic surge of inequality from 1979 to 1987, changes in the U.S. wage distribution subsequently "polarized," with a strong, persistent rise in inequality in the upper half of the distribution and a slowing then slight reversal of inequality in the lower half of the distribution. This polarization is seen in overall inequality, in residual inequality, and in educational wage differentials. We believe the evidence is clear that this polarization is not primarily explained by mechanical effects of labor force composition or by episodic, nonmarket events such as the falling real minimum wage of the 1980s. But canonical supply-demand models also fail to provide a satisfying explanation for wage polarization since the late 1980s.

One potentially viable hypothesis focuses on changing demand for job tasks and their link to computerization. As argued by Autor, Levy, and Murnane (2003, ALM hereafter) and amplified by Goos and Manning (2007), Spitz-Oener (2006), Autor, Katz, and Kearney (2006), and Dustmann, Ludsteck, and Schönberg (2007), skill-biased technical change is probably an insufficiently nuanced name for the shifts in skill demands that were induced or abetted by the rapid price declines in computer technology over the last three decades. In the ALM task framework, computerization has nonmonotone impacts on the demand for skill throughout the earnings distribution: sharply raising demand for the cognitive and interpersonal skills used by educated professionals and managers ("abstract tasks"); reducing demand for clerical and routine analytical and mechanical skills that comprised many middle-educated white collar and manufacturing production jobs ("routine tasks").[27] Somewhat paradoxically, computerization has probably had little direct impact on the demand for the nonroutine manual skills ("manual tasks") used in many "low-skilled" service jobs such as health aides, security guards, orderlies, cleaners, and servers. Because the interpersonal and environmental adaptability demanded by these manual tasks has proven extraordinarily difficult to computerize (to date), these manual activities may in fact grow in importance as a share of labor input.[28] The ALM framework suggests that computerization (among other forces such as international outsourcing) may have raised demand for skill among higher-educated workers, depressed skill demands for "middle-educated" workers, and left the lower echelons of the wage distribution comparatively unscathed.[29] Goos and Manning (2007) label this process a "polarization of work," and argue that it may have contributed to a hollowing out of the wage distribution in the United Kingdom from 1975 to 2000. Spitz-Oener (2006) and Dustmann, Ludsteck, and Schönberg (2007) report a similar polarization of employment for the former West Germany for 1979 to 1999.[30]

FIGURE 10.—TASK INTENSITY BY OCCUPATIONAL SKILL PERCENTILE, DEFINED AS OCCUPATIONAL RANK (IN PERCENTILES) IN MEAN YEARS OF SCHOOLING



Task intensity by occupational percentile is plotted using a locally weighted smoothing regression with bandwidth 0.5 (i.e., one-half of one percentile). Occupational skill is measured as the employment-weighted percentile rank of an occupation's mean years of education in the Census IPUMS 1980 5% extract. Mean education in each occupation is calculated using workers' hours of annual labor supply times the Census sampling weight. Data on task intensity by occupation is from the *Dictionary of Occupational Titles* and compiled by Autor, Levy, and Murnane (2003). Task intensities are measured as percentiles of the baseline distribution of job tasks in 1960. Thus, an occupation with the median intensity of routine task input in 1960 would receive a score of 50.

To illustrate the relevance of shifts in task demands for changes in skill demands, we link data on task intensity by occupation from the *Dictionary of Occupational Titles* to data on skill level by occupation in the 1980 Census. In this analysis, occupational skill level is measured by the mean years of education of an occupation's workforce (weighting workers by their annual hours worked). Figure 10 uses a locally weighted smoothing regression to plot task intensity by occupational skill for each of the three broad task categories above: abstract, routine, and manual tasks.[31] Task

[27] A related earlier model along these lines is developed in Juhn (1994).

[28] See Levy and Murnane (2004) for numerous paradigmatic examples. The fact that computerization causes manual tasks to grow as a share of labor input may be understood as a form of Baumol's disease.

[29] Welch (2000) and Weinberg (2000) argue that these technical changes are particularly likely to have been favorable to demand for female labor.

[30] Acemoglu (1999) offers an alternative theory of job polarization based on endogenous changes in production techniques as a response to a rise in the availability of skilled labor.

[31] The task intensity data are constructed by matching Census 1980 data by occupation and gender with task measures from the *Dictionary of Occupational Titles* (*DOT*). Task intensities by occupational skill percentile are plotted using a locally weighted smoothing regression with bandwidth 0.5 (that is, one-half of one percentile). Details on the processing and matching of *DOT* task measures to occupations are given in Autor, Levy, and Murnane (2003). The abstract task category we use in figure 10 is the arithmetic average of ALM's "nonroutine cognitive/analytic" and "nonroutine cognitive/interactive" category and, similarly, our routine task category is the average of ALM's "routine manual" and "routine cognitive" categories. Our manual category is equivalent to ALM's "nonroutine manual" category.

intensities are measured as percentiles of the baseline distribution of job tasks in 1960. Thus, an occupation with the median intensity of "routine" task input in 1960 would receive a score of 50. This figure shows that the intensity of abstract skill input is monotonically rising in occupational skill (that is, education) and, conversely, the intensity of manual task input is falling in occupational skill. Most significantly, there is a distinctly nonmonotone relationship between occupational skill and routine task input. Routine task use is highest between the 20th and 60th percentiles of the skill distribution, and falls off sharply on either side of this range. This nonmonotonic relationship is highly relevant because, as documented by ALM, routine task input saw the sharpest decline of all task categories over the last two decades (relative to its initial 1960 level). The substitution of information technology for routine tasks might be expected to contribute to polarization by reducing demand for "middle-skill" occupations relative to either high- or low-skill occupations.

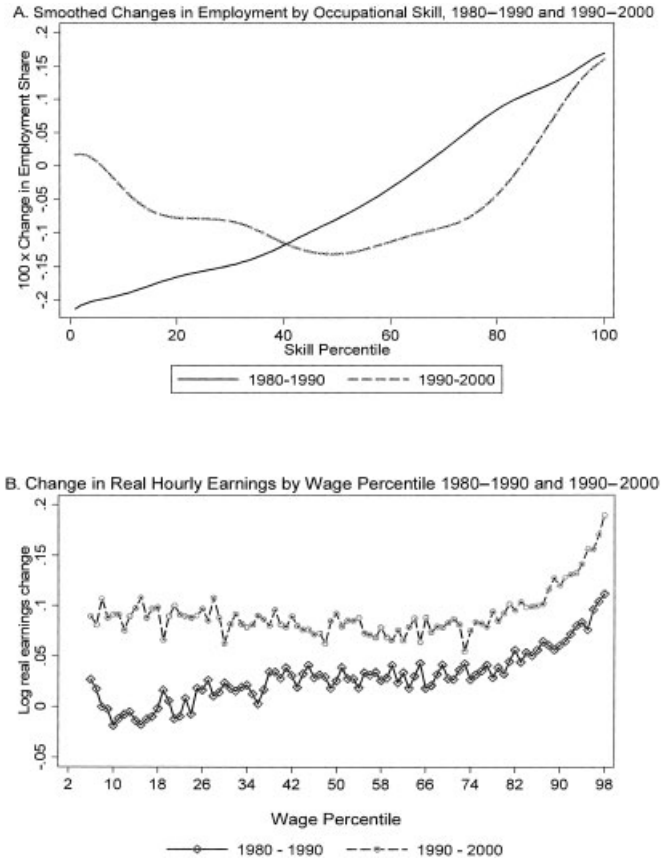
An implication of the polarization hypothesis is that the twisting of the wage structure observed in recent years is, in significant part, a demand-side phenomenon, induced by rising relative demands for both high- and low-skill tasks. This implication is testable, and we provide a simple evaluation here. Following Goos and Manning's (2007) analysis for the United Kingdom, we use U.S. Census data to explore how employment growth by occupation over the last two decades is related to occupational skill as proxied by educational levels.[32] Our hypothesis is that, if the wage structure changes observed in the 1980s and 1990s are driven in substantial part by demand shifts, wage changes by earnings level and employment changes by skill level should positively covary in both decades.

To test this implication, we plot in figure 11 the change in the share of total hours worked in the economy from 1980 to 1990 and 1990 to 2000 by occupation skill percentile, using the education-based occupational skill measure developed above.[33] For the decade of the 1980s, we see substantial declines in employment shares at the bottom end of the skill distribution with strongly monotone increases in employment shares as we move up the skill distribution. In contrast, employment growth in the 1990s appears to have polarized. There is rapid employment growth in the highest-skill jobs (at or above the 75th percentile), a decline in the employment shares of middle-skill jobs (those at percentiles 30 to 75), and flat or rising employment shares in the lowest-skill jobs, those in deciles one through three.

[32] Autor, Katz, and Kearney (2006) present a similar analysis using Census data for changes in occupational employment and CPS May/ORG data for changes in wage levels by earnings percentile. In the present analysis, we use exclusively Census data covering the same time periods.

[33] We employ a consistent set of occupation codes for Census years 1980, 1990, and 2000 developed by Meyer and Osborne (2005). We use a locally weighted smoothing regression (bandwidth 0.8 with 100 observations) to fit the relationship between decadal growth in occupational employment share and occupations' initial skill percentile in the 1980 skill distribution.

FIGURE 11.—(A) CHANGE IN OCCUPATION'S EMPLOYMENT SHARES, 1980–1990 AND 1990–2000, BY OCCUPATIONAL SKILL PERCENTILE IN 1980 (B) CHANGES IN REAL HOURLY EARNINGS BY WAGE PERCENTILE, 1980–1990 AND 1990–2000 (CENSUS 1980, 1990, AND 2000, ALL HOURLY WORKERS)



Data source: Census IPUMS 5% extracts for years 1980, 1990, and 2000. Sample includes respondents ages 18 through 64 who were currently employed in the civilian labor force at the time of the survey, who were not in unpaid family work, and who did not live in correctional institutions, mental institutions, or other noninstitutional group quarters. The labor supply measure is equal to the Census sample weight times the product of weeks worked and usual weekly hours in the prior year. For respondents with missing hours, we impute labor supply using the mean of workers of the same occupation and education group (high school dropout, high school graduate, some college, college graduate). Where the occupation-education cell is empty, we assign the mean for the education group.

Panel A plots log changes in employment shares by 1980 occupational skill percentile rank using a locally weighted smoothing regression (bandwidth 0.8 with 100 observations), where skill percentiles are as defined in the notes to figure 10. Consistent occupation codes for Census years 1980, 1990, and 2000 are from Meyer and Osborne (2005).

Panel B plots decadal changes in real log wages by wage percentile. Wage calculations exclude the self-employed and those with missing hours or weeks worked. Hourly wages are calculated as total wage and salary income divided by hours of labor supply. The bottom 1% of hourly earners are dropped. Hourly wages of top-coded earners are multiplied by 1.5. We limit the maximum hourly wage (via truncation) to the 1.5 times the maximum annual income amount divided by 1,750 (35 hours per week for 50 hours per year). All wage calculations are weighted by the product of Census person weights and calculated or imputed annual labor supply. Earnings are deflated using the PCE deflator.

This pattern of job growth corresponds closely with the observed pattern of wage structure changes in each decade, as is shown in the lower panel of figure 11. Real wage growth was essentially monotone in wage percentile in the 1980s, with especially sharp wage growth above the 75th percentile and especially sharp declines below the 30th percentile. In the decade of the 1990s, however, wage growth was more U-shaped. Wage growth was stronger below the 30th percentile and especially above the 80th percentile of the distribution than throughout the remainder of the distribution. Thus, despite substantial differences in the evolution of inequality between the 1980s and 1990s,

labor market prices and quantities (as measured by wage and skill percentiles) appear to positively covary in each decade.

To provide a slightly more rigorous assessment of this observation, we estimate a set of OLS models of the form

$$\Delta E_{p\tau} = \alpha_\tau + \beta_\tau \Delta \ln W_{p\tau} + \varepsilon_{p\tau}, \tag{8}$$

where changes in log employment share by skill percentile are regressed on changes in log wages by wage percentile in each decade. Here, $\Delta E_{p\tau}$ represents the change in occupational log employment share at skill percentile p in decade τ and $\Delta \ln W_{p\tau}$ is the change in real log hourly earnings at the corresponding wage percentile in the same decade.[34] Using data for the 4th through 97th percentiles of the earnings and skill distributions (thus trimming outliers at the tails), we estimate that $\beta_\tau = 3.00$ ($t = 3.75$) for the 1980s and that $\beta_\tau = 2.96$ ($t = 1.90$) for the 1990s. Thus, both the monotone rise of wage inequality in the 1980s and the "polarized" growth of wage inequality in the 1990s are mirrored by conformable changes in employment by skill. This is consistent with a demand-side explanation for observed wage changes.[35]

We have further experimented with these simple models by including linear terms in wage percentiles in addition to (or instead of) estimated wage changes by percentile. For the decade of the 1980s, we find that a linear function of wage percentiles fits the observed pattern of skilled employment growth better than does the observed change in earnings by percentile. In the 1990s, by contrast, the linear term is insignificant and the estimate of $\beta_{90\text{–}00}$ is hardly affected by its inclusion (either in magnitude or precision). These simple models do not, of course, take into account the substitutability and complementary among various skill groups (as measured by skill percentiles) and so lack a well-grounded production function interpretation. We nevertheless view them as suggestive evidence that labor demand shifts have favored low- and high-wage workers relative to middle-wage workers over the last fifteen years—a pattern that stands in contrast to the shifts in labor demand during the 1980s, which appear to have been monotonically rising in skill.

VI. Conclusion

The recent divergence of upper-tail and lower-tail wage inequality and the convexification in the returns to education present puzzles for the traditional and revisionist interpretations of changes in the U.S. wage structure. Revisionist explanations focusing on nonmarket factors, particularly minimum wages and the spurious effects of labor market composition on residual inequality, appear unable to provide a compelling explanation for the strong, secular rise in upper-tail inequality over at least the last 25 years and the polarization of employment growth since 1990. Explanations based on relatively steady-trend relative demand growth for more-skilled workers and fluctuations in relative skill supplies based on canonical supply-demand models, such as the two-factor CES used by Katz and Murphy (1992) and Autor, Katz, and Krueger (1998), imply a puzzling slowdown in relative demand growth starting in the early 1990s.

Despite our rejection of "unicausal" explanations for the complex pattern of wage changes we document, the simple analysis above demonstrating that employment and wage growth by skill percentile are positively correlated in each of the last two decades leaves us confident that skill demand shifts have played a central role in reshaping the wage structure, both during the monotone rise of inequality during the 1980s and the polarization of wage growth that followed. These trends toward wage structure "polarization" can, we believe, be partially reconciled by a reinterpretation of the skill-biased technical change hypothesis along the lines developed by Autor, Levy, and Murnane (2003), Autor, Katz, and Kearney (2006), and Goos and Manning (2007).

Nevertheless, there is more to be understood about the interactions among supply and demand, labor market institutions, and macroeconomic conditions than is accommodated by this simple "polarization" model. Equally important, the factors contributing to the evolution of U.S. inequality in the decade(s) to come may differ significantly from those identified above. International trade and outsourcing factors appear likely to become increasingly important, because of both rapid economic development in Asia and improvements in computer and communications technology that have dramatically reduced the costs of large-scale international trade in goods and services. Devising innovative and rigorous means to evaluate the impacts of more realistic form of technical change and these evolving globalization forces on inequality and economic well-being constitutes a significant agenda item for further research in this field.

[34] In contrast to figure 11, we use raw changes in employment shares by occupational wage percentile as the dependent variable rather than smoothed changes. If we were to instead use smoothed changes, this would not affect the point estimates by much but would substantially increase the precision of the estimates.

[35] Notably, this pattern appears inconsistent with the hypothesis that a declining minimum wage played a leading role in the expansion of lower-tail inequality in the 1980s. A decline in a binding wage floor should have led to a (modest) rise in low-wage employment rather than a sharp contraction.

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DATA APPENDIX

1. Basic Processing of May/ORG CPS Data

We use the May CPS for 1973 to 1978 and the CPS Merged Outgoing Rotation Groups for years 1979 to 2005. All samples include wage/salary workers ages 16 to 64 with 0 to 39 years of potential experience in current employment. Earnings weights, equal to the product of CPS sampling weights and hours worked in the prior week, are used in all calculations. Hourly wages are the logarithm of reported hourly earnings for those paid by the hour and the logarithm of usual weekly earnings divided by hours worked last week (not usual weekly hours) for nonhourly workers. We use hours last week instead of usual weekly hours because usual weekly hours are not consistently available: starting with the CPS redesign in 1994, workers who report that their weekly hours vary are not asked to report usual weekly hours, yielding a nonreport rate of 7% to 9% of workers in 1994 to 2005. To check sensitivity, we have tabulated and plotted overall and residual inequality measures using imputed usual weekly hours in place of hours last week in all years. This has little impact on our results.

Top-coded earnings observations are multiplied by 1.5. Hourly earners of below $1.675/hour in 1982 dollars ($2.80/hour in 2000 dollars) are dropped, as are hourly wages exceeding 1/35th the top-coded value of weekly earnings. All earnings numbers are deflated by the chain-weighted (implicit) price deflator for personal consumption expenditures (PCE). Allocated earnings observations are excluded in all years, except where allocation flags are unavailable (January 1994 to August 1995). As discussed by Hirsch and Schumacher (2004), only about 25% of allocated observations in the MORG CPS are actually flagged as allocated between 1989 and 1993. Following Lemieux (2006b), we identify and drop nonflagged allocated observations by using the unedited earnings values provided in the source data.

2. Basic Processing of March CPS Data

We use the March Current Population Survey for sample years 1964 to 2006 (covering earnings years 1963 to 2005) for workers age 16 to 64 (during the earnings year) with 0 to 39 years of potential experience whose class of work in their longest job was private or government wage/salary employment. Most analyses are limited to full-time, full-year workers, which we define as those who work 35 hours per week (using the Census Bureau's full-time worker flag) and forty-plus weeks in the prior year. Full-time weekly earnings are calculated as the logarithm of annual earnings over weeks worked for the full-time, full-year sample. Allocated earnings observations are excluded after (sample year) 1966 using family earnings allocation flags (1964 to 1975) or individual earnings allocation flags (1976 forward). Weights are used in all calculations. Full-time earnings are weighted by the product of the CPS sampling weight and weeks worked. Online reference tables 1a, 1b, and 2 additionally use March CPS hourly earnings for earnings years 1975 forward. Hourly earnings are calculated as annual earnings divided by the product of weeks worked and usual hours in the prior year and are weighted by the product of the CPS sampling weight, weeks worked, and hours worked in the prior year (available starting in earnings year 1975).

Prior to the March 1988 survey, all wage and salary income in the March CPS was reported in a single variable, which was top-coded at values between $50,000 and $99,999 in years 1964 to 1987. For these cases, we multiply the top-coded earnings value by 1.5, following Katz and Murphy (1992). Commencing in 1989, wage and salary incomes were collected in two separate earnings variables, corresponding to primary and secondary labor earnings. After adjusting for top-coding, we sum these values to calculate total wage and salary earnings. Starting in 1988, top-codes are handled as follows. For the primary earnings variable, top-coded values are reported at the top-code maximum up to 1995. We multiply these values by 1.5. Starting in 1996, top-coded primary earnings values are assigned the mean of all top-coded earners. In these cases, we simply reassign the top-coded value and, again, multiply by 1.5. For the secondary earnings value, the top-coded maximum is set at 99,999 from 1988 to 1995, falls to 25,000 for 1996 through 2002, and rises to 35,000 in 2003 through 2006. For lack of a superior alternative, we again use the top-coded value multiplied by 1.5. Earnings numbers are deflated using the PCE deflator.

3. Processing of Census Samples

We use the Census IPUMS 5% extracts for years 1980, 1990, and 2000 available at http://www.ipums.org. Our sample includes respondents ages 18 through 64 who were currently employed in the civilian labor force at the time of the survey, were not in unpaid family work, and who did not live in correctional institutions, mental institutions, or other noninstitutional group quarters. Our labor supply measure is the product of weeks worked and usual hours worked in the prior year. For respondents with missing hours, we impute labor supply using the mean of workers of the same occupation and education group (high school dropout, high school graduate, some college, college graduate). Where the occupation-education cell is empty, we assign the mean for the education group. We employ a consistent set of occupation codes for Census years 1980, 1990, and 2000 developed by Meyer and Osborne (2005). For wage calculations using Census data, we further exclude the self-employed and those with missing hours or weeks worked. Hourly wages are calculated as total wage and salary income divided by hours of labor supply. We drop the bottom 1% of hourly earners and multiply hourly wages of top-coded earners by 1.5. We limit the maximum hourly wage (via truncation) to 1.5 times the maximum annual income amount divided by 1,750 (35 hours per week for 50 hours per year). This correction prevents part-time workers from having a higher feasible hourly wage than full-time workers. All calculations are weighted by the product of Census person weights and calculated or imputed annual labor supply. Earnings numbers are deflated using the PCE deflator.

4. Coding of Education and Potential Experience in CPS and Census Samples

To attain comparable educational categories across the redefinition of the Census Bureau's education variable introduced in 1992 in the CPS and the 1990 Census, we use the method proposed by Jaeger (1997). In samples coded with the older education question, we defined high school dropouts as those with fewer than twelve years of completed schooling; high school graduates as those having twelve years of completed schooling; some college attendees as those with any schooling beyond twelve years (completed or not) and less than sixteen completed years; and college-plus graduates as those with sixteen or more years of completed schooling. In samples coded with the revised education question, we define high school dropouts as those with fewer than twelve years of completed schooling; high school graduates as those with either twelve completed years of schooling and/or a high school diploma or G.E.D.; some college as those attending some college or holding an associate's degree; and college plus as those with a B.A. or higher.

To calculate potential experience in data years coded with the revised education question, we use figures from Park (1994) to assign years of completed education to each worker based upon race, gender, and highest degree held. Years of potential experience were calculated as age minus assigned years of education minus 6, rounded down to the nearest integer value.

5. Construction of Relative Wage Series

We calculate composition-adjusted college/high school relative wages overall and by age or experience using the March and May/ORG samples described above. These data are sorted into sex-education-experience groups based on a breakdown of the data into two sexes, five education categories (high school dropout, high school graduate, some college, college plus, and greater than college), and four potential experience categories (0–9, 10–19, 20–29, and 30+ years). Log weekly wages of full-time, full-year workers (March CPS) and all hourly workers (May/ORG) are regressed in each year separately by sex on the dummy variables for four education categories, a quartic in experience, three region dummies, black and other race dummies, and interactions of the experience quartic with three broad education categories (high school graduate, some college, and college plus). The (composition-adjusted) mean log wage for each of the forty groups in a given year is the predicted log wage from these regressions evaluated for whites, living in the mean geographic region, at the relevant experience level (5, 15, 25, or 35 years depending on the experience group). Mean log wages for broader groups in each year represent weighted averages of the relevant (composition-adjusted) cell means using a fixed set of weights, equal to the mean share

of total hours worked by each group over 1963 to 2005 from the March CPS.

6. Construction of Relative Supply Measures

We calculate college/high school relative supply measures using the March and May/ORG samples above. We form a labor "quantity sample" equal to total hours worked by all employed workers (including those in self-employment) with 0 to 39 years of potential experience in 400 gender × education × potential experience cells: experience groups are single-year categories of 0 to 39 years; education groups are high school dropout, high school graduate, some college, college graduate, and post-college. The quantity data are merged to a corresponding "price sample" containing real mean full-time weekly (March CPS) or real mean hourly (May/ORG CPS) wages by year, gender, potential experience, and education. (Wage data used for the price sample correspond to the earnings samples described above.) Wages in each of the 400 earnings cells in each year are normalized to a relative wage measure by dividing each by the wage of high school graduate males with ten years of potential experience in the contemporaneous year. (The choice of the base earnings group is innocuous.) We compute an "efficiency unit" measure for each gender-experience-education cell as the arithmetic mean of the relative wage measure in that cell over 1963 through 2005.

The quantity and price samples are combined to calculate relative log college/high school supplies. Define the efficiency units of labor supply of a gender × education × potential experience group in year t as the efficiency unit wage measure multiplied by the group's quantity of labor supply in year t. Following Autor, Katz, and Krueger (1998) and Card and Lemieux (2001), we calculate aggregate college-equivalent labor supply as the total efficiency units of labor supplied by college or college-plus workers plus half of the efficiency units of labor supplied by workers with some college. Similarly, aggregate high school–equivalent labor supply is the sum of efficiency units supplied by high school or lower workers, plus half of the efficiency units supplied by workers with some college. Our college/high school log relative supply index is the natural logarithm of the ratio of college-equivalent to noncollege-equivalent labor supply (in efficiency units) in each year. This measure is calculated overall for each year and by ten-year potential experience groupings.